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Article

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FINANCIAL ECONOMICS | RESEARCH ARTICLE

Linkages among commodity futures prices in the recent financial crisis: An application of cointegration tests with a structural break

Yoichi Tsuchiya^{1*}

Abstract: In this study, we investigate the existence of long-term co-movements among the prices of commodity futures contracts. We use a cointegration test, which accounts for the presence of a structural break. We show that while there is a long-term relationship among agricultural and among non-agricultural commodity futures prices when a structural break is taken into account, there is no such relationship without allowing for a structural break. We also show that these break points, in fact, occur a few months before the recent global financial crisis. Although the previous literature broadly casts doubt on such price co-movements, our results confirm that market performance improved during the sample period.

Subjects: Econometrics; Investment & Securities; Macroeconomics

Keywords: commodity futures; price discovery; futures pricing; financial crises

JEL classifications: Financial Crises (G01); Contingent Pricing; Futures Pricing (G13); Information and Market Efficiency; Event Studies; Insider Trading (G14)

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His research interests are applied econometrics, economic forecasting and macroeconomics. He has published his papers in academic journals including *Applied Economics*, *Economics Letters* and *International Review of Economic and Finance*.

PUBLIC INTEREST STATEMENT

This study on price linkages among commodity futures contracts would be of interest to a wide range of individuals including policy-makers and investors. Commodities and their futures prices have attracted significant attention and have become increasingly important in recent years, because of their numerous advantages. For instance, they can reduce investor risks in financial markets and play an important role in explaining macroeconomic phenomena such as the early stages of recessions for advanced economies, boom and bust cycles for developing countries and sovereign defaults. A key factor in facilitating these advantages is price co-movements among commodities. Unless they are not related, we are not able to detect and use valuable information. Therefore, this study's findings are a starting point for investors and policy-makers and indicate that more attention should be paid to these price dynamics.

1. Introduction

Pindyck and Rotemberg (1990) demonstrate that the prices of several commodities such as wheat, cotton, copper, gold, crude oil, lumber and cocoa are likely to move together. They suggest that this co-movement of commodity prices is caused in part by the herd behaviour of traders in financial markets—a phenomenon called the excess co-movement hypothesis (ECMH).

Considerable attention has been paid to commodity and their futures prices and they seem to become increasingly important in recent years.¹ Gorton and Rouwenhorst (2006) show that investors in financial markets can reduce their risk to take advantage of commodity futures. Chinn and Coibion (2014) present optimistic results of the broader use of futures prices as predictors of subsequent spot price movements. Commodity prices have also arguably played a crucial role in explaining macroeconomic fluctuations. Hamilton (2009) discusses that a large increase in oil price can account for much of the early stage of the recent recession. Not limited to oil prices, Bosworth and Lawrence (1982) point out that the twin oil price shocks of the 1970s were accompanied by the twin food price shocks. Furthermore, some of developing countries have been dependent on primary commodities for a large part of their exports and suffered sharp boom and burst patterns as a result of commodity price changes. Reinhart and Rogoff (2009) suggest that spikes in commodity prices are almost invariably followed by waves of new sovereign defaults. In addition to commodity futures, Acharya and Gaikwad (2014), for example, investigate the impact of introduction of pre-open call auction on price co-movement in Indian stock exchanges.

It is an urgent issue for both investors and policy-makers to examine dynamics and linkages among commodity futures prices. Interdependencies between agricultural commodity futures prices have been investigated and past studies show mixed evidence while relatively less attention has been paid to non-agricultural commodity futures prices and existing studies show negative evidence of the ECMH among them.

Geoffrey Booth and Ciner (2001) find that the prices of the four agricultural commodities traded on the Tokyo Grain Exchange (TGE) move independently, thus indicating the rejection of the ECMH. Their data range from July 1993 to March 1998. Bhar and Hamori (2006) also analyse the interdependence of the futures prices of these four commodities using a more recent data-set, from August 1994 to December 2003. They find that the prices move independently for the total sample period. Their results also imply that there are price co-movements from 2000 to 2003 but not during the 1990s. They suggest that this difference could reflect the notion that price discovery, as emphasized by Malliaris and Urrutia (1996), helps the market function better. Dawson and White (2002) show no interdependence among soft commodities such as barley, cocoa, coffee, sugar and wheat on the London International Financial Futures Exchange. On the other hand, Malliaris and Urrutia (1996) show a long-term relationship among the six commodity futures contracts (corn, wheat, oats, soybean, soybean meal and soybean oil) traded on the Chicago Board of Trade from January 1981 to October 1991, which is consistent with the ECMH.

Tsuchiya (2010) shows that the prices of the four precious metal commodity futures traded at the Tokyo Commodity Exchange (TOCOM) do not move together in the long-run. The data range from May 2002 to May 2010. Ciner (2001) also examines the long-run relationship between the prices of gold and silver futures contracts traded at the TOCOM and find that there is not a long-term relationship among the two in the period from 1992 to 1998. Ciner (2002), however, finds evidence suggesting that trading volume for gold, platinum and rubber futures contracts traded on the TOCOM conveys valuable information to the market about absolute value of returns.

This study aims to investigate whether long-term price co-movements exist in each commodity futures market while taking into account the presence of a structural break. The recent global financial crisis caused a significant negative level shift in those commodity futures prices and thus can be considered a structural change. If we do not assume it is a structural change, we would be unable to uncover long-term price co-movements because price linkages are probably broken during turmoil

periods. Although linkages among agricultural commodity futures prices have been examined following the findings of Pindyck and Rotemberg (1990), non-agricultural commodity futures prices have attracted limited research attention. Further, this study examines whether the functioning of these futures markets improved from their performance in the 1980s and 1990s. In particular, we use the price indexes for four precious metal and four agricultural commodity futures contracts traded on the TOCOM and TGE, respectively.

We find that there is a long-run relationship among the four precious metal commodity futures prices when a structural break is assumed; however, there is no such relationship without this assumption. The estimated date of the structural break is approximately one month before 15 September 2008, the starting point of the recent financial crisis called the “Lehman shock”. We also find similar results for the four agricultural commodity futures prices. The estimated date of the structural break is 26 August 2008, which is approximately three weeks before the Lehman shock. These results indicate that the precise date for the recent financial crisis is actually the Lehman shock. Our results also imply that the futures markets of precious metals and agricultural commodity have improved its functioning over the 10-year sample period, in line with the findings of Malliaris and Urrutia (1996) and Bhar and Hamori (2006).

The remainder of this article is organized as follows. Section 2 describes the data-set. Section 3 presents the results of the unit root tests. Section 4 describes the cointegration test with a structural break and presents the empirical results. Section 5 summarizes the findings.

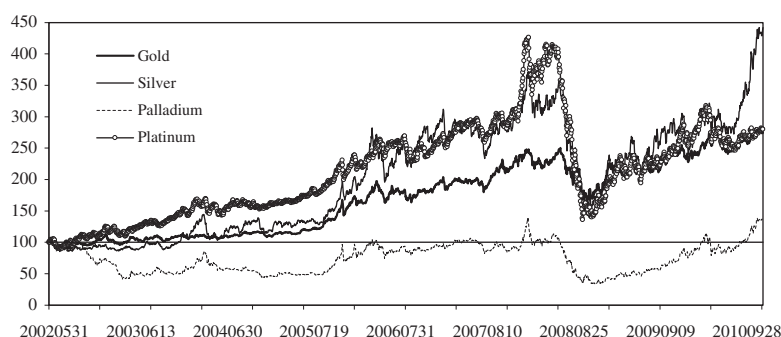
2. Data

We use the Nikkei-TOCOM Sub Commodity Indexes² for four precious metal futures contracts traded at the TOCOM: gold, palladium, platinum and silver. Each commodity listed at the TOCOM has its Sub Index. Each Sub Index is the figure of the fluctuation rates of the daily settlement prices in the core contract-month (fifth or sixth contract-month³), from base date (31 May 2002, on which the Index price is set at 100) to the applicable date. The sample period covers the end of May 2002 to the end of December 2010. The data comprise a total of 2104 observations for each contract with no observations missing.

We also use the TG Sub Indexes comprising the TG Indexes⁴ for the four agricultural futures contracts traded on the TGE: corn, soybean, red bean and sugar. Each commodity has its own Sub Index, which represents the fluctuation rates of the daily settlement prices in the core contract-month (furthest contract-months⁵), from the base date (31 March 2003, on which the Index price is set at 100) to the applicable date. The sample covers 31 March 2003 to 30 December 2010. The data comprise 1901 observations for each contract.

Figure 1 plots the four precious metal futures prices. For gold, platinum and silver, they have an upward trend since the beginning of the sample period and experience a large drop around the mid-July of 2008. Since then, they started to increase again. Note that the silver price increased

Figure 1. Nikkei-TOCOM sub commodity indexes.



considerably toward the end of the sample period. On the other hand, the palladium price has stagnated until a sudden drop around the mid-July 2008 and, however, it continued to increase afterward as the other three futures prices.

Figure 2 plots the four agricultural futures prices. In contrast to the precious metal futures prices, there is no clear trend. However, all except red bean exhibits similar price development since the mid-2007 with the previous metal futures prices.

Table 1 shows basic statistics of each commodity futures price. Among the precious metal futures, the prices of platinum and silver record more than four times high as the initial point of the sample while that of palladium stagnates. Silver varies the most among the four prices. Gold is the modest in terms of both price level and variability. The agricultural futures prices exhibit less divergence than the precious metal futures prices in terms of both price level and variability.

3. Unit root tests

First, we test stationarity for the level and first difference of the natural logarithm of the futures prices. We use two tests to see the robustness of the result: the augmented Dickey–Fuller (ADF) unit root tests and the Kwiatkowski–Phillips–Schmidt–Shin (KPSS) unit root tests. Then, we apply the Zivot–Andrews unit root tests with endogenous structural break detection.

Table 2 shows that the prices of all the commodity futures contracts are integrated of order one, $I(1)$. The ADF test⁶ tests the null hypothesis (that a variable contains a unit root) against the alternative: that a variable contains no unit root. We employ three specifications for the test: the auxiliary regression includes a constant and time trend (I); auxiliary regression includes a constant (II); and auxiliary regression includes none (III). For the price level, the ADF tests indicate that the futures price contains a unit root in every case. For the first difference of price, the tests indicate that there is no unit root in all cases.

Figure 2. TG sub-commodity indexes.

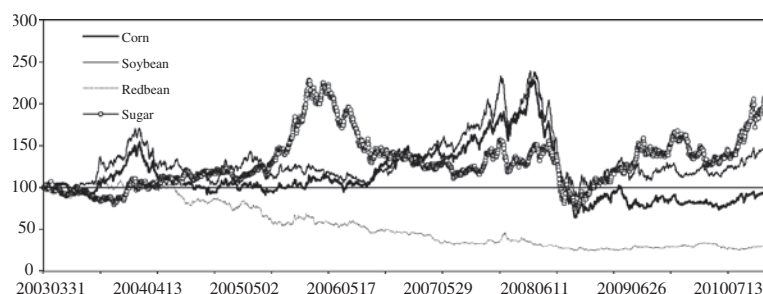


Table 1. Descriptive statistics

	Mean	Standard error	Coefficient of variation	Minimum	Maximum
<i>Nikkei-TOCOM Sub Index</i>					
Gold	170.5	56.5	0.33	89.2	279.9
Palladium	74.0	22.6	0.31	33.7	140.1
Platinum	214.2	74.4	0.35	89.7	426.3
Silver	203.5	87.0	0.43	87.0	442.0
<i>TG Sub Index</i>					
Corn	112.8	31.1	0.28	64.0	229.3
Red bean	54.6	27.1	0.50	24.5	115.3
Soybean	131.8	28.6	0.22	88.4	239.4
Sugar	130.8	30.5	0.23	75.9	228.4

Table 2. Augmented Dickey–Fuller test

(a) Precious metal futures prices				(b) Agricultural futures prices			
Commodity	Specification	Lag length	Test statistics	Commodity	Specification	Lag length	Test statistics
<i>Precious metal futures price levels</i>				<i>Agricultural futures price levels</i>			
Gold	I	23	−2.32	Corn	I	1	−1.64
	II	23	−0.77		II	1	−1.50
	III	23	2.24		III	1	−0.24
Palladium	I	1	−1.74	Soybean	I	28	−2.35
	II	1	−1.28		II	28	−2.36
	III	1	0.31		III	2	0.31
Platinum	I	29	−2.47	Red bean	I	5	−2.16
	II	21	−2.10		II	1	−0.92
	III	21	1.18		III	1	−1.91
Silver	I	10	−2.03	Sugar	I	2	−1.73
	II	10	−0.58		II	2	−1.41
	III	10	1.62		III	2	0.87
<i>First difference of precious metal futures prices</i>				<i>First difference of agricultural futures prices</i>			
Gold	I	22	−10.98**	Corn	I	1	−28.79**
	II	22	−10.99**		II	1	−28.80**
	III	22	−10.73**		III	1	−28.80**
Palladium	I	1	−31.12**	Soybean	I	1	−27.94**
	II	1	−31.10**		II	1	−27.95**
	III	1	−31.10**		III	1	−27.95**
Platinum	I	20	−9.91**	Red bean	I	4	−17.89**
	II	20	−9.86**		II	4	−17.90**
	III	20	−9.77**		III	4	−17.80**
Silver	I	9	−15.15**	Sugar	I	1	−29.67**
	II	9	−15.15**		II	1	−29.68**
	III	9	−15.05**		III	1	−29.66**

**Null hypothesis can be rejected at the 1% significance level.

Notes: The auxiliary regression includes a constant and time trend (I), a constant (II) and none (III). Lag length is chosen according to the AIC with a maximum lag of 30.

Table 3 shows the results of the KPSS tests for the level, and the first difference of the natural logarithm of all the commodity futures prices. Unlike the ADF test, the KPSS test⁷ tests the null hypothesis that a variable contains no unit root against the alternative that a variable contains a unit root. We employ two specifications for the test: the auxiliary regression includes a constant and time trend (I); and auxiliary regression includes a constant (II). For the price level, the KPSS tests indicate that the futures price contains a unit root in every case. For the first difference of price, the tests indicate that there is no unit root in all cases.

We obtain consistent results from the ADF and KPSS tests. Therefore, this leads us to the conclusion that the prices of the four precious metals and four agricultural futures contracts are integrated of order one, $I(1)$.

Furthermore, we apply the Zivot–Andrews unit root tests with endogenous structural break detection since the recent financial crisis might have caused a structural break for those futures prices.

Table 3. KPSS test

(a) Precious metal futures prices			(b) Agricultural futures prices		
Commodity	Specification	Test statistics	Commodity	Specification	Test statistics
<i>Precious metal futures price levels</i>			<i>Agricultural futures price levels</i>		
Gold	I	2.04**	Corn	I	2.40**
	II	22.21**		II	2.94**
Palladium	I	1.09**	Soybean	I	1.29**
	II	2.81**		II	2.06**
Platinum	I	3.13**	Red bean	I	2.48**
	II	15.76**		II	20.41**
Silver	I	2.86**	Sugar	I	1.94**
	II	19.41**		II	4.51**
<i>First difference of precious metal futures prices</i>			<i>First difference of agricultural futures prices</i>		
Gold	I	0.0399	Corn	I	0.0762
	II	0.0398		II	0.1179
Palladium	I	0.099	Soybean	I	0.0541
	II	0.2538		II	0.0648
Platinum	I	0.0521	Red bean	I	0.0613
	II	0.1304		II	0.0898
Silver	I	0.0696	Sugar	I	0.0933
	II	0.0717		II	0.1003

**Null hypothesis can be rejected at the 1% significance level.

Notes: The auxiliary regression includes a constant and time trend (I) and a constant (II).

Perron (1989) showed that the inference drawn from a Dickey–Fuller type test becomes unreliable in the case of a potential structural break, and proposed models where the structural break point is assumed to be known. However, Zivot–Andrews (1992) pointed out the problem of setting the break point exogenously and proposed a test that circumvents that possibility by endogenously determining the most likely occurrence of a structural break. The estimation procedure is to choose the date of the structural shift for the null hypothesis of a unit root as shown below. The test statistics⁸ is the t-ratio as in Perron (1989):

$$t_{\hat{\alpha}^i} \left[\hat{\lambda}_{\inf}^i \right] = \inf_{\lambda \in \Delta} t_{\hat{\alpha}^i}(\lambda) \text{ for } i = \text{I, II, III} \quad (1)$$

where $\Delta = [0, 1]$. The three different models are

$$y_t = \hat{\mu}^{\text{I}} + \hat{\theta}^{\text{I}} DU_t(\hat{\lambda}) + \hat{\beta}^{\text{I}} t + \hat{\alpha}^{\text{I}} y_{t-1} + \sum_{i=1}^k \hat{c}_i^{\text{I}} \Delta y_{t-i} + \hat{\varepsilon}_t \quad (2)$$

$$y_t = \hat{\mu}^{\text{II}} + \hat{\beta}^{\text{II}} t + \hat{\gamma}^{\text{II}} DT_t^*(\hat{\lambda}) + \hat{\alpha}^{\text{II}} y_{t-1} + \sum_{i=1}^k \hat{c}_i^{\text{II}} \Delta y_{t-i} + \hat{\varepsilon}_t \quad (3)$$

$$y_t = \hat{\mu}^{\text{III}} + \hat{\theta}^{\text{III}} DU_t(\hat{\lambda}) + \hat{\beta}^{\text{III}} t + \hat{\gamma}^{\text{III}} DT_t^*(\hat{\lambda}) + \hat{\alpha}^{\text{III}} y_{t-1} + \sum_{i=1}^k \hat{c}_i^{\text{III}} \Delta y_{t-i} + \hat{\varepsilon}_t \quad (4)$$

where $DU_t(\lambda) = 1$ if $t > T\lambda$ and 0 otherwise, and $DT_t^*(\lambda) = t - T\lambda$ for $t > T\lambda$ and 0 otherwise.

Table 4 shows the results of the Zivot–Andrews tests⁹ for the level and first difference of the natural logarithms of all the commodity futures prices. We use three specifications for the test: the auxiliary regression includes a constant and time trend (I); auxiliary regression includes a constant (II); and auxiliary regression includes none (III). Each specification corresponds to a model with the corresponding superscript for estimated parameters. The lag length of each specification for each variable is set as chosen in the ADF tests.

First, we focus on the precious metal futures prices. For the level, the hypothesis of a unit root is rejected in specification I and III of platinum and silver futures prices. For the first differenced series, the hypothesis of a unit root is rejected in each specification of all futures prices. In contrast to Dawson and White (2002), we obtain evidence that there is a structural break in each series.

The estimated break points for the futures prices of palladium and platinum lie between March 2008 and December 2008, which includes 15 September 2008, the starting point of the “Lehman shock”. However, those of gold and silver vary widely. The Zivot–Andrews test estimates 12 May

Table 4. Zivot–Andrews test

(a) Precious metal futures prices				(b) Agricultural futures prices			
Commodity	Specification	Test statistics	Break point	Commodity	Specification	Test statistics	Break point
<i>Precious metal futures price levels</i>				<i>Agricultural futures price levels</i>			
Gold	I	−4.44	14 September 2005	Corn	I	−5.43**	25 August 2008
	II	−2.88	27 March 2007		II	−2.21	4 September 2007
	III	−4.42	14 September 2005		III	−4.73	28 August 2008
Palladium	I	−2.97	11 July 2008	Soybean	I	−4.44	27 August 2008
	II	−2.25	8 July 2009		II	−2.50	27 September 2007
	III	−4.76	11 July 2008		III	−4.24	27 August 2008
Platinum	I	−7.00**	11 July 2008	Red bean	I	−4.05	9 July 2009
	II	2.68	18 January 2007		II	−3.78	23 October 2008
	III	−6.53**	4 September 2008		III	−3.98	27 May 2008
Silver	I	−5.32*	22 July 2008	Sugar	I	−2.64	19 September 2006
	II	−2.70	7 April 2006		II	−2.12	30 June 2010
	III	−5.32*	4 August 2008		III	−3.06	27 August 2008
<i>First difference of precious metal futures prices</i>				<i>First difference of agricultural futures prices</i>			
Gold	I	−11.14**	12 May 2006	Corn	I	−29.28**	25 June 2008
	II	−11.01**	2 August 2002		II	−29.05**	15 October 2008
	III	−11.20**	5 March 2008		III	−29.49**	3 July 2008
Palladium	I	−45.19**	3 March 2008	Soybean	I	−28.31**	22 February 2008
	II	−45.11**	29 September 2008		II	−28.22**	9 October 2008
	III	−45.33**	3 March 2008		III	−28.46**	3 July 2008
Platinum	I	−7.97**	1 December 2008	Red bean	I	−18.15**	14 November 2003
	II	−10.01**	26 September 2008		II	−18.09**	1 August 2005
	III	−10.49**	10 July 2008		III	−18.27**	14 November 2003
Silver	I	−15.43**	19 April 2006	Sugar	I	−30.16**	6 February 2006
	II	−15.35**	1 July 2010		II	−29.97**	30 April 2010
	III	−15.58**	14 July 2008		III	−30.17**	6 February 2006

**Null hypothesis can be rejected at the 1% significance level.

Notes: The auxiliary regression includes a constant and time trend (I), a constant (II) and none (III). The lag length of each specification is set as chosen in the ADF unit root tests (Table 2).

2006 as a break point for the futures price of gold in specifications I and 2 August 2002 in specification II. Furthermore, it estimates 19 April 2006 as a break point for the futures price of silver in specifications I and 1 July 2010 in specification II.

Next, we focus on the agricultural futures prices. For the level, the hypothesis of a unit root is rejected in only one specification (corn futures price). For the first differenced series, the hypothesis of a unit root is rejected in each specification of all futures prices. Therefore, it can be concluded that all prices are integrated of order one, $I(1)$.

The estimated break points for the futures prices of corn and soybean centre on 15 September 2008, but those of red bean and sugar do not. The Zivot–Andrews test estimates 14 November 2003 as a break point for the futures price of red bean in specifications I and III and 1 August 2005 in specification II. Furthermore, it estimates 6 February 2006 as a break point for the futures price of sugar in specifications I and III and 30 April 2010 in specification II.

Those dispersions of break points and rejections of the null for the price levels might indicate that there are more than two break points. However, investigating this issue is beyond the scope of the present paper.

4. Cointegration tests

We use Johansen's full information maximum likelihood cointegration analysis, in line with the past literature, to test for long-run co-movements. Johansen (1991) develops two likelihood ratios to test for the number of cointegration vectors in an unrestricted vector autoregression (VAR) model: the maximum eigenvalue test and the trace test. The former tests the null hypothesis of r cointegrating vectors against the alternative hypothesis of $r + 1$ cointegrating vectors. The latter tests the null hypothesis of r cointegrating vectors at most against the alternative hypothesis of more than r vectors. Neither test has a standard asymptotic distribution; however, Monte Carlo simulations based on critical values exist. The critical values are obtained from Osterwald-Lenum (1992).

To assess the robustness of our findings, we apply lag lengths of one and two, and we use three specifications of deterministic trends since the result of a cointegration test depends on these two factors. A lag length of two is selected according to the AIC. Since AIC is likely to choose a long lag length, a lag length of one is also examined.¹⁰ The specifications are as follows: the cointegrating equations have no intercepts and the level data have no deterministic trends (I); the cointegrating equations have intercepts but the level data have no deterministic trends (II); and both the cointegrating equations and the level data have linear trends (III).

Table 5 shows the results of the maximum eigenvalue tests and trace tests. The panel (a) shows the results of the four precious metal futures prices. The trace tests indicate that there are no cointegrating relations for all lag lengths and specifications. The maximum eigenvalue tests also show that there are no cointegration relations for all lag lengths and specifications. Therefore, we conclude that the four precious metal futures prices move independently during the sample period as shown in Tsuchiya (2010).

The panel (b) shows the results of the four agricultural futures prices. The trace tests indicate that there are no cointegrating relations for all lag lengths and specifications. The maximum eigenvalue tests show that there is no cointegration at the 1% significance level in two of the three specifications for both lag lengths, whereas specification III implies one cointegrating relation in both lag lengths. Therefore, we conclude that the four agricultural futures prices move independently during the sample period.

However, as shown in Table 4, the Zivot–Andrews unit root tests indicate an endogenous structural break for those futures prices. Therefore, we apply the method developed by Lutkepohl, Saikkonen, and Trenkler (2004), in which the structural shift is a simple shift in the level of the

Table 5. Cointegration tests

	Trace test			Maximal eigenvalue test		
	I	II	III	I	II	III
<i>(a) Precious metal futures prices</i>						
Lag = 1						
$r = 0$	32.50	28.00	45.79	13.73	12.71	22.41
$r \leq 1$	18.77	15.29	23.38	11.34	10.92	12.55
$r \leq 2$	7.44	4.37	10.83	4.32	3.30	7.91
$r \leq 3$	3.12	1.07	2.93	3.12	1.07	2.93
Lag = 2						
$r = 0$	34.28	30.18	49.31	13.73	13.46	23.72
$r \leq 1$	20.55	16.71	25.59	12.94	11.86	13.34
$r \leq 2$	7.61	4.85	12.25	4.46	3.47	9.03
$r \leq 3$	3.16	1.38	3.22	3.16	1.38	3.22
<i>(b) Agricultural futures prices</i>						
Lag = 1						
$r = 0$	39.88	34.17	60.09	25.44	25.20	45.82**
$r \leq 1$	14.44	8.97	14.28	6.92	5.00	6.46
$r \leq 2$	7.52	3.97	7.82	4.96	3.04	4.93
$r \leq 3$	2.56	0.93	2.89	2.56	0.93	2.89
Lag = 2						
$r = 0$	39.92	34.11	58.71	25.56	25.34	44.44**
$r \leq 1$	14.35	8.76	14.27	6.89	5.08	6.78
$r \leq 2$	7.46	3.68	7.50	4.91	2.85	4.79
$r \leq 3$	2.55	0.83	2.71	2.55	0.83	2.71

**Null hypothesis can be rejected at the 1% significance level.

Notes: I: The cointegrating equations have no intercepts and the level data have no deterministic trends.

II: The cointegrating equations have intercepts but the level data have no deterministic trends.

III: Both the cointegrating equations and the level data have linear trends.

process. This assumption is consistent with observations that most asset prices showed a large decrease during the crisis.

Lutkepohl et al. (2004)¹¹ assume that the $K \times 1$ vector process $\{y_t\}$ is generated by a constant, a linear trend and level shift terms:

$$y_t = \mu_0 + \mu_1 t + \delta d_{t\tau} + x_t \quad (5)$$

where $d_{t\tau}$ is a dummy variable defined by $d_{t\tau} = 0$ for $t < \tau$ and $d_{t\tau} = 1$ for $t \geq \tau$. The shift assumes that the shift point τ is unknown and that it is expressed as a fixed fraction of the sample size:

$$\tau = [T\lambda] \text{ with } 0 < \underline{\lambda} \leq \lambda \leq \bar{\lambda} < 1 \quad (6)$$

where $\underline{\lambda}$ and $\bar{\lambda}$ define real numbers and $[\cdot]$ defines the integer part. It is also assumed that the process $\{x_t\}$ can be represented as a VAR(p) and that the components are at most $I(1)$ and cointegrated with rank r .

The estimation of the break point is based on the regressions:

$$\mathbf{y}_t = \nu_0 + \nu_0 t + \delta d_{tr} + \mathbf{A}_1 \mathbf{y}_{t-1} + \dots + \mathbf{A}_p \mathbf{y}_{t-p} + \varepsilon_{tr} \text{ for } t = p+1, \dots, T \quad (7)$$

where \mathbf{A}_i with $i = 1, \dots, p$ assigns the $K \times K$ coefficient matrices and ε_t is the error process.

The estimator for the break point \hat{t} is defined as:

$$\hat{t} = \arg \min_{t \in \Psi} \det \left(\sum_{t=p+1}^T \hat{\varepsilon}_{tr} \hat{\varepsilon}'_{tr} \right) \quad (8)$$

where $\Psi = [T\lambda, T\bar{\lambda}]$ and $\hat{\varepsilon}_{tr}$ are the least-squares residuals of Equation 3. Once the break point \hat{t} has been estimated, the data are adjusted according to

$$\hat{\mathbf{x}}_t = \mathbf{y}_t - \hat{\mu}_0 - \hat{\mu}_1 t - \hat{\delta} d_{t\hat{t}} \quad (9)$$

Finally, Johansen's (1991) procedure for determining the cointegration rank can be applied to these adjusted series.

Table 6¹² shows the results of Lutkepohl et al.'s (2004) cointegration test with a structural break. The panel (a) shows one cointegrating relationship among the four futures prices in specification I and two cointegrating relationships in specification II. The estimated dates of the break points are

Table 6. Cointegration tests with a structural break

	I	Break point	II	Break point
<i>(a) Precious metal futures prices</i>				
Lag = 1				
$r = 0$	53.98**	14 July 2008	65.51**	1 August 2008
$r \leq 1$	21.62		31.86*	
$r \leq 2$	9.56		13.27	
$r \leq 3$	1.84		1.73	
Lag = 2				
$r = 0$	52.07**	14 July 2008	68.8**	1 August 2008
$r \leq 1$	23.40		34.74**	
$r \leq 2$	10.29		14.39	
$r \leq 3$	2.08		1.98	
<i>(b) Agricultural futures prices</i>				
Lag = 1				
$r = 0$	63.03**	26 August 2008	69.41**	26 August 2008
$r \leq 1$	27.40*		31.42*	
$r \leq 2$	6.07		8.42	
$r \leq 3$	2.70		2.07	
Lag = 2				
$r = 0$	64.97**	26 August 2008	71.44**	26 August 2008
$r \leq 1$	28.34*		32.25*	
$r \leq 2$	5.90		8.45	
$r \leq 3$	2.66		1.99	

*Null hypothesis can be rejected at the 5% significance level.

**Null hypothesis can be rejected at the 1% significance level.

Notes: I: Equation 5 does not include a linear trend. II: Equation 5 includes a linear trend.

14 July 2008 in specification I and 1 August 2008 for specifications I and II, respectively. The estimated dates of the structural break are approximately one month and two weeks before the Lehman shock. This fact implies that the commodity futures prices are good indicators of macroeconomic conditions represented by those spot commodity prices; this finding is thus consistent with the results of Chinn and Coibion (2014). The estimated dates of the structural breaks differ by only two weeks. The number of cointegrating relationships and estimated dates of break points are robust within each specification. These results indicate that the precise date for the recent financial crisis is actually the Lehman shock. In contrast to the results by Johansen's (1991) cointegration analysis, we obtain evidence that there is at least one cointegrating relationship among the four precious metal futures prices. Therefore, we conclude that there are long-term relationships among the four commodity futures prices, which is in contrast to the findings of Ciner (2001) and Tsuchiya (2010). In addition, Figure 1 indicates an upward trend for the four precious metal futures prices. It thus seems more reasonable to consider the two cointegrating relationships in specification II.

The panel (b) shows two cointegrating relations among the four agricultural futures prices and the estimated date of the break point is 26 August 2008 in all model specifications. In contrast to Johansen's (1991) cointegration analysis, we obtain robust results. Therefore, we conclude that there are long-run relationships among the four commodity futures prices since there are cointegrating relations between them.

Furthermore, our results imply that the futures market of precious metals and agricultural commodity have improved its functioning over the 10-year sample period, in line with the findings of

Table 7. VECM results: estimates of α

(a) Precious metal futures prices

Dependent variable	$\Delta gold$	$\Delta palladium$	$\Delta platinum$	$\Delta silver$
Lag = 1				
e_1	-0.0005 (-0.12)	0.0079 (1.00)	0.0160 (1.77)	0.0229 (3.13)
e_2	-0.012 (-5.04)	-0.0256 (-5.69)	-0.0173 (-3.40)	-0.0166 (-4.02)
Lag = 2				
e_1	-0.0009 (-0.20)	0.0080 (0.95)	0.0156 (1.63)	0.0228 (2.93)
e_2	-0.0120 (-5.17)	-0.0238 (-5.72)	-0.0174 (-3.69)	-0.0165 (-4.32)

(b) Agricultural futures prices

Dependent variable	$\Delta corn$	$\Delta soybean$	$\Delta redbean$	$\Delta sugar$
Lag = 1				
e_1	-0.0205 (-7.82)	-0.0020 (-0.82)	-0.0135 (-4.12)	-0.0217 (-8.96)
e_2	0.0080 (3.14)	-0.0087 (-3.65)	0.0107 (3.36)	0.0150 (6.64)
Lag = 2				
e_1	-0.0219 (-4.30)	0.0052 (1.92)	-0.0111 (-5.53)	-0.0207 (-8.06)
e_2	0.0088 (3.46)	-0.0082 (-3.47)	0.0103 (3.25)	0.0160 (6.77)

Notes: Estimates of specification II for precious metals futures and those of specification I for agricultural futures prices are presented. The t-values are in parentheses.

Malliaris and Urrutia (1996) and Bhar and Hamori (2006). One of the possible reasons for the improvement of market functioning is that the depth of the market has increased as pointed out by Chinn and Coibion (2014).

Finally, we examine a vector error correction model (VECM) representation because the existence of cointegration implies that the model can be represented as a VECM. The VECM is represented as a VAR in the first differences of Equation 7, augmented by the error correction terms $\alpha \mathbf{e}$, where $\mathbf{e} = \beta \mathbf{y}$. α can be interpreted as speed-of-adjustment parameters, β is a normalized cointegration vector and \mathbf{e} represents deviation from long-term equilibrium. Therefore, the VECM illustrates that the short-term dynamics of the commodity futures prices in the system are influenced by the deviation from long-term equilibrium and adjust according to speed-of-adjustment parameters.

We examine whether these parameters are equal to zero because if they are all equal to zero, the long-term relationship does not exist. For the precious metal futures prices, we consider specification II because they seem to demonstrate an upward trend. In contrast, for the agricultural futures prices, we consider specification I because Figure 1 indicates no such trend. Note that there are two cointegrating relationships for both markets. We denote \mathbf{e}_1 the error correction term for the first cointegrating relationships and \mathbf{e}_2 for the second cointegrating relationship.

Table 7 shows the estimates for the speed-of-adjustment parameters.¹³ It indicates the suitability of the VECM because most of the estimates are significantly different from zero. Among the precious metal futures prices, the speed-of-adjustment of gold is low, while those of platinum and silver are high. Among the agricultural futures prices, speed-of-adjustment of soybean is low, while those of red bean and sugar are high.

5. Conclusion

We examine long-term price relationships among the four precious metal commodity futures contracts traded at the TOCOM and the four agricultural commodity futures contracts traded on the TGE using Johansen's cointegration analysis and one with a structural break developed by Lutkepohl et al. (2004). The recent financial crisis forces us to use the latter to investigate long-term co-movements among commodity futures prices; however, structural breaks have been rarely taken into consideration in the literature.

The cointegration tests indicate that there are long-run relationships among those commodity futures prices, allowing for a structural break. However, there are no such relations without that premise. Therefore, our finding supports the ECMH in the precious metals and agricultural commodity futures markets whereas past literature such as Ciner (2001) and Tsuchiya (2010) show negative evidence of the ECMH among non-agricultural commodity futures prices. The estimated date of the structural break is between the mid-July and the beginning of August in 2008 for the precious metals and the end of August for the agricultural commodity. These results indicate that the precise date for the recent financial crisis is actually the Lehman shock. This fact indicates that futures prices are good indicators of their spot prices, thus of macroeconomic fluctuations.

Our findings also imply that the futures markets have improved its functioning over the 10-year sample period, in line with the findings of Malliaris and Urrutia (1996) and Bhar and Hamori (2006). This might reflect the fact that the depth of the futures market has increased. In recent years, it is likely that there could be long-run relationships among commodity futures prices because of the growing use of futures markets. Further investigation of other meal and commodity futures markets is needed.

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Notes

1. See, for example, Chinn and Coibion (2014) for more discussions.
2. The Nikkei-TOCOM Commodity Index is the aggregated figure of all the Sub Index, multiplied by a weight percentage for each component.
3. This is because those contract-month are the most active. Unlike the Japanese futures contracts, the nearest contracts are usually the most active in the US futures contracts.
4. The TG Index is the aggregated figure of all the Sub Indexes multiplied by a weight percentage for each component. It represents fluctuations in domestic agricultural commodity prices and is used as a benchmark for agricultural commodity investors.
5. This is because those contract-months are the most active. Unlike Japanese futures contracts, the nearest contracts are usually the most active in US futures contracts.
6. See Dickey and Fuller (1979) for details.
7. See Kwiatkowski et al. (1992) for details.
8. The presentation of the Zivot-Andrews test with endogenous structural break detection follows Pfaff (2008).
9. Dawson and White (2002) also applied the Zivot-Andrews tests and showed that no significant structural break is present in any series.
10. The lag length of one is selected according to the BIC.
11. The presentation of the cointegration test with a structural break follows Pfaff (2008).
12. The lag lengths of two and one are chosen following the previous Johansen's cointegration tests.
13. The other specifications yield consistent results although they are not reported here.

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