Is the high crude oil prices cause the soaring global food prices?

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Abstract

This paper explores shared trends and shared cycles between crude oil market and global food commodity markets for wheat, rice, sugar, beef, coffee, and groundnut. The results of the paper indicate there is no evidence of shared trend and common cycle between the two markets, suggesting that each market responds differently to cycle generating shocks. This result implies that change in food commodity prices do not mimic systematically the cyclical behavior in crude oil price changes. Thus, non-oil factors, such as the speculative activities in commodity future markets, may have more important role in recent years food commodity price explosion.

قياس ارتباط التضخم العالمي في أسعار السلع الغذائية بارتفاع أسعار النفط

ملخص

تهدف الورقة إلى قياس تأثير ارتفاع أسعار النفط على ازدياد الأسعار العالمية لمجموعة من السلع الغذائية الأساسية يتم تداولها في أسواق السلع العالمية. توضح نتائج الدراسة بأنه ليس هنالك ما يدعم، بمستوى معتبر إحصائياً، تأثير أسعار النفط على ارتفاع أسعار السلع الغذائية في أسواق السلع العالمية. هذه النتيجة ترجح أهمية دور المؤثرات الأخرى مثل المضاربات في أسواق العقود المستقبلية للسلع، بالإضافة لتسارع وتيرة النمو الاقتصادي لدول كالصين والهند وتداعيات ذلك على الطلب العالمي للسلع الغذائية، هذا بجانب استمرار تحول المزارع الأمريكي لإنتاج الوقود الحيوي نتيجة لارتفاع أسعاره في الأعوام الماضية.

1. Introduction

The sharp increase in agricultural commodity prices over the past few years have raised concerns about global inflation in both developed and developing countries. In 2007 the food price index calculated by the Food and Agricultural Organization (FAO) increased by 40 percent, reflecting substantial increase in all agricultural commodity prices.

Analysts attribute such dramatic rise in food commodity prices to variety of reasons, including macroeconomic policies in G7 countries, as well as structural changes in the global demand for food commodities(1). The increase in energy prices and its ramifications on bio-fuel uses has also been viewed as important factor behind the recent rise in food commodity prices(2).

Beside the above mentioned factors, there is even more important factor believed to have a significant role in the global food commodity price rises of recent years. A recent report released by Institute for Agriculture and Trade Policy (IATP, 2008, page 40) stated that "Amidst the food price crisis, speculation is a major contributor to extreme price volatility.." Also a report released by the United Nations Food and Agricultural Organization (Food outlook, 2008), iterates a similar voice about the impact of speculative effects on food commodity prices: "Massive commodity market speculation has pushed the prices of wheat, maize, rice, and other basic foods out of the reach of hundreds of millions of people around the World".

Regardless, of what may have been the major cause behind the increase in global food prices, it seems more accurate to say that there are variety of conditions, including energy prices, speculative effects, and may be even permanent supply and demand side shocks to global food commodity production, that contributed to the soaring food commodity prices. The main objective of this paper is to investigate if the crude oil price has a significant role in shaping the dynamics of food commodity prices by exploring if there are common trend and common cyclical features linking the two markets. The paper employs a methodology developed by Engle and Kozicki (1993), Vahid and Engle (1993) designed for testing for common trends and common cycles in less persistent stationary processes(3). The methodology of shared stochastic trends and common cycles has been employed in the past on energy markets by Serletis (1994), Serletis and Herbart (1999), Plourde and Watkin (2000), and Serletis and Ricardo (2004).

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(1) Guillermo Calvo (20-June-2008) argue that recent years explosion of commodity prices is mainly due to the global financial crisis linked with excess liquidity in non-G7 countries and nourished by low interest rates set by G7 central banks. However, Jeffrey Frankel (25-March 2008) attributes the soaring commodity prices to the high rapid economic growth of China and India.

(2) As oil prices kept at high levels above US$ 70 a barrel in 2009 and above US$100 a barrel in most of 2008, US farmers have shifted to grow bio-fuel feed stocks to take advantage of rising demand for bio-fuel crops. More directly, energy prices can also influence agricultural prices as high energy prices raise production and transportation costs.

(3) The term cycle in Vahid and Engle methodology refers to the stationary remainder after subtracting the random walk trend.

While cointegration is also a relevant approach for tackling the long-term persistence and co-movements of the variables under investigation, it only deals with association of variables that are non-stationary. However, when the co-movements are of stationary nature, or when the common shocks are less persistent than in the case of unit root, the common feature approach becomes more suitable for the detection of common cyclical features. Thus, commonality in cycles implies that variables respond in similar pattern to cycle generating shocks.

The paper is divided into four sections. Section two explains the methodology of the research. Section three deals with the estimation results. In the final section we conclude the research findings.

2. Methodology

2.1 Stochastic Common Trend

To illustrate the common feature testing procedure developed by Engle and Kozicki(1993), Vahid and Engle (1993), we decompose each variable into a trend (w_t), a cycle (c_t), and stationary innovation (e_t), so that:

$$y_t = \beta_1 w_{1t} + \gamma_1 c_{1t} + e_{1t} \qquad (1)$$
$$x_t = \beta_2 w_{2t} + \gamma_2 c_{2t} + e_{2t} \qquad (2)$$

where (y_t) stand for a food commodity price (in our case), (x_t) is oil price, and β_i *and* γ_i $(i = 1,2)$ are the corresponding trend and cycle coefficients. Assuming that there is a common trend among the two series, so that $w_{1t} = \theta w_{2t}$, where θ is a factor of proportionality between the two trends, then a linear combination between the two price series, $y_t - \lambda x_t$ can be expressed as:

$$y_t - \lambda x_t = (\beta_1\theta - \beta_2\lambda) w_{2t} + (\gamma_1 c_{1t} - \lambda\gamma_2 c_{2t}) + (e_{1t} - \lambda e_{2t}) \qquad (3)$$

where λ is the linear combination factor linking the two prices. Given that the parameter, λ is such that $\lambda = \beta_1\theta / \beta_2$, then there is a unique linear combination of y_t *and* x_t such that the common trend between the two series no longer exists. In such a case equation (3) reduces to the cyclical components and the innovations, so that:

$$h_t = (\gamma_1 c_{1t} - \lambda\gamma_2 c_{2t}) + (e_{1t} - \lambda e_{2t}) \qquad (4)$$
where
$$h_t = y_t - \lambda x_t$$

Thus, to test for a common trend between y_t *and* x_t in the equations (1) and (2) we test for cointegration (ie., the null of no common trend against the alternative of a significant trend) in the equation:

$$y_t = \lambda x_t + h_t \qquad (5)$$

where h_t is stationary innovation defined in (4), and λ is the contegration vector.

To test for common trend between oil price and the food commodity prices we employed the nonparametric rank test developed by Breitung (2001). An advantage of using this test includes, unlike the more popular parametric cointegration tests, of Johansen (1988), it captures the nonlinear long-term relationship that may exist between the two sets of data, in addition it is robust to structural changes in the time series data.

In the following we give a brief exposition of the rank cointegration test.

Given the two variables $z_{1t} = f_1(x_{1,t})$, and $z_{2t} = f_2(x_{2,t})$ are both I(1) series, where $x_{1,t}$ *and* $x_{2,t}$ are observed, whereas $f_1(.)$ *and* $f_2(.)$ are monotonically increasing function but are unknown; nonlinear cointegration between $x_{1,t}$ *and* $x_{2,t}$ is computed when the difference between z_{1t} *and* z_{2t} is integrated of order zero, or $\mu_t = z_{1t} - z_{2t}$ is I(0).

Since the sequence of ranks is invariant to monotonic transformations of the original data, the unknown $f_1(.)$ *and* $f_2(.)$ can be replaced by the ranks, R(x) so that: $R(z_{1t}) = R(x_{1t}), \quad and \quad R(z_{2t}) = R(x_{2t})$.

To test for ranks cointegration we need to calculate the following two statistics:

$$k_T = T^{-1} \sup |d_t| \qquad (6)$$

$$\zeta_T = T^{-3} \sum_{t=1}^{T} d_t^{\,2} \qquad (7)$$

where $d_t = R(x_{1t}) - R(x_{2t}) \quad and \quad \sup|d_t|$ is the maximum value of $|d_t|$ over t=1,2,…T. The null-hypothesis to be tested is no cointegration, and it is rejected if the statistics are smaller than the critical values at an appropriate significance level. The statistics expressed in (6) and (7) depend on the assumption that $z_{1t} \quad and \quad z_{2t}$ are not correlated. To correct for the possibility of correlation, Breitung (2001) proposed corrections based on the size of the correlation. When the absolute value of the correlation coefficient of the two series is minimal or close to zero, the test statistic should be corrected so that(4):

(4) When the absolute value of the correlation coefficient is close to one, Breiting (2001) also indicate modification to the equations (8) and (9).

$$k_T^{\ *} = \frac{k_T}{\hat{\sigma}_{\Delta d}} \qquad (8)$$

$$\zeta_T^{\ *} = \frac{\zeta_T}{\hat{\sigma}^2_{\ \Delta d}} \qquad (9)$$

$$where \quad \hat{\sigma}^2_{\ \Delta d} = T^{-2}\sum_{t=2}^{T}(d_t - d_{t-1})^2$$

2.2 Common Cycle Analysis

The test for common cycle follows a similar approach as that of the common trend. Given that y_t *and* x_t are non-stationary processes of order one (I(1)), then each series can be reduced to stationary process, I(0), by detrending equations (1) and (2) so that:

$$\Delta y_t = \alpha_1 c_{1t} + \varepsilon_{1t} \qquad (10)$$

$$\Delta x_t = \alpha_2 c_{2t} + \varepsilon_{2t} \qquad (11)$$

where Δ is the first difference, $\alpha_i\,(i = 1,2)$ are the coefficients corresponding to cyclical components and ε_{it} are stationary error terms.

To test whether there is a common cyclical feature linking the two series we test if there is a linear combination such that:

$$u_t = \Delta y_t - \delta \Delta x_t \qquad (12)$$

which does not have the cyclical component. Thus, the common cyclical feature test includes minimization of equation (12) with respect to δ, or more formally:

$$s(\hat{u}_t) = \min_{\delta} \ s(\Delta y_t - \delta \Delta x_t) \qquad (13)$$

Engle and Kozicki (1993), show that equation (13) can be reduced to[(5)]:

$$s(\hat{u}_t) = \min_{\delta} \ \hat{u}' M_x \Delta x_t (\Delta x_t' M_x \Delta x_t)^{-1} \Delta x' M_x \hat{u} / \hat{\sigma}_h^2 \qquad (14)$$

(5) See Engle and Kozicki (1993), pages 370-371, for verification of equation (14).

where M_x is a projection matrix, such that $M_x = [I - \Delta x(\Delta x'\Delta x)^{-1}\Delta x]$ and

$\hat{\sigma}_h^2 = \hat{u}'M_x\hat{u}' / N$

Where N is the number of observations.

Since minimization of equation (14) requires nonlinear procedure because the parameter δ appears in the denominator and the numerator of the equation (14), then the estimation procedure can be carried out using nonlinear estimation approach employing Limited Information Maximum Likelihood (LIML) method. However, more simpler and asymptotically equivalent estimator (Engle and Kozicki , 1993) can be obtained by minimizing only the numerator of equation (14), which is equivalent to estimation of equation (12) using Two Stage Least Square (2SLS) after augmenting it with instrumental variables and then testing for legitimacy of the instrumental variables. Such a process involves two steps. First we employ 2SLS in the following:

$$\Delta y_t = \delta_0 + \delta_1 \Delta z_t + \delta_2 \Delta x_t + \mu_t \qquad (15)$$

Where ($z_t = x_{t-1}, y_{t-1}$) stand for instrumental variables. Then using the estimated residuals from equation (15) (that is $\hat{\mu}_t$) we use the OLS to conduct LM test statistic:

$$\hat{\mu}_t = w_0 + \lambda_1 \Delta y_{t-1} + \lambda_2 \Delta x_{t-1} + \lambda_3 \hat{e}_{t-1} + \varepsilon_+ \qquad (16)$$

With the LM statistic distributed as Chi-square with two degrees of freedom. So, the test of the LM statistic in (16) is a test for the legitimacy of the IV variables used in 2SLS estimates.

2.3 Serial Correlation

As discussed in the previous section, the test of the null hypothesis of a common cycle between variables reduces to testing the linear combination in equation (12). This is because Vahid and Engle (1993) indicate (page 344, proposition 1) if the first differences of non-stationary, unit root variables are serially correlated, there exists a linear combination of the differences that maintain only innovations if and only if the levels of the variables have common cycles. In other words, to test for a common cycle among a group of variables we need first to verify existence of a cyclical component in each individual price series, which is guaranteed if the price series passes the test of serial correlation.

To perform the serial correlation test we adopt the same approach of Engle and Kozicki (1993), by setting price changes in VAR specification:

$$\Delta y_t = \beta_1 + \beta_{11} \Delta y_{t-1} + \beta_{12} \Delta x_{t-1} + \beta_{13} \hat{e}_{t-1} + \varepsilon_1 \qquad (17)$$

$$\Delta x_t = \beta_2 + \beta_{21} \Delta y_{t-1} + \beta_{22} \Delta x_{t-1} + \beta_{23} \hat{e}_{t-1} + \varepsilon_2 \qquad (18)$$

Where $\hat{e}_{t-1}$ is the lagged error terms from the equation (16). The test for a serial correlation can be computed using the LM test, which is NR^2, where R^2 is the coefficient of determination and N is the sample size, so that NR^2 is distributed as chi-square with degrees of freedom equal to the number of lagged variables coefficients in the above equations.

3. Results

The analysis in this paper is based on monthly data for six food commodity prices, beside Brent crude oil price, during the sample period from October 1984 to September 2009 – see Figures (1) – (3), for graphical exposition of the price series. The food commodities included wheat, rice, sugar, beef, coffee, and groundnut. All price series are collected from Index Mundi website, which in turn is extracted from the IMF, Primary Commodity Price Tables[(6)].

To investigate the presence of stochastic trend (unit root) in each price series we employed two alternative unit root tests, the augmented Dickey-Fuller (ADF, 1981) test, which tests the null of a random walk, and the stationarity test developed by Kwiatkowski, Phillips, Schmidt, and Shin (1992), which is often referred as KPSS test. An advantage of using KPSS test beside the ADF test is that it is a safeguard against the low power evidence of ADF test when the roots are near unity and when the data is fractionally integrated. Since the objective is to test for the presence of a common trend between food commodity prices and crude oil price, then we need first to verify the evidence of significant stochastic trend (unit root) in each price level. Table (1) reports the ADF and KPSS tests results, which indicate all price series exhibit stochastic trend, at levels, but such a trend can be removed by the first-differencing transformation. As a result, the non-parametric rank cointegration test statistics, reported in table (2), do not support evidence of long term relation between food commodity prices and crude oil price as the null hypothesis of no cointegration cannot be rejected for all commodities in the table. The LM test statistics reported in table (3) reject the null hypothesis of no serial correlation feature in all price changes, except for the change in coffee price. Given that we established a significant evidence of serial correlation feature in food commodity prices and crude oil price, which in turn imply presence of cyclical feature in each individual price series, then table (4) reports the estimation results of the common cyclical feature between food commodity prices and crude oil price. To explore if a decoupling had occurred during the whole sample period between food prices and crude oil price, we also tested for common cyclical feature for a sub-sample period extending from Oct 1984 to Dec 1993 (100 observations), beside for the full sample period[7].

Results in table (4) reject the null hypothesis of a common cycle between all food commodity prices and crude oil price, except for the beef product which exhibit a common cycle behavior with crude oil during the sub-sample period.

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(6) http://www.indexmundi.com/

(7) The structural break as represented by the sub-sample period determined based on Chow test results. The test results are not reported in this paper, but available upon request.

Table(1):Unit Root Tests

Series	ADF Level	ADF 1st diff	KPSS Level	KPSS 1st diff
Wheat			1.10	0.018*
Rice	2.30	10.89*	0.24	0.004*
Sugar	3.04	13.17*	1.89	0.14*
Beef	1.45	10.26*	1.98	0.019*
Groundnut	4.63	8.21*	0.56	0.024*
Coffee	2.26	10.88*	0.58	0.010*
Crude oil	1.96	313.5*	2.87	0.022*

* Reject the null-hypothesis of a unit root at 5% significance level for the ADF test, and fails to reject the null of trend stationarity for KPSS.
Note: All series are log transformed.

Table (2): Rank Cointegration Test

series	K_T stat	ζ_T stat	Correlation (ρ_T)
Wheat	0.67	0.098	0.06
Rice	0.71	0.14	0.07
Sugar	0.68	0.15	0.02
Beef	0.66	0.11	0.04
Groundnut	0.71	0.11	0.12
Coffee	0.68	0.19	0.05

Note: Critical values for the test statistics are provided in Breitung (2001), table (1).

Table (3): Serial Correlation Feature

Dependent Variable	LM test Statistic (1)	LM test Statistic (2)
Δ Wheat	16.91*	13.37*
Δ Rice	37.09*	12.40*
Δ Sugar	19.74*	8.33*
Δ Beef	18.48*	10.24*
Δ Groundnut	43.32*	11.17*
Δ Coffee	5.97	7.81*

Note: LM statistic(1) and (2) refers respectively to equations (16) and (17).
*significant at 5% significance level.

Table (4): Common Cycle Test

Dependent Variable	LM test stat (Full sample)	LM test stat (Sub-period)
Δ Wheat	9.6*	13.2*
Δ Rice	22.6*	5.9*
Δ Sugar	19.5*	7.12*
Δ Beef	14.2*	2.23
Δ Groundnut	24.7*	7.79*
Δ Coffee	2.5	13.1*

*significant at 5% significance level.
Note: The sub-period extend from Oct 1984 to Dec 1993.

4. Conclusion

To explore the dynamic association between crude oil market and major food commodity markets for wheat, rice, sugar, beef, coffee, and groundnut, this paper investigates shared trends and shared cycles between the two markets, using monthly data sample, spanning from October 1984 to September 2009 (299 observations). The finding of the paper indicates that there is no significant evidence of shared common trend and common cycles linking the oil market to each of the food commodity markets. This imply that change in food prices do not mimic systematically the cyclical behavior in crude oil price changes. This also suggests that food commodity markets and crude oil market respond differently to shocks, and each market reacts to different shocks. Accordingly, crude oil price changes is not a good predictor of food commodity price changes, as non-oil factors such as speculative effects in future commodity markets, may have greater role in shaping the dynamics in food commodity prices.

Fig (1):Log of food commodities and oil prices

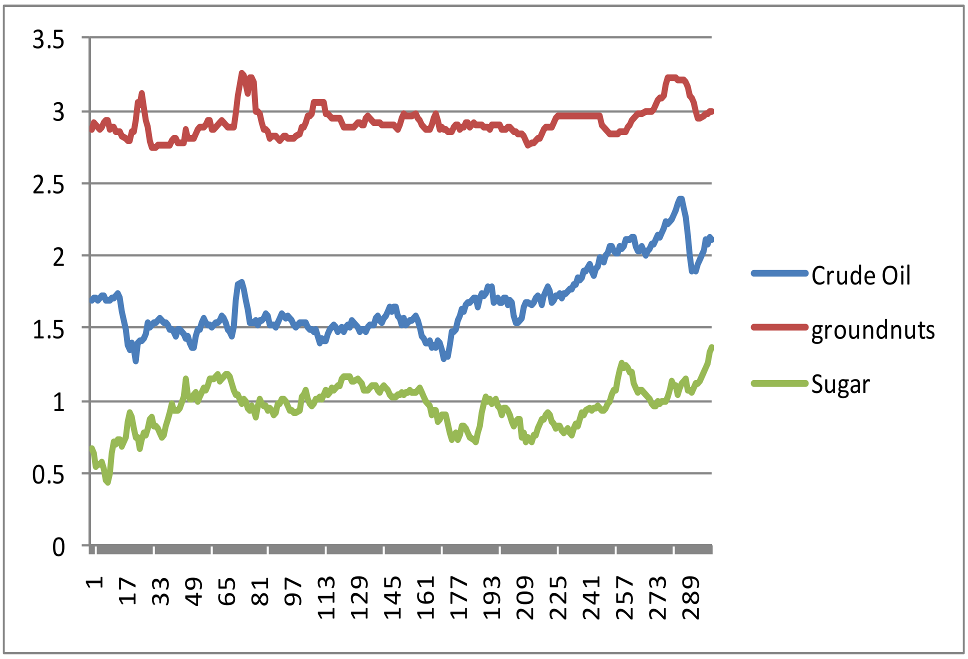


Fig (2):Log of food commodities and oil prices

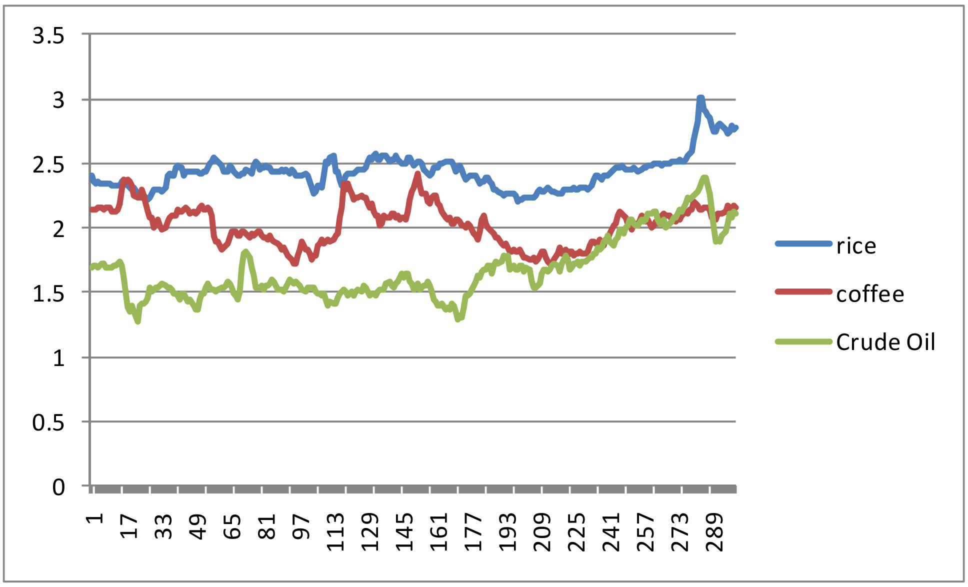
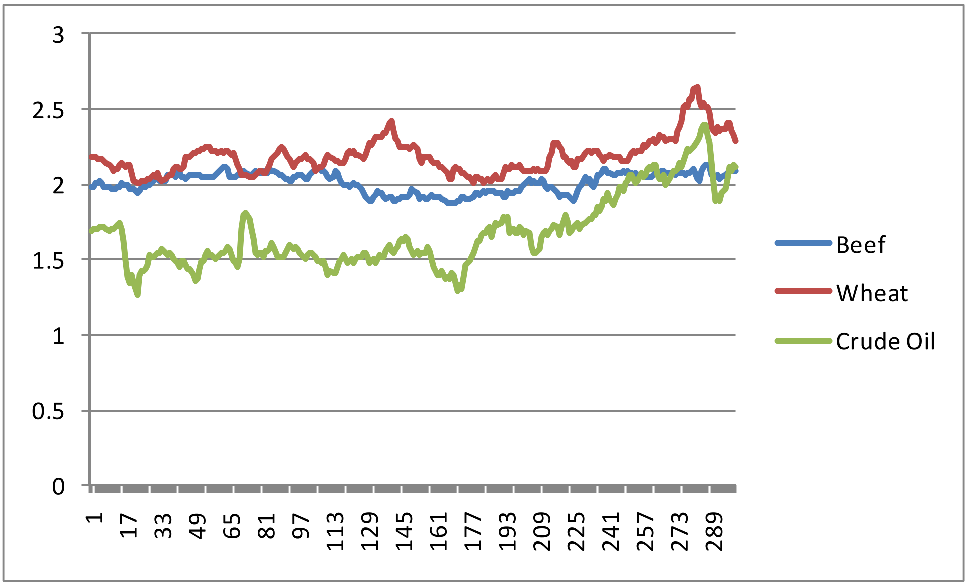


Fig (3):Log of food commodities and oil prices



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