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Price, Inventories, and Volatility in the Global Wheat Market

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ABSTRACT

The study estimates a conditional mean model for international wheat prices and inventories. Endogenous price volatility and exogenous shocks in the price and inventory series are controlled for in estimation. Redressing the empirical linkage between volatility, prices, and inventories is important because volatility increases returns to inventories, which in turn may imply prices. The problem is also important from the regulator perspective, because publicly funded inventory programs have been traditional measures in stabilizing prices and improving food security by providing a buffer against adverse yield shocks and stock-outs. The structural model underlying the estimating equations is based on a dynamic inventory optimization problem.

The data suggest that the price of both wheat and wheat inventories is nonstationary and that they are significantly linked to each other in the short run but do not exhibit a stationary long-run equilibrium relationship. Price volatility is an important determinant in the short-run conditional mean processes for both the price and inventories. The pairwise causal relationships have only one direction each. Inventories imply price volatility, price volatility implies price, and price implies inventories, but not vice versa.

The parameter estimates suggest that when inventories decrease, price volatility increases. Thus, low inventories have likely been among the necessary conditions, but have not been a sufficient condition by themselves, for the price surge observed in 2008. The price and inventory movements have a significant negative relationship in the very short run, but it is leveled off over time. A decreasing price implies either inventory build-ups or postponement of inventory withdrawals. Overall, the current and past inventory and price movements are not very valuable in predicting the future price movements, and it is likely that the inventory information announced each month is already in the prices.

Keywords: international price, inventories, volatility, wheat

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1. INTRODUCTION

The surge of world market grain prices in spring 2008 resulted in an extensive debate about the foundations of international agricultural commodity markets and factors that truly triggered such a dramatic surge (for example, Mitchel 2008). The recent mathematical equilibrium analysis of Dewbre et al. (2008) and, in particular, the reports of Abbott, Hurt, and Tyner (2008) and Heady and Shenggen (2008) provide a detailed description, a broad discussion, and a synthesis about the issues and market developments in the context of the 2007–2008 food crisis.

Even though different stakeholders and research groups weight the factors underlying the dramatic price spike differently, there is a wide consensus that the low buffers against adverse demand and supply shocks, such as the end-of-season inventories, have likely been among the most critical and decisive factors that triggered the crisis. There are also serious concerns that the record-low inventories in 2008 encouraged speculation that further exacerbated the price movements, since the shrinking inventories and price increases may have raised arbitrage opportunities and expectations in favor of further continuum of increasing prices. The empirical study of Cooke and Robles (2009), using a broad set of different economic, financial market, and monetary indicators, supports the view that trading activities in the futures markets increased in particular among noncommercial traders and that speculation played a role in the price surge. The situation became even worse when countries threatened by severe food shortages imposed new export restrictions and export bans to protect their own inventories and consumers (Braun and Torero 2008).

Now that grain prices have decreased and are closer to the precrisis levels, there is additional support for the view that the market and policymakers were overreacting and that, as such, a large price surge may not have been fully justified by the market fundamentals alone (Braun and Torero 2009). The crisis has also triggered policy actions that might result in irreversible long-term economic developments. Capital- and population-rich countries are considering new means of rebuilding their stocks, increasing their self-sufficiency, and securing their consumers' access to food (Braun and Torero 2009). Thus, the drastic price surge may also have persistent implications for the international grain market, and a better understanding of these implications requires new knowledge about the relationship between inventories, prices, and price volatility.

Even if low levels of food inventories are among the core factors underlying the food crisis, we still lack recent rigorous empirical analysis on the causal relationship between the global grain inventories, international grain prices, and the volatility of these prices. This study complements the existing literature and quantitative equilibrium analyses with a statistical approach. We estimate a conditional mean and conditional volatility model for the global wheat inventories and the international wheat price using time-series econometrics. Our goal is to examine how informative the global inventories and price data are and to identify the empirical relationship between price, inventories, and volatility.

The economic foundations of our analysis are in the structural form inventory models of Pindyck (1994, 2004). These models define equilibrium inventory behavior as the solution to a stochastic dynamic optimization problem. They first solve for the intertemporal optimality conditions (Euler equations) or policy rules by dynamic programming and then estimate the model parameters in these optimality and orthogonality conditions using econometric techniques. Concerning the conditional means (the first moments), these studies suggest that a highly nonlinear negative relationship exists between the price and inventories and further, as an asymptotic condition, that the price should approach infinity if the end-of-season inventories approach zero. Compared with earlier inventory models, including those of Deaton and Laroque (1995, 1996) and Miranda and Rui (1997), our approach is more straightforward. We estimate the model in the reduced form, which is feasible because the United States Department of Agriculture (USDA) publishes data on the predicted end-of-season inventories and we are not circumvented by an absence of such critical data in the estimation.

In addition to the conditional mean processes, volatility plays a decisive role as a determinant of spot prices and optimal inventory rules through two effects (Pindyck 2004). First, volatility directly affects the marginal returns to storage, usually referred to as the *convenience yield*. The higher the variance and volatility of spot prices, the higher the marginal returns to storage. Thus, high volatility may result in inventory build-ups, or it may implicitly impose high opportunity costs for storage withdrawals and increase prices in the short run. Second, volatility may affect the marginal production costs and decrease production volumes by creating option premiums.

Based on the propositions and results of Deaton and Laroque (1995, 1996) and Pindyck (1994, 2004), we test the following questions concerning the volatility, price, and inventories:

- Does a long-term equilibrium between the price and the end-of-season inventories exist? Is there a negative price–inventories relationship between the price and inventories?
- Do exogenous price shocks change the relationship between the price and inventories?
- Do inventories and the price cause volatility, or is the volatility exogenous so that it affects inventories and the price?

The rest of the paper is structured as follows. The subsequent section describes the data. Thereafter, the econometric models and the estimation results are presented. The last section gives concluding remarks.

2. DATA

The price data are weekly quotations for nominal spot price (USD/metric ton) for hard wheat at U.S. Free on board (Fob) Gulf as given at Food and Agriculture Organization of the United Nations (FAO) statistics.[1] The data span January 1998 to July 2009. The monthly price series is constructed from the weekly series by taking the price quotation right after the monthly inventory update has been announced by the USDA. This price quotation has been, with only few exceptions, the price on the second Tuesday of each month. The resulting monthly price series has 139 observations, whereas the underlying weekly series has 572 observations (Figures 1 and 2).

The inventory data and consumption data, used to compute the inventory-to-consumption ratios (also referred to as the stock-to-use ratio), are obtained at World Markets and Trade Archives of Foreign Agricultural Service (FAS) in the USDA (Figures 1–3).[2] The inventory data are the predicted end-of-season global wheat inventories as they are announced in the monthly USDA-FAS reports. Thus, the inventories measure the predicted amount of grain reserves carried from the ongoing harvest year to the new harvest year. The definition of the harvest year is based on local harvest years. Therefore, the local harvest years overlap and depend on the location of the inventories. The largest trader of wheat in the international market is the United States, where the marketing season starts at the beginning of June and ends at the end of May. The consumption data are the predicted season's consumption levels. The descriptive statistics of the data are presented in Table 1.

Table 1. Descriptive statistics of predicted season's consumption levels

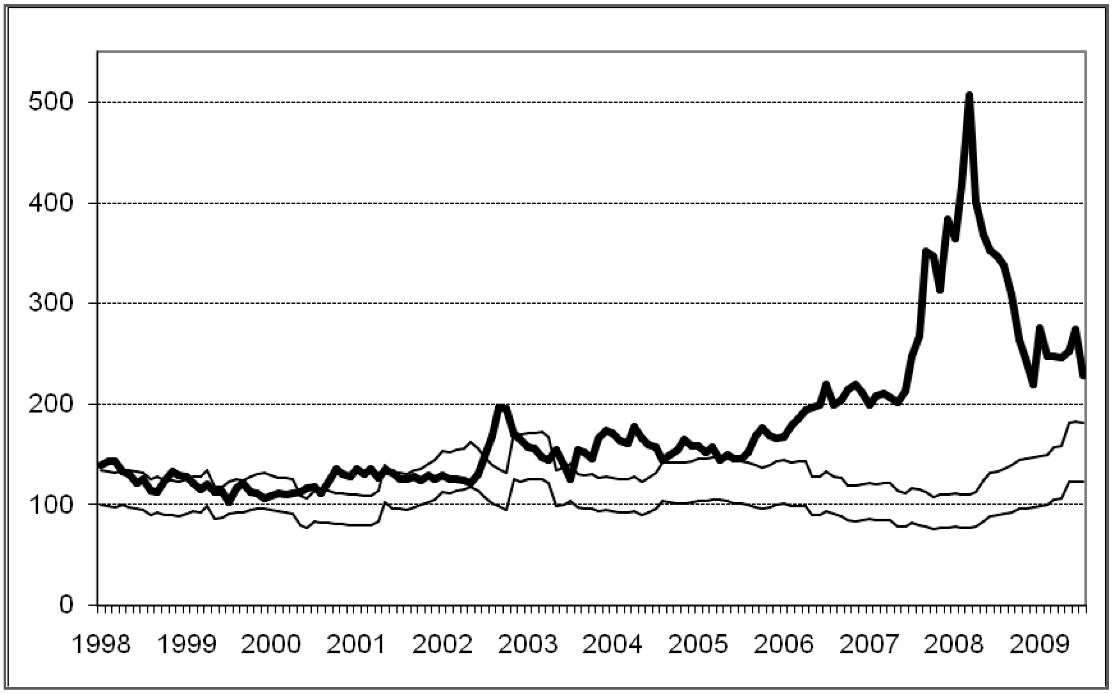
Description	Number of observations	Mean	Maximum	Minimum	Std. Dev.
Monthly wheat price (USD/metric ton)	139	178.2	507.0	103.0	75.1
End-of-season inventories (1,000 metric tons)	139	133.4	182.7	106.3	16.7
Predicted season's consumption (1,000 metric tons)	139	606.7	656.5	582.2	18.8
Inventories-to-consumption ratio	139	0.220	0.2903	0.1736	0.0270

Source: FAO: http://www.fao.org/es/esc/prices and USDA-FAS: http://www.fas.usda.gov/grain_arc.asp.

Seemingly, the patterns in the observed data support the view that there is a negative correlation between the inventories (relative to consumption) and price movements, so that low levels of inventories-to-consumption ratios are associated with periods of high prices (Figures 1 and 2). The global end-of-season wheat inventories decreased notably between 2006 and 2008, a result of two subsequent poor harvest years. The inventory-to-consumption ratio reached historically low levels in 2008. At this point, the price for wheat peaked to a record high, and it more than doubled within a few months. Nevertheless, it is not evident from the observed data that the low inventories alone would be a sufficient condition for the price surge. For example, the low inventories in 2000 did not result in a price surge and market reactions similar to 2008.

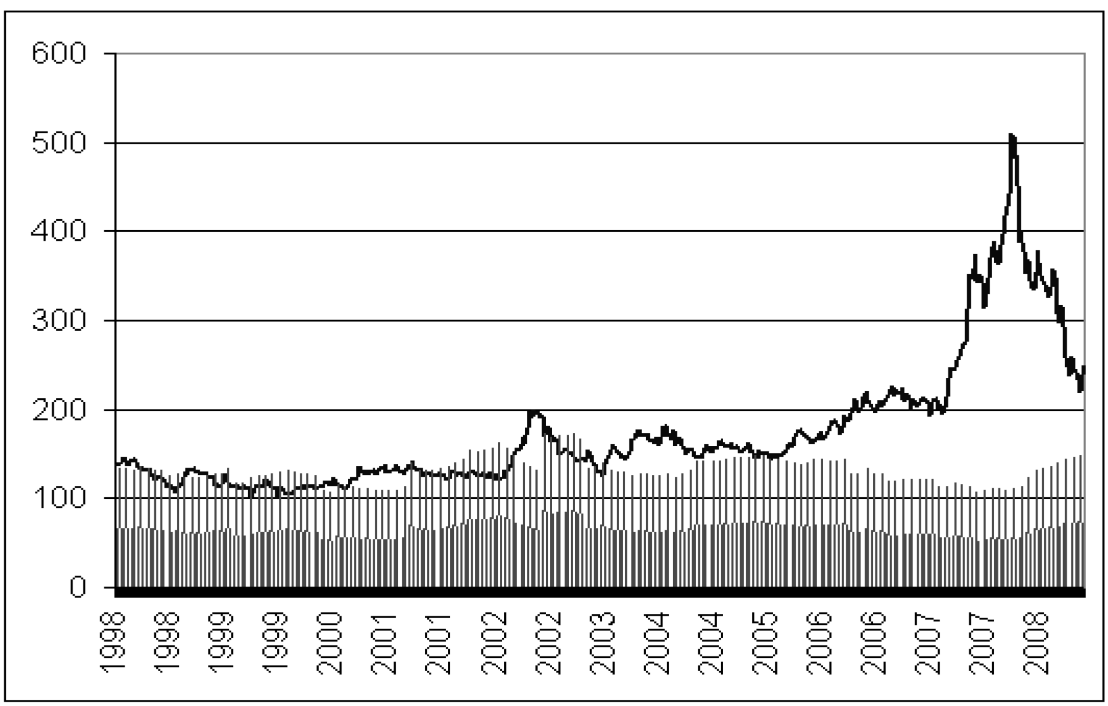
[1] U.S. No.2, Hard Red Winter ord. Prot, US Fob Gulf (Tuesday). The web page is http://www.fao.org/es/esc/prices.
[2] The web page is http://www.fas.usda.gov/grain_arc.asp.

Figure 1. The monthly price for wheat (USD/metric ton; thick line), end-of-season inventories (million metric tons; upper thin line), and inventories-to-consumption ratio (Jan 1998=100; lower thin line), January 1998 to July 2009



Source: FAO.

Figure 2. The weekly price notation for wheat (USD/metric ton; continuous line) and monthly end-of-season inventories (million metric tons; uppermost frontier shaped by the peaked lines)



Source: FAO.

3. ECONOMIC INVENTORY MODEL

Our model is a simplified version of Pindyck (2004). We consider price-taking and risk-neutral firms that maximize the expected present value (V_t) of the flow of their one-period profits. Firms choose production (z_t), inventory levels (s_t), and sales (q_t) to maximize:

$$V = \max_{\{z_t, s_t, q_t\}_{t=0}^{\infty}} E_0 \sum_{t=0}^{\infty} \beta^t (p_t q_t - C_t) \tag{1}$$

subject to:

$$s_t = s_{t-1} + z_t - q_t \text{ and } s_t \geq 0 \; for \; all \; t,$$

where E is the expectations operator, β is the discount factor, p is the output price, q is the sales or marketed quantity, C is the one-period cost function, z is the quantity of production, and t is the time index. Maximization is constrained by the transition equation describing the evolvement of inventories and by nonnegative inventories (that is, selling short is not allowed). The constraint of "no short selling" is the standard in estimating inventory models and implies that the commodity has to be produced before it can be sold in the spot market.[3] The one-period cost function (C_t) is the sum of the production cost (C^p), the marketing cost (C^m), and the constant per-unit storage cost (k) times the inventories:

$$C_t = C^p(z_t) + C^m(s_t, p_t, \sigma_t) + ks_t, \tag{2}$$

where σ_t refers to the price volatility. As compared with the depletable resource model of Pindyck (2004), we have dropped the option valuation component from the total costs because the current wheat production does not increase the next period's production costs and does not create option values comparable to those of depleting nonrenewable natural resources.

The marketing cost (Cm) is an aggregate of the costly activities facilitated by inventories, such as delivery scheduling, stock-out avoidance, and the cost of adjusting production over time. The marketing cost approaches infinity as the inventories approach zero (meaning $\lim_{s \to 0} C^m = \infty$). This asymptotic condition also guarantees that the constraint of nonnegative inventories in equation (1) will never bind. The marketing cost (Cm) is a decreasing function on the inventory level (s_t) and an increasing function on price volatility (σ_t). For simplicity, Cm does not depend on wheat price (p_t), as in Pindyck (2004). Hence a marginal unit of inventory reduces the total marketing cost and therefore delivers a convenience yield (ψ), such that:

$$\psi_t(s_t, \sigma_t) = -\partial C_t^m(s_t, \sigma_t) / \partial s_t, \tag{3}$$

which is decreasing in storage and increasing in volatility: $\partial \psi_t / \partial s_t < 0$ and $\partial \psi_t / \partial \sigma_t > 0$. Thus, when storage is decreasing and approaches zero, the value of a marginal unit of storage (the convenience yield) increases and approaches infinity. Also, when price volatility increases, the convenience yield increases.

Replacing the law of motion of inventories into equation (1) and then maximizing the result with respect to production (z_t) and inventories (s_t), we get the following first-order conditions:

$$p_t = \frac{\partial \, C^p(z_t)}{\partial z_t} \tag{4}$$

and

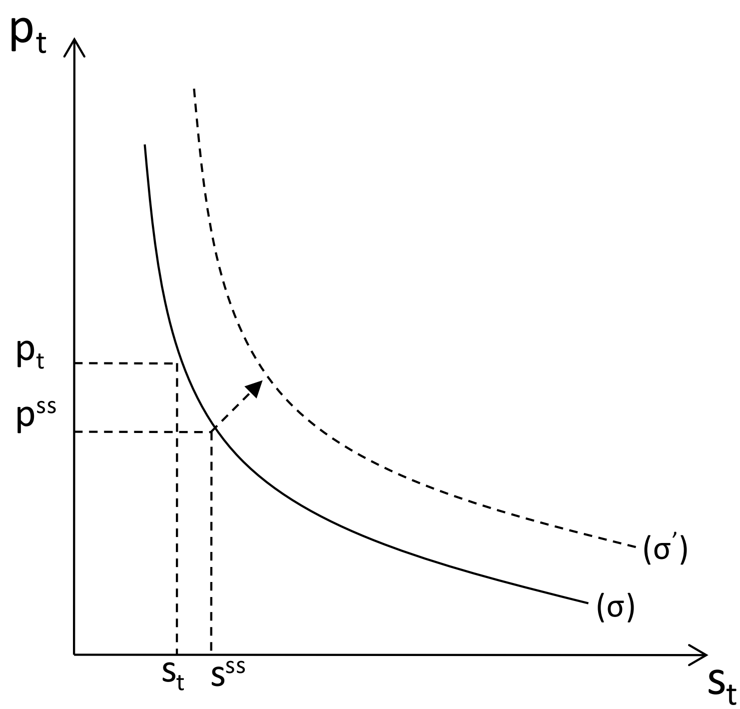
[3] The nonnegativity constraint could be relaxed by introducing derivatives and allowing for speculation.

$$\beta E_t[p_{t+1}] = p_t + \partial C_t^m / \partial s_t + k$$
$$\beta E_t[p_{t+1}] = p_t - \psi_t(s_t, \sigma_t) + k \quad (5)$$
,

which assert that, in equilibrium, the expected next-period price equals the current price plus the marginal cost of marketing and storage. This expression depicts an implicit function relating inventories to spot prices, expected prices, price volatility, and fixed inventory cost.

From equations (4) and (5), one can derive optimal choices and equilibrium relationships. Notice that in this simple partial equilibrium model, firms are price takers so that the price and price volatility are both exogenous objects. One can also assume without loss of generality that price expectations are exogenous and that they are independent of the current price level p_t. So a transitory shock affecting the current price does not affect price expectations for next period. From equation (4), one gets the optimal production decision z_t, given a price p_t; this traces a standard increasing supply (or production-level) curve $z(p_t)$. From equation (5), the optimal level of inventories can be pinned down given expectations, current price, and price volatility. The current price p_t and inventory level s_t have a negative relationship, which depicts a demand curve for inventories (Figure 3). Also, despite that this is a partial equilibrium model, it is informative to evaluate equations (4) and (5) at a steady-state price p^{ss}, such that $E(p^{ss}) = p^{ss}$. From equation (4), one gets a steady-state production level (z^{ss}). From equation (5), we get a steady-state inventory level s^{ss} and we verify that steady-state sales q^{ss} equal the steady-state production level (z^{ss}).

Figure 3. Relationship between price, inventories, and volatility



Source: Authors' creation.

From this simple model we get the following predictions:

1. There is a short-run equilibrium with an inverse relationship between current prices and inventory level. Suppose the price level is at its long-run steady-state level p^{ss} and then suddenly jumps up to p_t, due for example to a transitory demand or supply shock. As future expectations remain constant, inventory levels decrease from s^{ss} to s_t (see Figure 3). In other words, firms react to the transitory shock by selling out current production and some previous period stocks.
2. An increase (decrease) in price volatility shifts the short-run price inventory relationship to the right (left). One can verify that for higher price volatility, given a steady-state price, a larger inventory level is needed to satisfy equation (5).
3. An increase (decrease) in price volatility causes an increase (decrease) in both prices and inventory level. In the short run, we already showed in (b) how inventory level is affected. The effect on prices comes through the reduction in sales. Notice that in the absence of a price change, the production level remains constant; hence the only way to increase inventories is by keeping some output as inventory, which implies reducing sales. However, in a general equilibrium perspective, this will require a higher price, as we haven't assumed any shock to the demand (not modeled explicitly here). In the case of a permanent increase (decrease) in price volatility, a new higher (lower) steady-state price will result.

As Pindyck (2004) notes, one approach to continue in specifying the empirical model would be to first specify the stock-out costs and the costs of adjusting production and delivery scheduling and then solve for the marketing cost function (C^m) in the dynamic optimization model. Also, the dual approach and stochastic Itō calculus could be applied in the original optimization problem to get stochastic policy rules, but this would require data on supply and demand (for example, Pietola and Myers 2000). However, we take a more straightforward approach and specify reduced-form equations in order to provide empirical support to the inventory model and marketing cost function as described on page 7.

Since the time frequency on price–inventories relationship is one month in our analysis, we further simplify our model by approximating the discount factor by unity so that it drops out from the equations. This is merely rescaling of the estimating equation and can be done without loss of generality. We also impose a fixed storage cost *k*.

We proceed to empirical modeling with first testing for the stationarity properties of individual series. Then we test for the existence of the long-run equilibrium between the price and inventories. Since the data suggest that the price and inventories do not have a long-run equilibrium relationship, we move on to the short-run dynamics. The linkage between the price, inventories, and volatility is estimated in a vector autoregressive model, augmented by exogenous price and inventory shocks and the price volatility. The results suggest that the price and inventories have a significant negative relationship in the conditional mean process. The price Granger-causes inventories so that increasing price decreases inventory withdrawals and inventories increase in the very short run, but over a longer time period the price effect cancels out in the inventory process. In addition, an increasing price volatility encourages the build-up of inventories, as suggested by the economic inventory model. In some specifications, the linkage from inventories to the price is also significant through inventory shocks.

4. THE CONDITIONAL MEAN PROCESS

Stationarity

The first step is to test for stationarity and the order of integration of the individual price and inventory series. Both stationarity and nonstationarity are used as the null hypothesis because the unit root tests are known to have low power problems in small samples or in samples with structural breaks; therefore the results depend on the direction of the null hypothesis (for example, Leybourne and Newbold 2000). The null hypothesis of stationarity or trend-stationarity against the alternative of unit root is tested for by the η-test developed by Kwiatkowski et al. (KPSS test 1992). Using a null hypothesis of nonstationary unit root process against the alternative of stationary series is tested for Augmented Dickey-Fuller test (ADF test 1979) and Phillips-Perron test (PP test).

To obtain the test statistics, the mean processes of the price (xj, j=1), the inventory (xj, j=2), and the stock-to-use ratio (x^j, *j=3*) are estimated in the form

$$\Delta \ln x_{t+1}^{j} = \phi_0 + \rho \ln x_t^{j} + \sum_{i=0}^{k} \mu_i \Delta \ln x_{t-i}^{j} + S(t)^j + \varepsilon_{t+1}^{j}\,, \tag{6}$$

where $\phi_0, \rho,\ and\ \mu_{1+i}$ are parameters and ε_{t+1} is an error. Natural logarithms (*ln*) of variables are taken to get the random model variables to follow normal distributions more closely.[4] Logarithmic transformation is also justified by imposing the (asymptotic) relationship between the price and the inventories as required by the cost and convenience yield functions in equation (3).

We first test for seasonal effects in the conditional means by augmenting the estimating equation by two lagged differences (*k=2*), as suggested by Akaike-Scwartz criterion, and by monthly dummy variables in *S(t).* The seasonal effects turn out to be insignificant, however, and they are dropped from the estimating equations.[5]

Because the PP test uses the error structure in computing the test statistics, the lagged differences are dropped out in equation (6) by imposing a constraint μ_i=0 for all *i* in estimating the PP test statistics. The KPSS test statistics are obtained simply by regressing the price on intercept and the time trend.

Both ADF test statistics and PP test statistics suggest that the price, the inventory, and the inventory-to-consumption series all include a unit root; that is, the null hypothesis of unit root is not rejected (Tables 2 and 3). The test results remain unchanged whether we include or exclude the trend in the estimating equations. The results are also robust with regard to increasing the number of lagged differences in the estimating equations. Further, with regard to the price series for which we also have access to higher-frequency weekly data, the result remains unchanged when the monthly series is changed to weekly series.

The KPSS test rejects the null hypothesis of stationarity for price series, and the result does not change whether the trend is included or excluded in testing or the lag truncation parameter is changed. Thus, the test supports that the price series are nonstationary. Nevertheless, the test fails to reject the level stationarity of the end-of-season inventories and inventory-to-consumption ratio at 5 percent risk level when trend is not included; however, the test rejects the stationarity as the significant level extends to 10 percent. In addition, as the trend is included, the hypothesis of stationarity is rejected at the 5 percent level. Therefore, we conclude that all series appear to be nonstationary, with a possibility that the

[4] Nonnormality decreases the power of Dickey-Fuller tests to reject the unit root hypothesis, particularly if the sample size is small. Nevertheless, in a larger sample we can justify the approximation of the central limit theorem such that estimators are approximately normally distributed. Violation of the normality condition is quite common in empirical price models (for example, Roche and McQuinn 2003).

[5] LR-ratio test statistics for the monthly dummy variables were estimated in the price process at 3.69, in the inventory process at 10.0, and in the inventory-to-consumption process at 12.4. The corresponding chi-squared critical value with 11 degrees of freedom at 5 percent risk level is 19.7.

inventory data are not informative enough to generate fully consistent test results when the null hypothesis is reversed from nonstationarity to stationarity.

When the series are differenced, the resulting differenced series are all stationary and the null hypothesis of nonstationarity is rejected. Thus, the data suggest that when differenced, price, inventory, and inventory-to-consumption ratio are stationary I(0) processes.

Table 2. Unit root test statistics in the weekly and monthly price series (estimated in equation 1)

	Monthly price		Weekly price	
Test*	One lag (i=1)	Two lags (i=2)	One lag (i=1)	Two lags (i=2)
ADF, no trend	–1.153	–1.092	–0.933	–0.940
(critical value at 5% level)	(–2.888)	(–2.888)	(–2.860)	(–2.860)
ADF, with trend	–2.859	–2.838	–2.611	–2.709
(critical value at 5% level)	(–3.446)	(–3.446)	(–3.410)	(–3.410)
PP Z_ρ test, no trend	–1.171		–2.335	
(critical value at 5% level)	(–2.88)		(–2.88)	
PP Z_ρ test, with trend	–2.815		–2.24	
(critical value at 5% level)	(–3.44)		(–3.44)	
KPSS η-test, no trend**	1.18		1.792	
(critical value at 5% level)	(0.463)		(0.463)	
KPSS η-test, with trend**	0.169		0.710	
(critical value at 5% level)	(0.146)		(0.146)	

Source: Authors' estimation.
Note: * ADF, Augmented Dickey-Fuller test; PP, Phillips-Perron Z_ρ test; KPSS, η-test of Kwiatkowski et al. (1992).
** The test statistics decrease with an increase in the lag truncation parameters. Following Kwiatkowski et al. (1992), the lag truncation parameter was set at 8 and the test result remains unchanged even if the parameter value is increased.

Table 3. Unit root test statistics in the monthly series for end-of-season inventories and inventory-to-consumption ratios (estimated in equation 1)

	End-of-season inventories		**Inventory-to-consumption ratio**	
Test*	**One lag**	**Two lags**	**One lag**	**Two lags**
ADF, no trend	–1.832	–1.732	–2.044	–1.931
(critical value at 5% level)	(–2.888)	(–2.888)	(–2.888)	(–2.888)
ADF, with trend	–1.967	–1.845	–1.97	–1.841
(critical value at 5% level)	(–3.446)	(–3.446)	(–4.030)	(–3.446)
PP Z_ρ test, no trend	–1.82		–2.06	
(critical value at 5% level)	(–2.88)		(–3.44)	
PP Z_ρ test, with trend	–1.96		–1.98	
(critical value at 5% level)	(–2.88)		(–3.44)	
KPSS η-test, no trend**	0.337		0.220	
(critical value at 5% level)	(0.463)		(0.463)	
KPSS η-test, with trend**	0.241**		0.211**	
(critical value at 5% level)	(0.146)		(0.146)	

Source: Authors' estimation.
Note: * ADF, Augmented Dickey-Fuller test; PP, Phillips-Perron Z_ρ test; KPSS, η-test of Kwiatkowski et al. (1992).
** The test statistics decrease with an increase in the lag truncation parameters. Following Kwiatkowski et al. (1992), the lag truncation parameter was set at 8 and the test result remains unchanged even if the parameter value is increased.

The Long-run Cointegration Relationship

Given that the data are nonstationary, we now apply the Johansen trace statistics for binary cointegration on price and inventories in addition to the price and inventory-to-consumption ratio shown as equation (7):

$$\Delta \ln X_t = \Phi D_t + \Pi X_{t-1} + \sum_{l=1}^{L} \Gamma_l \Delta X_{t-l} + \varepsilon_t , \tag{7}$$

where X_t is a (2×1) vector to be tested for cointegration, which either represents price and inventory or price and inventory-to-consumption, Δ is the difference operator, $\Phi, \Pi,$ and Γ are parameter matrixes, and ε_t is a vector of white noise errors with $N(0, \Sigma)$. The number of lagged price and inventory differences and volatility is indexed by *l=1,…L.*

The Johansen procedure is used to examine the rank of the parameter matrix Π. The number of cointegrating vectors equals the rank of Π, which should be less than the number of tested series. In our case, the rank of Π should be smaller than 2. In Johanson's test, D_t is a deterministic term that includes a trend term in the cointegration relationship. Thus, Johansen's approach can examine the model specification with and without restriction of trend, starting with the restricted model without trend (Johansen 1992).

The Johansen trace statistics result is shown in Table 4. The result suggests that the price and inventories are not cointegrated, whether or not the trend is included. Thus, the price and inventories do not share a common long-term equilibrium during the time period and sampled data we have used.

Table 4. Johansen trace statistics for cointegration relationship between the price and inventories; two lags included in the estimating equations*

	s_t = the level of inventories		s_t = the inventories-to-consumption ratio	
Model specification	**Trace test statistics for the rank of 1**	**Critical value at 5%****	**Trace test statistics for rank of 1**	**Critical value at 5%****
no trend	7.87	15.49	7.51	15.49
with trend	19.99	25.87	18.47	25.87

Source: Authors' estimation.
* Increasing the number of lagged variables in the model does not change the result.
** MacKinnon et al. (1999).

Short-run Dynamics

Because the price and the inventories are not cointegrated and the data do not support a long-run equilibrium between them, we estimate the short-run dynamics by imposing $\Pi = 0$, which results in the following vector autoregressive model[6]:

$$\Delta \ln X_t = \Phi D_t + \sum_{l=1}^{L} \Gamma_l \Delta \ln X_{t-l} + \varepsilon_t \tag{8}$$

In the first-stage estimation, we exclude price volatility in the system and estimate it for the price and inventories only. As suggested by the likelihood ratio (LR) statistics, four lags of price and inventory differences are included. The estimation results suggest that price and inventories exhibit a significant negative short-run relationship (Table 5).

[6] To simplify our analysis, we use here and hereafter only the inventories-to-consumption ratio as the measure of inventories. Using the level of inventories results in similar estimates.

Table 5. Parameter estimates for the conditional mean in equation (8), price volatility excluded in estimation

	Price equation		Inventory equation	
	Parameter	Standard error	Parameter	Standard error
Intercept	0.00567	0.00664	0.00137	0.00405
Price				
once lagged	–0.0483	0.0882	–0.135	0.0539
twice lagged	–0.0797	0.0894	0.102	0.0546
third lagged	0.0655	0.0908	–0.00975	0.0554
fourth lagged	–0.160	0.0909	0.174	0.0555
Inventories				
once lagged	–0.124	0.140	0.00194	0.0852
twice lagged	–0.0691	0.140	–0.0680	0.0854
third lagged	–0.190	0.143	0.0652	0.0876
fourth lagged	–0.281	0.141	0.0641	0.0861
R squared	0.0667		0.126	

Source: Authors' estimation.

As indicated by the economic inventory model, we continue to augment the estimating equations by the price volatility $\sigma^2_{p,t}$, computing the volatility measure as a rolling average of lagged squared errors:

$$\sigma^2_{p,t} = \left(\sum_{i=0}^{I} \varepsilon^2_{p,t-i} \right) / I \tag{9}$$

where $\varepsilon_{p,t-i}$ refers to the error term in the first-stage estimation of the price equation in (8) and $i=1,2 \ldots I$ is the number of lagged errors. The specification with six lagged squared errors ($I=6$), that is, a rolling average spanning a period of a half year, fits the data better than other alternatives.[7]

The significance of exogenous price and inventory shocks is also tested by augmenting the model by lagged price and inventory shocks. More specifically, the error in equation (8) is replaced by

$$\varepsilon_t = \sum_{h=1}^{H} \varsigma_h \varepsilon_{t-h} + \nu_t \ ,$$

where ς is a parameter, $h=1,\ldots H$ is the number of lags, and υ_t is the new *i.i.d.* error.

Granger causality tests suggest in this specification that prices Granger-cause inventories, but not vice versa. The chi-squared distributed Wald test statistics for the null hypothesis of "inventories do not Granger-cause the price" is 5.54, which is smaller than the 9.35 critical value at 5 percent risk level. Thus, the null hypothesis cannot be rejected. The corresponding test statistics for the reverse direction of Granger causality and the null hypothesis of "the price does not Granger-cause inventories" is 9.74. Now the null hypothesis is rejected in favor of the alternative that the price Granger-causes inventories. Thus, the current and past inventories have no significant value to predict the next period price, once the current and past prices are controlled for. The result is consistent with the phenomenon that the inventory information is already in the prices when announced and that prices imply short-run inventory behavior.

[7] For example, a volatility specification over three lagged errors results in similar parameter estimates. Stationarity of the volatility measure was tested, and it turned stationary.

In the re-estimated two-equations model, once-lagged errors and volatility variables receive significant parameters (Table 6). Among these, the lagged error of the inventory equation (meaning an inventory shock) has a significant negative parameter in the price equation, while all other lagged errors are insignificant. Thus, the result suggests that the price responds more significantly to the unexpected part of the inventory movements (exogenous inventory shocks) rather than the observed movements. In other words, if new information on inventories reveals lower-than-expected inventories, the price responds upward. The result depends on the model specification; when the number of lagged prices and inventories is decreased, the inventory shocks become more significant in the price equation.

To address the question of whether the price response is different with regard to positive or negative inventory shocks, we further separate the lagged errors into positive and negative parts, but this kind of asymmetry does not significantly improve the model. Therefore, the results suggest that the price response is symmetric with regard to the positive and negative inventory shocks. Overall, if the parameter estimates are aggregated over all lags, the resulting aggregate effects of inventory movements and the inventory shocks on the short-run price movements are predicted to be small.

Table 6. Parameter estimates in equation (8); right-hand side augmented by lagged errors and volatility

	Price equation		Inventory equation	
	Parameter	Standard error	Parameter	Standard error
Intercept	0.00685	0.0106	–0.00724	0.00658
Price				
once lagged	0.691	0.620	–0.0628	0.386
twice lagged	0.142	0.147	0.0687	0.0915
third lagged	–0.0152	0.105	–0.0186	0.0651
fourth lagged	–0.192	0.101	0.146	0.0627
Price shock	–0.735	0.624	–0.0913	0.388
Price volatility	–1.357	1.378	1.715	0.857
Inventories				
once lagged	1.453	0.787	–0.144	0.489
twice lagged	0.0552	0.161	–0.106	0.100
third lagged	–0.0714	0.161	0.0391	0.100
fourth lagged	–0.184	0.159	0.0659	0.0990
Inventory shock	–1.627	0.795	0.111	0.494
R squared	0.106		0.148	

Source: Authors' estimation.

The price volatility gets significant parameter only in the inventory equation, suggesting that high price volatility significantly encourages increasing inventories or delaying inventory withdrawals. The result is consistent with the stock-out problem as suggested by the economic inventory model. The price volatility turns out, nevertheless, endogenous in the inventory equation. So we continue to respecify the model by adding a third equation for the endogenous price volatility in equation (8). As above, the LM statistics suggest including four lags in the right-hand side, but FPE (Akaike's Final Prediction Error criterion) and AIC (Akaike's information criterion) statistics support only two lags. The exogenous price and inventory shocks get insignificant parameters in the four-lag specification and they are dropped.

The pairwise Granger causality tests again suggest significant causal relationships, but each of them in only one direction. The price volatility causes price, the price causes inventories, and inventories cause price volatility, but not vice versa (Table 7). Also, the parameter estimates in the price equation suggest that there is a negative relationship between the inventory and price movements, but the price responds most significantly to the price volatility (Table 8). The signs of the volatility parameters alternate across the lags, indicating that, in the very short run, high volatility increases the price but thereafter decreases the price. The result supports the view that if high volatility increases the price, then the price volatility persists and the price can come down even faster.

Table 7. Wald test statistics for pairwise Granger causality

Direction of causality	Test statistics, Chi squared
from price volatility to price	11.70*
from price volatility to inventories	5.295
from price to inventories	11.38*
from price to price volatility	7.687
from inventories to price	4.338
from inventories to price volatility	11.55*

Source: Authors' estimation.
Note: * The null hypothesis of no-Granger-causality rejected at 5% risk level.

The price parameters in the inventory equation support a negative relationship between the price and inventory movements in the very short run, but over time these effects cancel out as the signs of the parameter estimates alternate. The volatility equation reveals a significant negative relationship between the inventory movements and the price volatility. When inventories decrease, the price volatility unambiguously increases.

Table 8. Parameter estimates in equation (8) with endogenous price volatility

	Price equation		Inventory equation		Volatility equation	
	Parameter	Stand. error	Parameter	Stand. error	Parameter	Stand. error
Intercept	0.0229	0.0103	–0.00649	0.00657	0.000527	0.000258
Price						
once lagged	–0.101	0.0937	–0.138	0.0597	–0.00108	0.00234
twice lagged	–0.145	0.0964	0.0614	0.0614	0.00438	0.00241
third lagged	–0.148	0.0970	–0.0460	0.0619	–0.00419	0.00242
fourth lagged	–0.159	0.0949	0.140	0.0605	0.000704	0.00237
Inventories						
once lagged	–0.150	0.141	–0.404	0.0902	–0.000167	0.00353
twice lagged	–0.00706	0.142	–0.115	0.0905	–0.00906	0.00355
third lagged	–0.201	0.150	0.00435	0.0954	–0.00700	0.00374
fourth lagged	–0.201	0.144	0.0520	0.0917	–0.00623	0.00359
Price volatility						
once lagged	2.862	3.756	–0.0846	2.395	1.00855	0.0938
twice lagged	–3.290	5.218	2.356	3.326	0.0446	0.130
third lagged	7.877	5.057	1.453	3.224	–0.102	0.126
fourth lagged	–10.271	3.626	–2.161	2.312	–0.030	0.091
R squared	0.153		0.155		0.879	

Source: Authors' estimation.

5. CONCLUDING REMARKS

The data suggest that the price of wheat and wheat inventories are nonstationary and that they are significantly linked to each other in the short run, but they do not exhibit a stationary long-run equilibrium relationship. Price volatility is an endogenous and important determinant in the short-run conditional mean processes for both the price and inventories. The pairwise Granger causality tests indicate, nevertheless, that causal relationships have only one direction each, so that inventories imply price volatility, the price volatility implies price, and the price implies inventories, but not vice versa.

The parameter estimates suggest unambiguously that when inventories decrease, price volatility increases. Price volatility further increases the price, but only in the very short run. Over time, the effect of volatility on the price is leveled off, and high volatility may also imply quickly declining price. In other words, the estimates signal that if high volatility increases the price, the price volatility persists and the price can come down even faster.

The price movements and inventories have a significant negative relationship in the very short run, but it is leveled off over time. Thus, in the very short run, increasing price implies decreasing inventories. Decreasing price implies either inventory build-ups or postponement of inventory withdrawals.

Broadly stated, the parameter estimates indicate the view that the current and past inventory and price movements are not very valuable in predicting the future price movements. It is likely that the inventory information announced each month is already in the prices. But if the inventories had a significant effect on the price, the price more likely responded to the exogenous inventory shocks rather than observed inventory movements.

We conclude that public inventory programs may be effective, but also controversial, means for decreasing price volatility. In the very short run, inventories unambiguously decrease price volatility, but in longer run they may have controversial price effects, as the data do not suggest a significant long-run relationship between the price and inventories. The long-run trade-offs between the price, inventories, and volatility in the global grain market, as required for designing preferred inventory policies, remain an open research question and are left for future research.

REFERENCES

Abbott, P. C., C. Hurt, and W. E. Tyner. 2008. *What's driving food prices?* Issue Report. Oak Brook, I.L.: Farm Foundation.

Cooke, B., and M. Robles. 2009. *Recent food prices movements: A time series analysis*. Discussion Paper No. 00942. Washington D.C.: International Food Policy Research Institute.

Deaton, A., and G. Laroque. 1995. Estimating nonlinear rational expectations commodity price model with unobservable state variables. *Journal of Applied Econometrics* 10: 9–40.

________. 1996. Competitive storage and commodity price dynamics. *Journal of Political Economy* 104 (5): 896–923.

Dewbre, J., C. Giner, W. Thompson, and M. von Lampe. 2008. High food commodity prices: Will they stay? Who will pay? *Agricultural Economics* 39: 393–403.

Dickey, D. A., and W. A. Fuller. 1979. Distribution of estimates for autoregressive time series with unit root. *Journal of American Statistical Association* 74: 427–431.

Heady, D., and F. Shenggen. 2008. *Anatomy of a crisis: The causes and consequences of surging food prices.* Discussion Paper 00831. Washington D.C.: International Food Policy Research Institute.

Johansen, S. 1992. Determination of cointegration rank in the presence of a linear trend. *Oxford Bulletin of Economics and Statistics* 54 (3): 383–397.

Kwiatkowski, D., P. Phillips, P. Schmidt, and Y. Shin. 1992. Testing the null hypothesis of stationarity against the alternative of a unit root. *Journal of Econometrics* 54: 159–178.

Leybourne, S. J., and P. Newbold. 2000. Behavior of Dickey-Fuller *t*-tests when there is a break under the alternative hypothesis. *Econometric Theory* 16: 779–789.

MacKinnon, J.G, A. A. Haug, and L. Michelis. 1999. Numerical distribution functions of likelihood ratio tests for cointegration. *Journal of Applied Econometrics* 14 (5): 563–577.

Miranda, M. J., and X. Rui. 1997. Maximum likelihood estimation of the nonlinear rational expectations asset pricing model. *Journal of Economic Dynamics and Control* 21 (8–9): 1493–1510.

Mitchell, D. 2008. *A note on rising food prices*. Development Prospects Group Policy Research Working Paper 4682. Washington D.C.: World Bank.

Pietola, K., and R. Myers. 2000. A dynamic dual model of investment under uncertainty with an application to structural adjustment in the Finnish hog industry. *American Journal of Agricultural Economics* 82 (November): 956–967.

Pindyck, R. S. 1994. Inventories and the short-run dynamics of commodity prices. *RAND Journal of Economics* 25 (1): 141–159.

________. 2004. Volatility and commodity price dynamics. *Journal of Futures Markets* 24 (11): 1029–1047.

Roche, M. J., and K. McQuinn. 2003. Grain price volatility in a small open economy. *European Review of Agricultural Economics* 30 (1): 77–98.

Von Braun, J., and Meinzen-Dick, R. 2009. *"Land grabbing" by foreign investors in developing countries: Risk and opportunities.* Policy Brief 13. Washington D.C.: International Food Policy Research Institute.

Von Braun, J., and M. Torero. 2008. *Physical and virtual global food reserves to protect the poor and prevent market failure*. Policy Brief 4. Washington D.C.: International Food Policy Research Institute.

Young II, R. E., L. Wilcox, B. Willot, G. Adams, and D. S. Brown. 1997. Consumer food prices from the ground up. *American Journal of Agricultural Economics* 79: 1673–1676.

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