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The Analysis of Wheat Prices Using Multiple Structural Break-Point Cointegration Test

Summary: From 2005 to 2008, high volatility in the markets affected grain prices significantly. This high volatility in grain prices made many researchers curious, and many discussions aroused from this topic. This study analyzes wheat price behavior during this period of high volatility. We estimate a return index for wheat using spot and futures wheat prices with the help of a present value model. To analyze the cointegration between the wheat prices and return index, a new cointegration test with multiple structural breaks, developed by Daiki Maki (2012), is used. The long-run cointegration coefficients are estimated using the Dynamic Ordinary Least Squares methodology. The empirical results show that there is cointegration between the spot and futures wheat prices, which tends to change at breakpoints. In other words, there is an equilibrium relation between spot prices and futures prices; however, it becomes unstable during the crisis in 2008. The results may help in understanding the dynamics of wheat prices, especially during high-volatility periods.

Keywords: Futures market, Multiple structural-breakpoints, Price explosiveness, Price discovery, Wheat market.

JEL: C32, G13, G14, Q11, Q14.

Volatility in commodities' prices, under certain limits, is helpful for the smooth and effective working of the markets. However, if it goes beyond these limits, it may have a negative impact not only on consumers and producers but also on the whole economy. High volatility increases market uncertainty. An uncertain market environment makes market manipulation easier and causes losses to market participants.

The reports of the Europen Union Commission showed that there was an increase in the overall volatility of different agricultural commodity markets from 2005 to 2008 and 2008 to 2011 (Europen Union 2011)[1]. These developments were also on the agenda in Cannes Summit, 2011 of G-20 – the leading actors of agriculture markets with a 65% market share in the global agriculture market – and it was decided to control the position limits in agriculture derivatives markets, increase transparency, and tighten the regulations in commodity markets (Hilary Till 2011).

Especially, wheat prices showed fluctuations ranging from -40% to +280% over the period signaling speculation in the commodities market. Some researchers argued

[1] **Europen Union.** 2011. http://europa.eu/index_en.htm (accessed January 20, 2011).

the presence of an unprecedented purchase pressure leading to several large bubbles in agricultural futures prices (e.g., Michael W. Masters 2008, 2009). However, academicians have previously tested the presence of bubble components at agricultural prices and found inconsistent results (Christopher L. Gilbert 2010; Scott H. Irwin and Dwight R. Sanders 2011; Luciano Gutierrez 2013; Celso Brunetti, Bahattin Büyükşahin, and Jeffrey Harris 2016).

Wheat is the main ingredient of food production. Because households spend a relatively high portion of their income on basic food and energy, price increases negatively affect the social well-being of consumers, especially in developing countries (Chris Brooks, Marcel Prokopczuk, and Yingying Wu 2015). Moreover, wheat is among the highly traded agricultural commodity in spot and futures markets. Therefore, it is important to determine whether the prices of this product deviate from a random walk and whether it exhibits explosive behavior.

Based on all the above-mentioned developments, this study aims to analyze the rapid rise of wheat prices in the futures market of Kansas Trade Exchange, one of the most important markets for world wheat trading.

1. Previous Studies

The factors causing the sudden changes in food prices during 2005 to 2008 found by previous studies can be grouped under the following five categories:

1. Factors causing stock changes (e.g., increase in the demand of non-food items due to bio-gas programs of the USA and the EU, increase in food demand because of high growth in some developing countries like China – see Derek Headey and Sehenggen Fan 2008; Mark W. Rosegran et al. 2008; Teresa Serra, David Zilberman, and José M. Gil 2011; Serra and Gil 2013).
2. Factors affecting the supply (e.g. global environmental conditions – see Hanjra A. Munir and Qureshi M. Ejaz 2010; Brian D. Wright 2011).
3. Factors affecting the use of inputs (e.g. changes in energy prices – see Herve Ott 2014; John Baffes and Tassos Haniotis 2016).
4. Factors affecting monetary policies (e.g. macro variables of developed countries – see Scrimgeour Dean 2015).
5. Speculative behaviors (Jenifer Piesse and Colin Thirtle 2009; Gilbert and Simone Pfuderer 2014; Phil Dawson 2015).

The study models the wheat price behavior with the help of structural breakpoints. There are many studies analyzing price behaviors in agri-products markets using speculative movements (Martin T. Bohl and Patrick M. Stephan 2013), mostly focusing on the changes in speculative positions and their effect on prices. Most of them find no extreme speculative movements (Sanders, Irwin, and Robert P. Merrin 2010; Sanders and Irwin 2011). There are very few studies focusing directly on price bubbles. An important problem in these studies is the inability of empirical tests to answer whether there were price bubbles or not. To solve this problem, Peter Phillips and Tassos Magdalinos (2009) and Phillips and Jun Yu (2011) developed a new unit root test for the presence of speculative bubbles. This method is developed to identify the bubbles that are even on a medium degree. Ulrich Homm and Jörg Breitung (2012) checked the success of this method using other tests. Gilbert (2010) applied this method on wheat, soya, and corn; however, except for soya, no bubbling phenomena were found in any of them. Gutierrez (2013) improved this method of Phillips, Yangru Wu, and Yu (2011) and tried to find price bubbles in agro-product markets using a non-parametric bootstrap method. The results of this study were exactly opposite to the results of Gilbert (2010). Bursts and crashes were found in the prices of wheat and corn but not in the prices of soya. In another study, price bubbles in wheat and corn were found using Momentum Threshold Autoregressive (MTAR) approach (Philipp Adöammer and Martin Bohly 2015). Similar results are found in the study of Peter Liu and Ke Tang (2010). It found price bubbles in corn, soya, and sugar (wheat was not considered in this study). Xiaoliang Liu, Gunther Filler, and Martin Odening (2013) used the regime shift approach but could not find any speculative bubble except for soya. Rıza Emekter, Benjamas Jirasakuldech, and Peter Went (2012) took a longer period (1961-2005) and identified bubbles in the prices of corn, wheat, and soya. This study is conducted using a time dependence test. Recently, Phillips, Shu-Pink Shi, and Yu (2013) developed a test that allows multiple bubbles. Xiaoli Etienne, Irwin, and Philip Garcia (2014) observed bursts in cereal markets by using this test. However, their results showed that the bubbles formed in agro-products are very small and loses their significance in 2 weeks.

In this study, cointegration tests, derived from the present value model, are used. In time series analysis, the cointegration test is developed to analyze co-movement between two non-stationary time series. If a linear combination of two or more non-stationary time series is stationary, then it can be said that the series are co-integrated. Many researchers in the field of finance, like Behzad Diba and Herscehel Grossman (1988) and Bohl (2003), studied the presence of price bubbles using unit root tests or cointegration between dividend and price data. These methodologies are based on the argument that if there is cointegration between the series of dividends and price, then the two series cannot move away from each other for an unlimited period and will attain equilibrium in the long-run. Nowadays, the generally accepted rule is that if there is cointegration among the data, then it can be taken as an indication of a price bubble. However, three important fundamental factors determine this relationship. These are unit root tests, long period present value model, and the presence of structural breaks (Vicente Esteve, Navarra Ibáñez, and Maria Asuncion 2013). However, this topic is still being discussed in the literature. Some of the researchers think these structural breaks are formed because of the effects of future expectations on market (see Zacharias Psaradakis, Martin Sola, and Fabio Spagnolo 2004), whereas others interpret them as result of cyclical political changes because of interest rates dynamics (see Eugene F. Fama and Kenneth R. French 2002). Whatever the reason is, these structural changes seriously question the validity of the efficient market hypothesis. Keeping in mind the different opinions in the literature, this study focuses on explaining the price instability and its structural breaks.

There are a couple of studies analyzing the relation between spot and futures prices in commodity markets (e.g. Holly H. Wang and Bingfan Ke 2005; Yunxian Yan and Michael Reed 2014). The distinguishing feature of our study is its focus on the wheat price dynamics in spot and futures markets under the structural break.

Another distinguishing feature of this study is the use of a contemporary residual-based cointegration test with multiple structural breaks, developed by Maki (2012), to analyze wheat price dynamics. This test works well both for the large and small number of breaks. This method is based on the structural breakpoint tests of Jussan Bai and Pierre Perron (1998) and the unit-root test of George Kapetanios (2005). This study uses a modified version of the code used by Maki (2012) to identify the breakpoints. Given the breakpoint dates, the coefficients of the long-run relation are estimated by the Dynamic Ordinary Least Squares (DOLS) method.

2. Test and Data

Monthly wheat price and storage cost data for the period of 1990:01-2012:12 are taken from the USA Kansas futures market (CME Group's Kansas City Board of Trade 2013)[2]. Data up to 2012 is focused since it includes extraordinary movements in wheat prices both in spot and futures markets especially around 2008 and 2011. The data for wheat prices in the spot market are taken from the official website of the USA agricultural ministry (United States Department of Agriculture Economic Research Service 2015)[3]. Descriptive statistics of the spot and futures wheat prices are summarized in Table 1[4].

Table 1 Descriptive Statistics for Spot and Kansas Trade Exchange Future Wheat Prices Time Series

	Spot	Future
Mean	418.7754	452.2717
Median	357.5000	375.4000
Maximum	1050.000	1255.000
Minimum	222.0000	202.0000
Std. dev.	169.1619	186.9571
Skewness	1.397266	1.532834
Kurtosis	4.320912	4.842666
Jarque-Bera	109.8735	147.1280
Sum	115582.0	124827.0
Sum sq. dev.	7869334	9612060
Observations	276	276

Source: Authors' calculations.

Table 1 shows that the average spot prices and futures prices are \$418 and \$452, respectively. The highest spot price is \$1050, whereas the futures price is \$1255 over the study period. Because the kurtosis values of both spot and futures prices are greater than 3, it shows a fat tail than the normal distribution. On the other hand, the skewness values of both spot and futures price show positively skewed distributions.

The expected price in the futures market is taken as the actual price. The prices of the first and last few days of a month, which cause high volatility, are excluded from

[2] **CME Group.** 2013. Kansas City Board of Trade. http://www.kcbt.com/historical_data.asp (accessed March 12, 2013).
[3] **United States Department of Agriculture Economic Research Service.** 2015. Wheat Data. https://www.ers.usda.gov/data-products/wheat-data/ (accessed February 02, 2015).
[4] Gauss 6 and EViews 7 are used to analyze the data.

the data set. The data of the futures contracts, whose maturities are close, are also excluded to avoid high manipulation.

2.1 Present Value Model and Wheat Stock Prices

The concept of efficient markets is based on rational commodity pricing. According to the present value (PV) model, the rational price of an asset can be obtained by discounting all the future cash flows of the asset at a predetermined discount rate. The PV model for a rational commodity price can be written as follows:

$$P_t = E_t\left[\sum_{i=1}^{k} \beta^i D_{t+i}\right] + E_t[\beta^k P_{t+k}], \tag{1}$$

where P_t is the rational price of a storable commodity at time t, P_{t+k} is the price of the commodity at time $t+k$, D_{t+i} is the convenience yield of the commodity at time $t+i$, β is the constant discount factor ($1/1+R$), and E_t represents an operator for the expected future value of the asset based on the information at time t (Robert Pindyck 1993). Convenience yield "d" which comprises carrying charges, like storage costs, insurance costs, and finance charges paid to store a physical commodity. It can also be defined as the utility, obtained by inventory holders, associated with availability in periods when supplies are scarce (Gutierrez 2013). The use Equation (1) along with the above-mentioned convenience yield definition to explain the pricing of storable commodities. Because the aggregate storage cost is positive for agricultural commodities, the above PV model can be used to price a storable commodity.

The expected present value of the entity approaches zero as time t goes to infinity. It can be represented as follows:

$$\lim_{k\to\infty} E_t[\beta^k P_{t+k}] = 0. \tag{2}$$

If the discount rate is lower than the growth rate in the profit share, the only solution for the basic value can be written as follows:

$$P_t = \sum_{i=1}^{\infty} \beta^i + E_t[D_{t+i}]. \tag{3}$$

The current price of an asset reflects information about the future. The PV model is suitable for calculating a rational commodity price because it incorporates expected future cashflows given the current information. Many solutions can be made from Equation (1), which results in a difference between the market price and the base price. The base value is the present value of the future profits an asset will make. Price bubble (φ_t) is the deviation from this base. It can be represented as follows:

$$P_t^* = P_t + \varphi_t. \tag{4}$$

2.2 Multiple Structural Breakpoints Cointegration Test

Allan W. Gregory and Bruce E. Hansen (1996a) argued that standard cointegration tests, including the tests of Robert Engle and Clive W. J. Granger (1987) and Søren Johansen (1988, 1991), give inadequate results under structural breaks. To overcome this problem, Gregory and Hansen (1996a, b) developed a test that allows a single structural break. However, the validity of the test is questionable as the number of breakpoints increases. Abdulnasser Hatemi-J (2008) developed a cointegration test with two structural breaks. However, both of the above tests lose their validity when there are 3 or more breaks. Maki (2012) developed a cointegration test that allows an infinite number of structural breaks. In our study, we applied this test to detect the presence of price bubbling phenomena in the agricultural commodity market. This residual-based test works both for a large and small number of structural breaks. This test is based on the structural breaks tests of Bai and Perron (1998) and the unit root tests of Kapetanios (2005).

In the Maki (2012) test, every period is taken as a possible breakpoint, *t*-statistics are calculated and the dates with the lowest *t*-statistics are taken as breakpoints. There are four possible model specifications.

Model 0: Model with a breakpoint in intercept and without a trend.

$$y_t = \mu + \sum_{i=1}^{k} \mu_i C_{i,t} + \beta' X_t + u_t; \quad (5)$$

Model 1: Model with a breakpoint in intercept and slope and without a trend.

$$y_t = \mu + \sum_{i=1}^{k} \mu_i C_{i,t} + \beta` X_t + \sum_{i=1}^{k} \beta_i' X_t C_{i,t} + u_t; \quad (6)$$

Model 2: Model with a breakpoint in intercept and slope and with the trend.

$$y_t = \mu + \sum_{i=1}^{k} \mu_i C_{i,t} + \gamma t + \beta` X_t + \sum_{i=1}^{k} \beta_i' X_t C_{i,t} + u_t; \quad (7)$$

Model 3: Model with a breakpoint in intercept, slope and trend.

$$y_t = \mu + \sum_{i=1}^{k} \mu_i C_{i,t} + \gamma t + \sum_{i=1}^{k} \gamma_t t C_{i,t} + \beta` X_t + \sum_{i-1}^{k} \beta_i' X_t C_{i,t} + u_t, \quad (8)$$

where $t = 1, 2, \ldots, T$. y_t and $x_t = (\mathrm{x_{1t}}, \cdots, \mathrm{x_{mt}})'$ denote observable I(1) variables, and u_t is the equilibrium error. y_t is a scalar, and $x_t = (\mathrm{x_{1t}}, \cdots, \mathrm{x_{mt}})'$ is an $(m \times 1)$ vector. The test assumes that an $(n \times 1)$ vector z_t is generated by $z_t = (y_t, x_t')' = z_{t-1} + \epsilon_t$, where ϵ_t are i.i.d. with mean zero, positive definite variance-covariance matrix $\sum$ and $E|\epsilon|^s < \infty$ for some $s > 4$. μ, μ_i, γ, γ_t, $\beta' = (\beta_1, \ldots, \beta_m)$, and $\beta_i' = (\beta_i, \ldots \ldots \beta_m)$ are parameters.

$C_{i,t}$ denotes a dummy variable. Maki (2012) defined this variable as:

$$C_{i,t} \begin{cases} 1, & if\ t > T_{Bi} \\ 0, & otherwise \end{cases}$$

Here, T_{Bi} represents the break years in the series. The hypotheses of the test are as follows:

$$\begin{cases} H_0\text{:} & \text{No cointegration under structural breaks} \\ H_1\text{:} & \text{Cointegration under structural breaks} \end{cases}.$$

The critical values, needed for hypothesis tests, are calculated using Monte Carlo simulation and given in Maki (2012). If the calculated value is less than the critical value, H_0 is rejected, and it can be concluded that there is cointegration among the series. A necessary condition for Maki's test is that the series should be I(1), that is, stationary in their first differences. The stationarity is tested using Josep Lluís

Carrion-i-Silvestre, Dukpa Kim, and Perron's (2009) structural breaks unit root test that allows multiple unknown breaks.

2.3 Long-Run Relationships

The logic behind cointegration is that there is a long-run relationship among some non-stationary variables. If two non-stationary variables have a stationary linear combination, then, in the long-run, they tend to move together, that is, there is cointegration between them. In this study, the coefficients of a long-run relation between D_t and P_t are estimated using DOLS. To remove endogeneity and heteroskedasticity problems in the residuals, James Stock and Mark Watson (1993) propose the use of the lags and leads of the explanatory variables in the model. The estimates derived using the above-mentioned methodology can be used to make statistical inferences. The above procedure is better as compared with Johansen and FMOLS methodologies. However, in Johansen's system of equations, parameter estimates of one equation may affect the residuals of other equations, whereas in DOLS, as there is only one equation, the above problem is out of the question. All the variables used in this study are stationary at I(1). Thus, the cointegration relation between them can be written as:

$$P_t = a_0 + \alpha_1 t + \alpha_2 D_t + \sum_{i=-q}^{q} \beta_i \Delta D_{t-i} + \varepsilon_t. \tag{9}$$

Here, α_2 is a coefficient of a long-run relation between D_t and P_t; q represents an optimal number of lags and leads. After estimating long-run relation using DOLS, in the second step, we look for cointegration relation among the variables using a test developed by Yongcheo Shin (1994). In this test, one of the two statistics, named C_μ, is calculated assuming no trend by putting $\alpha_1 = 0$. Whereas the second one, named C_t, is calculated with the trend by putting $\alpha_1 \neq 0$. The test statistic C_μ tests the null hypothesis of the presence of cointegration relation against the alternative hypothesis of the absence of cointegration. In the equation, q represents lag structure. The optimal number of lags and lead is obtained using Akaike Information Criteria (AIC).

3. Findings and Discussion

The stationarity of the series is tested using Augmented Dickey-Fuller (ADF) unit root test. Unit roots are found in all the series at level as the calculated test statistics are found to be greater than the critical values (Table 2). In other words, series are not stationary at their levels. They become stationary at the first difference, that is, all the series are I(1).

Table 2 Unit Root Test Results

	First difference			
	With intercept	**Prob.**	**With trend**	**Prob.**
P_t	-10.46785	0.0000	-10.50095	0.0000
	[-3.454174]		[-3.992029]	
D_t	-13.24608	0.0000	-13.22101	0.0000
	[-3.454353]		[-3.992029]	

Source: Authors' calculations.

Traditional ADF unit root testing does not take into account structural breaks. The test developed by Carrion-i-Silvestre, Kim, and Perron (2009) can perform unit root testing in the presence of structural breakage. Also in this test technique, effective results can be obtained in small samples. The stationarity of the series is tested using Carrion-i-Silvestre et al. (2009) structural breaks unit root test. Unit roots are found in both series at the level as the calculated test statistics are found to be greater than the respective critical values (Table 3). In other words, the series is not stationary at their levels. They become stationary at the first difference, that is, all the series are I(1). To study the cointegration relation among the variables, Maki's (2012) cointegration test is used.

Table 3 Carrion-i-Silvestre Structural Breaks Unit Root Test Results

	P_t				D_t			
	Level		First difference		Level		First difference	
	Test stat.	Critical value	Test stat.	Critical value	Test stat.	Critical value	Test stat.	Critical value
PT	23.358	8.571	6.406	8.399	11.825	8.589	3.288	8.376
MPT	19.925	8.571	5.557	8.399	9.814	8.589	2.983	8.376
MZA	-19.009	-43.813	-72.971	-46.417	-38.969	-44.090	-125.631	-43.982
MSB	0.162	0.106	0.083	0.103	0.113	0.106	0.063	0.106
MZT	-3.081	-4.675	-6.025	-4.829	-4.414	-4.690	-7.924	-4.689

Source: Authors' calculations.

This study analyzes the presence of price breaks in the futures wheat market by using multiple breakpoint cointegration test developed by Maki (2012). Structural breaking dates are determined with this test. It is recommended to use regime-switching regression models to analyze the structural change phenomenon (Brooks and Apostolos Katsaris 2003).

Model (2), which is called the regime-shifts model, allows for structural breaks of β in addition to μ.

So, model 2 and model 3 of Maki's test are used in this study. The results of the test are presented in Table 4. The test has obtained a minimum t-statistic using a simple structure.

The results show for all the models that the calculated test statistics are greater than the critical values, and thus under structural breakpoints, there is cointegration between the series. The abovementioned results show that at a 5% level of significance, a regression model with two breakpoints is significant for model 2 ($\tau > \tau^i \sim -5.37/-5.36$). However, at the same level of significance, a regression model with one breakpoint is significant for model 3 ($\tau > \tau^i \sim -5.99/-4.62$). Thus, in the Kansas futures market, there is a structural breakpoint in June 2008. However, according to the two breaks model, there are structural breaks in March 2008 and June 2008 in the same exchange. Furthermore, in model 2, with three breaks, another statistically significant breakpoint (at a 10% level of significance) is found in December 2007, along with the other two breaks. In model 3 with two breaks, statistically significant breaks (at a 10% level of significance) are found in April 2008 and July 2010. It needs to be mentioned that no statistically significant breaks are found after July 2010, even at a 10% level of significance (Table 4).

Table 4 Structural Breaking Test Results

Model	Break	Test statistics	Critic value			Structural breaking history
			*** 1%	** 5%	* 10%	
Model 2	One break	-4.99**	-5.45	-4.89	-4.62	2008:M6
	Two breaks	-5.37**	-5.86	-5.36	-5.07	2008:M3; 2008:M6
	Three breaks	-5.44*	-6.25	-5.7	-5.40	2007:M12; 2008:M3; 2008:M6
	Four breaks	-5.46	-6.59	-6.01	-5.72	1996:M4; 2007:M12; 2008:M3; 2008:M6
	Five breaks	-5.84	-6.91	-6.35	-6.05	1996:M4; 2007:M12; 2008:M3; 2008:M6; 2011:M3
Model 3	One break	-5.38*	6.04	-5.54	-5.28	2010:M7
	Two breaks	-5.92*	-6.62	-6.10	-5.85	2008:M4; 2010:M7
	Three breaks	-5.92	-7.08	-6.52	-6.26	2007:M11; 2008:M4; 2010:M7
	Four breaks	-5.92	-7.55	-7.00	-6.71	2002:M4; 2007:M11; 2008:M4; 2010:M7
	Five breaks	-6.20	-8.00	-7.41	-7.11	1993:M1; 2002:M4; 2007:M11; 2008:M4; 2010:M7

Notes: Critical values are taken from Table 1 of Maki (2012). From now on, we don't consider the breakpoint dates identified at a 10% level of significance in analyzing the long-run relationship among the series.

Source: Authors' calculations.

The main reason for not choosing the findings of model 3 is that the results are not statistically significant at a 5% significance level. Although the two breaks version of model 2 is significant at a 5% significance level, it shows structural breaks in the 3rd and 6th months of 2008. Moreover, both the single-break and two-break versions of model 2 point to June 2008. It is difficult to predict a long-term relationship over such a short regime with reasonable accuracy. Therefore, it is logical to focus on only one break with a smooth transition that started in March 2008. Wheat prices exploded in June 2008.

The coefficients of a long-run relation between P_t and D_t are estimated using a test, based on DOLS methodology. We do this using a test developed by Stock and Watson (1993) and Shin (1994) based on DOLS methodology. We use two sub-periods: 1990:1 to 2008:6 and 2008:7 to 2012:12. The results of cointegration tests are shown in Table 5.

Table 5 Long-Run Relationships DOLS

Deterministic cointegration	Model without structural breaks	One break	
	Full sample	Model II First rejime (1990:1-2008:6)	Model II Second rejime (2008:7-2012:12)
Parameter estimates			
α_2	2.47	2.90	1.52
a_0	4.07	2.49	5.72
Cointegration test			
C_μ	0.14*	0.007*	0.03*
W_{DOLS}	6.30**	15.95**	80.81**

Notes: * The main hypothesis about the existence of a cointegration relationship is not rejected at the 5% significance level ($C_\mu < 0.314$). ** For the W_{DOLS} test basic hypothesis is rejected at the 5% significance level ($\alpha_2 \neq 1$, $\chi^2 > 3.841$).

Source: Authors' calculations.

According to the results, we do not reject the null hypothesis of the presence of cointegration relation between the series, because test statistic C_μ calculated for the whole period (1990:1-2012:12) is lesser than the critical value at 5% level of significance (Table 5). So, we can say the series are cointegrated.

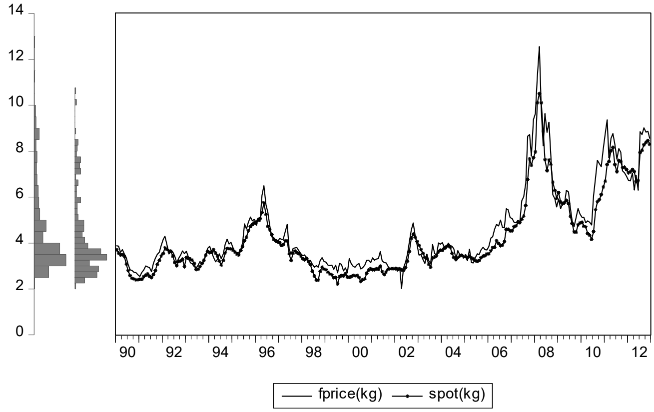
To check whether the cointegration relation is strong or weak, the Wald test is used to test whether the coefficient α_2 is equal to one or not. The null hypothesis of strong cointegration (α_2= 1) is rejected because the test statistic (W_{DOLS}) of the Wald test is greater than the critical value (3.841), taken from the Chi-square table, at a 5% level of significance. If we do not consider the structural breaks, the relation for the whole period (1990:1 to 2012:12) is found as [1, 2.47]. And if we consider structural breaks, the whole period is divided into two regimes because of the breakpoint identified in 2008:06. In the first regime (1990:01 to 2008:06), the null hypothesis of the presence of cointegration relation cannot be rejected because the calculated test statistic, C_μ, is lesser than the critical value (0.314) at 5% level of significance. The coefficient α_2 is also found to be different than 0 at a 5% level of significance for the said period. The null hypothesis of α_2= 1, in the Wald test, is rejected, implying weak cointegration. In the second regime (2008:07 to 2012:12), statistically significant cointegration is found between wheat price and return index. Analyzing both regimes shows there is a decreasing trend [2.90, 1.52] in the coefficients (α_2) of the long-run relation between P_t and D_t. The null hypothesis of strong cointegration, in the Wald test, is rejected for the said period also as the calculated test statistic W_{DOLS} is greater than the critical value (3.841) at a 5% level of significance. The empirical results show that the cointegrating relationship has changed over time, that is, it is not stable. There is cointegration between the spot and futures wheat prices, which tends to change at breakpoints.

The empirical results also show, in particular, the Maki tests for testing multiple structural breaks in cointegrated regression models suggest a model of three regimes, with the dates of the breaks estimated at 2018:M3 and 2008:M6, and a model of two regimes, with the date of the break estimated at 2008:M6. Accordingly, Kansas futures market wheat prices exploded in June.

Besides, the intrinsic bubble concept explains the asset prices deviating from the intrinsic/base value and the response of these deviations to the shocks (Kenneth A. Froot and Maurice Obstfeld 1991). Therefore, it is stated that tests with regime change are suitable for periodically exploding bubbles (Stephan G. Hall, Psaradakis, and Sola 1999). Based on the evidence, it can be said that the wheat price bubble, which originated in March 2008, collapsed in June 2008. The total duration of the wheat price exuberance was 3 trading months.

In another study, conducted on the wheat prices in the Chicago Board of Trade, the bubbles are identified in August 2007 and April 2008 (see Gutierrez 2013). Our results are in line with other studies indicating the presence of speculative bubbles in wheat prices (Liu and Tang 2010; Emekter, Jirasakuldech, and Went 2012; Roberto Espostia and Giulia Listorti 2013; Adöammer and Bohly 2015). Figure 1 shows the breakpoint relation in the spot and futures wheat prices in the Kansas Trade Market.

This study provides important evidence about the deviation of wheat prices from their market-based internal values from 2007 to 2008. In general, structural breaks are a problem for the analysis of economic series, because they are affected by either internal shocks or changes in policy regimes.



Source: Authors' elaboration.

Figure 1 The Breakpoint Relation between Spot and Futures Wheat Prices in Kansas Trade Market

The sub-prime credit crisis, starting in 2007, began to affect all the financial system of the USA rapidly, by March 2007, New Century Financial Corporation, the second-largest credit firm of the United States, along with many other finance and insurance companies had gone bankrupt. In 2008, the effect of the crisis increased further. The turbulence in the world market increased as the news that Bear Stearns, the world's largest investment and consultancy firm, was facing financial problems. So the crisis, starting from the United States, moved to the rest of the world. The financial crisis also affected the other sectors of the economy.

Overall, weak expectations in stocks and the housing sector can make a speculative investment in food products an attractive opportunity. Only part of the wheat production can be stored, the rest is sold quickly. Particularly low wheat stock levels encourage speculation on price increases more easily. It is observed that the stock levels are at very low levels (Piesse and Thirtle 2009) when the breaks are formed. Although this study determines where and when price breaks have occurred, further research is needed to understand which factors affect price dynamics.

4. Conclusion

This study analyzes price behavior in the KCBT wheat futures market based on bubbling phenomena. For this, the wheat price-return relation, taken from a present value model, is analyzed by using a new multiple structural breakpoint test of Maki (2012). The relation between structural breaks and cointegration is also considered to avoid the problems observed in the empirical literature. Furthermore, to estimate the cointegration coefficients of the relation between futures prices, taken from Kansas Trade Market, and returns, DOLS regression is used.

A significant cointegration relation is found in model 2 with one and two structural breaks. According to the results, in March 2008, the wheat prices in the Kansas futures market burst; and in June 2008, they crash. These structural breaks can be attributed to the global subprime crisis of 2007 to 2008.

The results obtained in our study are consistent with the existence of linear cointegration between the wheat prices and the returns, with a vector [1, 2.47]. Thus, the cointegration vector is not [1, 1], as predicted by the theory. The empirical results show the cointegration relation between the series tends to change over time because of the structural breaks.

The results suggest that ignoring structural changes in the cointegration relationship may understate the extent of correlation between the spot prices and the return index because the response of the present-value future price to changes in returns increases over time. Summing up, there is cointegration between the spot and futures wheat prices; and the equilibrium relation, that is, the cointegration vector, changes over time. In other words, the relation between spot prices and futures prices becomes unstable during the crisis.

The academic research for speculative bubbles in the agricultural commodity markets is still limited as compared to stock markets. The results of the study will enrich the literature and broaden the view of the researcher studying the dynamics of commodity markets.

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