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Kai-Yin Woo
Hing Lin Chan

June 2013

經濟及金融學系

Working Paper Series

Department of Economics and Finance

Hong Kong Shue Yan University

Working Paper Series
June 2013

ISBN: 978-988-16292-6-5
The URL is:
http://www.hksyu.edu/economics/working_paper/2013/Working%20Paper_2013_June_Woo.pdf

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Dr. Shu-kam Lee
Working Paper Coordinator
Department of Economics and Finance
Hong Kong Shue Yan University
10 Wai Tsui Crescent
Braemar Hill Road
North Point
Hong Kong
Fax: 2806-8044
Tel: 2806-5179 (Mr. Ken Tsui)
Email: sklee@hksyu.edu

An Investigation into the nonlinear relationship between the international and the domestic crude oil prices in China

Kai-Yin Woo*

Department of Economics and Finance

Hong Kong Shue Yan University

Hong Kong

kywoo@hksyu.edu

Hing Lin Chan

Department of Economics

Hong Kong Baptist University

Hong Kong

Keywords: Crude oil prices, nonlinearity, threshold cointegration, M-TAR

* Corresponding author

Introduction:

This paper attempts to study the relationship between the domestic crude oil prices of China with those prices in the international market. As China has recently replaced the US to become the largest oil importer in the world, it would be of interest to examine how the domestic and international prices affect each other. To understanding their relationships, it is necessary to know more about the reforming process of the Chinese crude oil prices in recent years.

Before the launch of the economic reform that took place in 1978, the state used state planning to allocate resources between competing users in the economy. In the absence of a fully functioning market, the crude prices were strictly controlled by the government and there was little adjustment in response to the demand and supply conditions of the oil industry. However, after the launch of reform in 1978, the regulation on oil prices becomes less restrictive. Starting from 1981, the state began to implement a "double-track" pricing system whereby the oil explorers could sell their products at a market price after they have fulfill their output targets laid down by the government. The largest oil field Da Qing, for example, was allowed to their products at international prices once its output target of 50 million tons was fulfilled (Wang et. al. 1991). Although this new pricing policy had encouraged the explorers to produce more than under strict planning, it has also generated widespread corruption. Therefore, the country entered into a new stage which began in 1994 and last until

1998, during which the dual track system was phased out and the state continued to control the setting of petroleum prices. However, the setting of oil prices was increasingly influenced by the market conditions. After 1998, domestic petroleum prices have been set by National Development and Reform Commission in accordance with the international energy market price due to the increasing reliance on imports to meet its energy oil demand (Hang and Tu, 2007). As a result of this change, the domestic oil prices have been staying increasingly close to the international prices since then.

The main purpose of the paper is to study the dynamic relationships between the international prices and domestic prices of crude oil. There are two reasons to study this issue. First, although it is the government policy to use international prices as a benchmark, the mechanism of this adjustment has never been spelt out. It is of our interest to see if the authority adjusted its domestic prices entirely in response to changes in the international prices without any delays. Second, as the import demand for oil has been growing as rapidly as its economy, China has now one of the largest oil importers in the world. In 2011, China accounted for 12% of the world oil import (British Petroleum, 2012). It would be interesting to see if there is any causality effect between the international and domestic prices. Knowing this relationship would help us forecast the movement of the domestic oil prices if the direction of causality is clarified.

However, since the adjustment process may be threshold asymmetric and nonlinear (Balk and Fomby, 1997), it motivates us to adopt the autoregressive distributed lag (ADL) test for cointegration proposed by Li and Lee (2010) which allow for nonlinear adjustment process in the threshold vector error correction model (TVECM). This method has been used in literature for the PPP testing (Chang, et al, 2012) but not attempted in the oil price analysis. The power of this method for cointegration test is much higher than the traditional cointegration tests with symmetric adjustment if the true adjustment process is asymmetric. In addition, it allows for threshold autoregressive (TAR) and the momentum threshold autoregressive (M-TAR) adjustments (Enders and Siklos, 2001). TAR adjustments allow the model to display differing speeds of autoregressive decay depending upon whether the discrepancies from equilibrium are larger or smaller than a threshold value and M-TAR adjustments allow the model to display speeds of error correction depending upon whether the growth of the discrepancies from equilibrium is increasing or decreasing. These threshold adjustments may be helpful if the Chinese authorities might take strong measure to offset shocks to the Da Qing in asymmetric manners. The two step approach of threshold cointegration suggested by Balke and Fomby (1997) involves the first step of cointegration test and the second step of threshold nonlinearity test. We use Li and Lee (2010) as a cointegration test and Hansen and Seo (2002) as a threshold nonlinearity test. After that, we estimate the

TVECM for describing the nonlinear error-correcting dynamics and Granger causality between Da Qing oil and Brent oil prices.

The rest of paper is organized as follows. The methodology is outlined in Section 2, the data description is done in Section 3, the empirical results are reported in Section 4, and then the main conclusions are summarized in Section 5.

Ⅱ. Methodology:

In this paper, we will adopt the method of Hansen and Seo (2002) to investigate the nonlinear threshold adjustment behavior that may exist between the international and domestic crude oil prices. To illustrate, let us consider an (n+1) dimensional observed series of price indexes $z_t = (y_t, x'_t)'$, at time t = 1,…,T, where y_t is a scalar and x_t is a n-vector of variables (in our case, n = 1). These two time series can be modeled by a 2-regime TVECM as shown below:

$$\Delta y_t = \pi_y d_t + \kappa_{11}\theta\ ' z_{t\text{-}1} I_{1\,t} + \kappa_{12}\theta\ ' z_{t\text{-}1} I_{2\,t} + \Phi_{1y}(L)\Delta y_{t-1} + \Phi_{1x}(L)\Delta x_{t-1} + \varepsilon_{1t} \quad (1a)$$

$$\Delta x_t = \pi_x d_t + \kappa_{21}\theta\ ' z_{t\text{-}1} I_{1\,t} + \kappa_{22}\theta\ ' z_{t\text{-}1} I_{2\,t} + \Phi_{2y}(L)\Delta y_{t-1} + \Phi_{2x}(L)\Delta x_{t-1} + \varepsilon_{2t} \quad (1b)$$

where d_t is a deterministic term which has a coefficient matrix $\pi = (\pi_y, \pi_x)'$; $\theta = (1, -\gamma')'$ is a cointegrating vector; $(\kappa_{11}, \kappa'_{21})'$ and $(\kappa_{12}, \kappa'_{22})'$ are the vectors of adjustment coefficients in regime 1 and 2, respectively; $\Phi(L) = (\Phi_1(L), \Phi_2(L))'$, which is composed of $\Phi_1(L) = (\Phi_{1y}(L), \Phi_{1x}(L))$ and $\Phi_2(L) = (\Phi_{2y}(L), \Phi_{2x}(L))$, is a p-th order polynomial matrix, and $\varepsilon_t = (\varepsilon_{1t}, \varepsilon'_{2t})'$ is an i.i.d. (n+1) vector of error terms. The coefficients in π and $\Phi(L)$ are assumed to be regime-invariant. However, the adjustment coefficients $(\kappa_{11}, \kappa'_{21})'$ and $(\kappa_{12}, \kappa'_{22})'$ are allowed to vary across regimes. Their values measure the proportion of the disequilibrium error of last period $e_{t-1} \equiv \theta' z_{t-1} = y_{t-1} - \gamma' x_{t-1}$ that has been corrected in the current period. The regime switching behavior of the adjustment coefficients is captured by the two Heaviside indicators I_{1t} and I_{2t}, where $I_{2t} = 1 - I_{1t}$. For the case of TAR model, we use regime 1 to denote those periods when e_{t-1} is smaller than the threshold variable $e^*_{t-1}(\tau)$, where $e^*_{t-1}(\tau)$ is the τ th percentile element of the empirical distribution of e_{r-1} with τ being selected by minimizing the sum of squared errors of model (1). Thus, $I_{1t} = 1$ when $e_{t-1} < e^*_{t-1}(\tau)$, and $I_{1t} = 0$ otherwise. Conversely, regime 2 appears when $e_{t-1} \geq e^*_{t-1}(\tau)$. For the case of M-TAR model, (Enders and Granger, 1998), the indicator I_{1t} is equal to 1 when $\Delta e_{t-1} < \Delta e^*_{t-1}(\tau)$ and equal to 0 otherwise, where $\Delta e^*_{t-1}(\tau)$ denotes the τ th percentile element of the empirical distribution of Δe_{t-1}. The choice of TAR and M-TAR model depends on how the market agents and policy makers react to the discrepancies in the equilibrium relationship between DQ and Brent oil price indexes. If the price adjustment is targeted at the levels of discrepancy rather than changes in the discrepancies from the equilibrium relationship, then TAR is more preferable than M-TAR and vice versa. Since there is no a priori theoretical justification as to whether TAR or M-TAR adjustments are correctly specified, both of these methods will be used in the empirical analysis.

In estimating the model (1), it is assumed that the variables in the system contain a cointegration relationship. However, since the error correction process of the system may be adjusted in a nonlinear manner, conventional cointegration methods may fail to detect their relationship. In response to this, we propose to adopt Li and Lee's (2010) single-equation ADL cointegration test to investigate the cointegration relationship between the Brent oil price index (Brent) and Da Qing oil price index (DQ). This method is proposed because we need to take the institutional factors of the Chinese pricing policy into account. In particular, since the policy of the Chinese government is to use the international oil prices as benchmark to adjust its DQ while the price of Brent is determined mainly by international market forces, it is reasonable to assume that the Brent, not DQ, is a weakly exogenous variable in model (1). To formulate the model, we use y_t and x_t to denote DQ and Brent in model (1) respectively, with x_t being assumed to be weakly exogenous to θ (Johansen, 1992), which means κ_{21} and κ_{22} of Eq.(1b) can be set equal to zero.[1] Therefore, model (1) can be modified into a model that is represented by a conditional model of Δy_t and a marginal model for Δx_t as shown below:

$$\Delta y_t = \pi_y d_t + B_1 z_{t-1} I_{1t} + B_2 z_{t-1} I_{2t} + \alpha' \Delta x_t + \Phi_1^*(L)\Delta z_{t-1} + \upsilon_{1t} \tag{2a}$$

$$\Delta x_t = \pi_x d_t + \upsilon_{2t} \tag{2b}$$

where $B_1 = \kappa_{11}\theta'$, $B_2 = \kappa_{12}\theta'$, $\Phi_1^*(L) = \Phi_1(L) - \alpha'\Phi_2(L)$, $\upsilon_{1t} = \varepsilon_{1t} - \alpha'\varepsilon_{2t}$ and $\upsilon_{2t} = \Phi_2(L)\Delta z_{t-1} + \varepsilon_{2t}$. Li and Lee (2010) propose using the ADL BO test to examine the

[1] The assumption of weak exogeneity of Brent to the system will be formally tested later.

cointegration relationship that may exist in the nonlinear threshold model as given in model (2).[2] This test is operated by examining the null hypothesis of $B_1 = B_2 = 0$ in the conditional model of Δy_t, i.e. Eq. (2a). Li and Lee (2010) proposed to use the supremum of Wald (sup Wald) statistics to conduct the test. The sup Wald statistic is free of nuisance parameters in the limit distributions and hence the ADL threshold tests do not need to use bootstrap procedure to derive their critical values. Instead, they can be obtained from simulation, which is much less time-consuming. In general, the critical values of the sup Wald depends upon the dimension of x_t, the type of indicators (TAR or M-TAR) and the type of deterministic term d_t that have been used.[3] The rejection of the ADL BO tests implies the existence of cointegrating relationship.

Once the cointegration relationship has been obtained, we can proceed to investigate the threshold effects in model (1) using supremum of Lagrange Multiplier (sup LM) tests proposed by Hansen and Seo (2002). However, due to the problem of unidentified parameters, the limit distribution of the sup LM statistics is unlikely to be standard. As a result, the p-values are obtained using the fixed regressor bootstrap and residual bootstrap procedures (Hansen and Seo, 2002). If the sup LM statistic significantly rejects the null hypothesis of no threshold effect, the existence of threshold adjustment effects can be confirmed. Also, in estimating the model (1), we

[2] The ADL BO test is extended from the linear ADL test of Boswijk (1994),
[3] We have employed three different types of deterministic terms, namely 0, 1, (1, *t*)' in the subsequent analysis to ensure our results are robust to the choice of deterministic terms.

can also check whether x_t (y_t) is weakly exogenous by testing the null hypothesis that $\kappa_{21} = \kappa_{22} = 0$ ($\kappa_{11} = \kappa_{12} = 0$).

Ⅲ. Data

To carry this analysis, we use the Brent oil prices (Brent) to proxy the international oil price and use the oil price of Da Qing, one of the largest oil fields, to represent the domestic oil prices. All the daily data of both Brent oil price index (Brent) from the US Energy Information Administration.[4] While the Da Qing oil price index is obtained from the website of International Oil Network (in Chinese).[5] The sample periods span from 2nd January, 2003 to 26th November 2012. The total number of observation is equal to 2517. All price indices are taken in natural logarithm and seasonally adjusted using X12 method.

IV. Empirical results:

Before conducting the cointegration tests, we apply four standard unit root tests to examine the null hypothesis of having a unit root against the alternative of stationarity for both Brent and DQ. A constant and a linear time trend are included in the test regressions. The four tests are: (1) the Augmented Dickey-Fuller (ADF) test, (2) the Phillips-Perron (PP) test, (3) the Elliott-Rothenberg-Stock (ERS) Dickey-Fuller with GLS de-trending (DF-GLS) method, and (4) the ERS point optimal (PO) test. As can

[4] Source: http://www.eia.gov/dnav/pet/hist/LeafHandler.ashx?n=pet&s=rbrte&f=d
[5] http://oil.in-en.com/quote/spot-oil-info.asp?cpname=Daqing

be seen from the results in Table 1, all the four tests that have been used cannot reject the null of a unit root process for the indices of Brent and DQ in level. However, they can significantly reject the null when the two indices are expressed in the first difference. As a result, the price indices under study are all I(1).

Table 1 Unit root tests

Tests	Brent				DQ			
	level	Lag	First difference	lag	Level	Lag	First difference	Lag
ADF	-2.227	0	-49.005*	0	-2.232	0	-51.500*	0
PP	-2.346	7	-48.995*	10	-2.296	4	-51.500*	6
DF-GLS	-2.182	0	-6.515*	13	-2.243	0	-4.843*	18
PO	9.320	0	0.072*	0	9.049	0	0.134*	0

Notes:
The bandwidth or truncation lag for PP test is chosen based on the Newey-West automatic selection method using Bartlett kernel. For all other tests, the choice of lags is based on Schwarz Criterion (SC).
The 5% and 1% critical values for ADF and PP are -3.411 and -3.961 respectively; for DF-GLS are -2.890 and -3.480 respectively; for PO are5.620 and 3.960 respectively.
* denotes significance at the 1% level.

After the unit root tests, we implement the single-equation ADL BO cointegration test proposed by Li and Lee (2010) on the assumption that Brent is weakly exogenous. We first estimate the error-correction term $e_{t-1} = DQ_{t-1} - c - \gamma' Brent_{t-1}$ using the OLS method:

$$\begin{array}{ccccccc} e_{t-1} & = & DQ_{t-1} & + & 0.02336 & -1.002 & Brent_{t-1}, \\ & & & & (-1.172) & (178.288) & \end{array} \tag{3}$$

where the figures in parentheses indicate the fully-modified OLS (FM-OLS) t-statistics (Phillips, 1995).

To continue, we estimate the conditional equation (2a) for both the cases of TAR and M-TAR adjustments. The number of lag length (p) is selected by using the Schwartz Criteria (SC). Three specifications of the deterministic term, namely $d_t = 0, 1,$ and $(1, t)'$, are included. As shown in Table 2, the BO Sup W statistics reject the null hypothesis of $B_1 = B_2 = 0$ at the 1% level for all cases, which provide evidence of a cointegration relationship that exists between DQ and Brent for both of the TAR and M-TAR adjustment processes.[6] The existence of cointegration relationship implies proportional movements between Brent and DQ in the long run. The estimated threshold values obtained from the estimation are 0.0169 in TAR and 0.0174 in M-TAR models, respectively, and they are the same for three specifications of deterministic term that have been used. Also, from equation (3), the FM-OLS t-statistics show that the null hypothesis of $c = 0$ cannot be rejected and the coefficient γ is significantly different from zero at the 1% level.[7] Since the FM-OLS t-statistics for the null hypothesis of $\gamma = 1$ is equal to 0.591, we conclude that the restriction of $\gamma = 1$ cannot be rejected.

[6] We have tried to assume weak exogeneity of DQ in model (2) and find the same results of cointegration (unreported). However, our weak exogeneity tests shown later support the weak exogeneity of Brent only.

[7] In case of cointegration, the FM-OLS t-statistics do not diverge and follow the standard distribution asymptotically so as to be valid for hypothesis testing.

Table 2 ADL BO cointegration tests

d_t	P	SC	Indicator	BO Sup Wald
0	2	-8.180	TAR	84.248*
1	2	-8.181	TAR	92.665*
$(1,t)'$	2	-8.177	TAR	92.665*
0	2	-8.175	M-TAR	71.379*
1	3	-8.198	M-TAR	64.042*
$(1,t)'$	3	-8.197	M-TAR	64.473*

Notes:

The 5% and 1% critical values of the ADL BO Sup W tests with TAR adjustments are 16.60 and 20.96 respectively when no deterministic term is included in the model (2); are 19.04, and 24.00 respectively when a constant is included; are 22.07 and 26.98 respectively when a constant and a linear trend are included.

The 5% and 1% critical values of the BO Sup W tests with MTAR adjustments are 15.65 and 19.98 respectively when no deterministic term is included in the model (2); are 18.66 and 23.88 respectively when a constant is included; are 21.44 and 26.15 when a constant and a linear trend are included.

* denote significance at the 1% level.

After confirming that these two price indexes contain a cointegration relationship, the next step is to estimate the model their non-linear threshold relationship. To do that, we test for the two-regime threshold effects in the model (1) under the null hypothesis of $(\kappa_{11},\kappa'_{21})' = (\kappa_{12},\kappa'_{22})'$ for both of the cases of TAR and M-TAR adjustments.[8] The sup LM statistics in the case of TAR adjustment are 31.681, with its fixed regressor bootstrap and residual bootstrap p-values are equal to 0.010 and 0.026, respectively. Similarly, the sup LM statistics in the case of M-TAR adjustment are 38.885 with both of its fixed regressor bootstrap and residual bootstrap p-values

[8] A constant is included in the model (1) as in Hansen and Seo (2002).

equal to 0.009. Hence, the sup LM statistics are all significant, indicating that the threshold nonlinearities in the error-correcting dynamics are observed for TAR and M-TAR adjustments. In addition, we have also conducted the weak exogeneity tests, which involves using the usual Wald statistics (Hansen and Seo, 2002; Seo, 2011) to examine if $\kappa_{21} = \kappa_{22} = 0$ and $\kappa_{11} = \kappa_{12} = 0$ hold in model (1). The Wald statistics for the null hypothesis of $\kappa_{21} = \kappa_{22} = 0$ are 4.616 and 2.748 for the cases of TAR and M-TAR adjustments, respectively, while for null hypothesis of $\kappa_{11} = \kappa_{12} = 0$ the corresponding statistics are 49.417 and 31.079, respectively. These tests imply that the Wald tests cannot significantly reject the weak exogeneity of Brent at the 5% level, but can reject the weak exogeneity of DQ for both of the cases of TAR and M-TAR.

The above analyses justify the estimation of the TVECM in the form of model (2) whose estimation results are reported in Table 3. The parameters are estimated by using seemingly unrelated regression (SUR). The threshold variable for TAR model is set equal to $DQ_{t-1} - Brent_{t-1}$ by imposing the restrictions *c* = *0* and γ = 1 on Eq.(3), while the threshold variable for M-TAR model is set equal to $\Delta DQ_{t-1} - \Delta Brent_{t-1}$. The estimated threshold value for TAR model is 0.0011 and for M-TAR is 0.0174.

For the conditional model of ΔDQ with TAR adjustment, we found that κ_{11} (-0.088) is larger than κ_{12} (-0.026) in absolute value and both of them are statistically significant. It indicates that DQ adjusts faster in response to the previous disequilibrium in regime 1 where $DQ_{t-1} - Brent_{t-1} - 0.0011 < 0$[9] than in regime two, and about 8.8% (2.6%) of the disequilibrium in regime 1 (regime 2) would be eliminated daily by the changes in DQ when DQ is lower (higher) than Brent. In other words, the DQ price adjusts faster to close the gap with the Brent when it is lower than when it is higher than Brent. Also, $\text{Wald}(\text{Brent} \rightarrow \text{DQ})$ is significant at the 1% level, indicating that the direction of Granger causality is running from Brent to DQ through the lagged changes in Brent. The coefficient of the current changes in Brent (0.396) is also significant, meaning that 1% increase in Brent oil price leads to 0.396% increase in Da Qing oil on the same day. In other words, changes in DQ are caused by current and lagged changes in Brent, and the short-run adjustments in response to disequilibrium. However, $\text{Wald}(\text{DQ} \rightarrow \text{Brent})$ in the marginal model of Δ Brent is insignificant, which leads us to conclude that Brent displays strong exogeneity to θ, i.e. DQ cannot Granger cause Brent through the error-correcting term and its lagged changes.

For the TVECM with M-TAR adjustment, κ_{11} and κ_{12} are both statistically significant. As in the case of TAR adjustment, κ_{11} (-0.056) is larger than κ_{12} (-0.039) in absolute term, suggesting that deviations from equilibrium (such that $\Delta DQ_{t-1} - \Delta Brent_{t-1} - 0.0174 < 0$) caused by increases in Brent are eliminated faster whereas decreases in Brent display slower speed of adjustments.[10] Moreover, similar

[9] There are 1,780 cases in total for regime 1 where $DQ_{t-1} - Brent_{t-1} - 0.0011 < 0$. Nevertheless, there are 1,742 cases where DQ – Brent < 0. Hence, most of the cases in regime 1 indicate that DQ is below Brent.

[10] There are 2,040 cases in total for regime 1 where $\Delta DQ_{t-1} - \Delta Brent_{t-1} - 0.0172 < 0$. There are 1,240

to the case of TAR, $\text{Wald}(\text{Brent} \rightarrow \text{DQ})$ in the conditional model of ΔDQ is significant at the 1% level, supporting the short-run form of Granger causality of Brent to DQ. However, unlike the TAR, $\text{Wald}(\text{DQ} \rightarrow \text{Brent})$ in the marginal model of ΔBrent is also significant, rejecting the strong exogeneity of Brent, i.e. DQ can Granger cause Brent through its lagged changes although this causality effect does not go through the disequilibrium term.

Table 3: Estimation of TVECM

A. TVECM with TAR adjustment

$$\Delta DQ_t = -0.0008 - 0.0887\, e_{t=1} I_{1t} - 0.0268\, e_{t=1} I_{1t} + 0.396\, \Delta Brent_t - 0.281\, \Delta DQ_{t-1}$$

(-1.926) (-7.301) (-2.359) (26.280) (-14.029)

$$-0.021\, \Delta DQ_{t-2} + 0.511\, \Delta Brent_{t-1} + 0.104\, \Delta Brent_{t-2} + \upsilon_{1t}$$

(-1.219) (28.693) (5.336)

$$\text{Wald}(\text{Brent} \rightarrow DQ) = 206.834 \ [0.000]$$

$$\Delta Brent_t = 0.0004 + 0.043\, \Delta DQ_{t-1} + 0.030\, \Delta DQ_{t-2} + 0.005\, \Delta Brent_{t-1}$$

(1.109) (1.666) (1.349) (0.246)

$$-0.020\, \Delta Brent_{t-2} + \upsilon_{2t}$$

(-0.818)

$$\text{Wald}(\text{DQ} \rightarrow \text{Brent}) = 3.609 \ [0.164]$$

B. TVECM with M-TAR adjustment

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cases where $\Delta DQ_{t-1} - \Delta Brent_{t-1} < 0$. Hence, over 60% of the cases in regime 1 indicate that ΔDQ is below ΔBrent.

$$\Delta DQ_t = -0.00004 - 0.056\,e_{t=1}I_{1t} - 0.039\,e_{t=1}I_{2t} + 0.394\,\Delta Brent_t - 0.299\,\Delta DQ_{t-1}$$
(0.137) (-6.782) (-2.511) (26.126) (-14.790)

$$-0.060\,\Delta DQ_{t-2} - 0.013\,\Delta DQ_{t-3} + 0.522\,\Delta Brent_{t-1} + 0.124\,\Delta Brent_{t-2}$$
(-2.913) (-0.753) (29.278) (6.083)

$$+0.060\,\Delta Brent_{t-3} + \upsilon_{1t}$$
(3.036)

$\text{Wald (Brent} \rightarrow \text{DQ)} = 863.564$ [0.000]

$$\Delta Brent_t = 0.0004 + 0.044\,\Delta DQ_{t-1} + 0.047\,\Delta DQ_{t-2} + 0.060\,\Delta DQ_{t-3}$$
(1.068) (1.729) (1.747) (2.680)

$$+0.003\,\Delta Brent_{t-1} - 0.029\,\Delta Brent_{t-2} - 0.025\,\Delta Brent_{t-3} + \upsilon_{2t}$$
(0.148) (-1.121) (-0.993)

$\text{Wald (DQ} \rightarrow \text{Brent)} = 10.105$ [0.017]

Notes:
SUR method is used for estimation.
The t statistics are shown in parentheses
Wald statistics follow the usual chi-squared distribution with degrees of freedom equal to the number of restrictions. The figures in the squared brackets are the p-values.
** and * denotes the statistical significance at the 1%, 5% and 10% level, respectively.

V. Conclusion:

This article explored the cointegration relationship between Da Qing oil and Brent oil prices using the newly developed testing methods of threshold cointegration with both TAR and M-TAR adjustments. In addition, this paper also tested for Granger causality of these two oil prices.

The results suggest that Da Qing oil and Brent oil prices are threshold cointegrated with the restricted unit cointegrating coefficient, indicating that these two oil prices move in the same proportion in the long run. It is also possible for the disequilibrium to eliminate in the TAR or M-TAR adjustment processes. Da Qing oil price adjusts faster in face of negative discrepancies in TAR model or in face of falling discrepancies in M-TAR model. Both indicate that the government may take measure the smooth out the discrepancies when the domestic price or its growth rate is below the international oil price. Also, the weakly exogeneity tests unsurprisingly show that Da Qing oil price is Granger caused by the Brent oil price in response to deviations from equilibrium and not vice versa.

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The working paper series is a series of occasional papers funded by the Research and Staff Development Committee. The objective of the series is to arouse intellectual curiosity and encourage research activities. The expected readership will include colleagues within Hong Kong Shue Yan University, as well as academics and professionals in Hong Kong and beyond.

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