Estimating Relative Price Impact: The Case of Brent and WTI

Shiyu Ye
Graduate Student
Department of Agricultural & Applied Economics
The University of Georgia
ysy319@uga.edu

Berna Karali
Associate Professor
Department of Agricultural & Applied Economics
The University of Georgia
bkarali@uga.edu

Selected Paper prepared for presentation at the 2016 Agricultural & Applied Economics Association Annual Meeting, Boston, Massachusetts, July 31-August 2

Estimating Relative Price Impact: The Case of Brent and WTI

Abstract

Brent and Western Texas Intermediate (WTI) are the two dominant benchmarks in global crude oil markets. Historically, the prices of Brent and WTI moved together and their price spread was narrow and negative. Since 2010, WTI has traded at a sizable discount compared to Brent, resulting in a widening and positive spread. This paper tests for possible structural breaks in the Brent-WTI spread during 1993-2016 time period using the Bai-Perron (1998) test and measures the price impacts of the shifts found in the equilibrium relationship using the relative price of a substitute method of Carter and Smith (2007). While the spread is found to experience multiple structural changes during the sample period, the price impact of the break occurred in the later time period is found to be larger.

Key words: Brent, cointegration, crude oil, relative price of substitute, structural break, WTI

Estimating Relative Price Impact: The Case of Brent and WTI

1. Introduction

In global oil markets, the two dominant benchmarks are Brent and West Texas Intermediate (WTI) for light crude oil. Brent is produced in the North Sea and plays a key role in pricing of about 70% of the international oil trade (Chen, Huang, and Yi, 2015), whereas WTI is used in pricing of oil imported to the Unites States (Fattouh, 2007, 2010). Despite slight quality differences, WTI and Brent are largely considered as substitutes (Worstall, 2015). Historically, both spot and futures contracts of these two benchmark products had been traded with a small price differential, with Brent slightly discounted compared to WTI due to quality difference and transportation costs. However, the WTI price became disconnected from the other benchmarks after the mid-2000 and traded at large discounts compared to Brent. This raised the question of whether an alternative benchmark that better reflects the demand and supply conditions in the oil market should be used instead of WTI (Bentzen, 2007; Fattouh, 2007).

The "break" in the WTI benchmark is attributed to the storage and pipeline capacity constraints at Cushing, Oklahoma, an oil-trade hub and the delivery location for NYMEX crude oil futures contracts, which resulted in a downward pressure on WTI price (Buyuksahin et al., 2013; Liu, Schultz, and Swieringa, 2015). Several important events that have the potential to substantially impact the prices of Brent took place after late 2010 (Buyuksahin et al., 2013). These include the Tunisian revolution in December 2010, the increased weight of Brent and decreased weight of WTI in Standard and Poor's S&P GSCI commodity index in January 2011, the Libyan crisis in February 2011, the Fukushima-Daiichi nuclear disaster in Japan in March 2011, and the inclusion of Brent in in Dow Jones' DJ-USB commodity index for the first time in January 2012. These events put upward pressure on the price of Brent (Buyuksahin et al., 2013), whereas the upsurge in U.S. oil production put downward pressure on the price of WTI, resulting in further divergence of the prices which led to a larger spread.

Although changes in the Brent-WTI spread's dynamics affect production, operations, and risk management decisions of oil-related enterprises as well as international trade, the literature on the impacts of structural breaks in the Brent-WTI spread is rather limited, and the existing results are mixed. For instance, Buyuksahin et al. (2013) find two structural breaks (November 2008 and December 2010), whereas Chen, Huang, and Yi (2015) identify only one break (December 2010). However, neither of these studies measured the real price impact of these structural breaks on each commodity's price. Given the importance of the Brent-WTI spread in global oil trade and the mixed results in the literature, the goal of this paper is to shed more light on the spread dynamics by studying an extended period of time and applying econometric methods designed to detect multiple structural breaks (Bai and Perron, 1998) and to measure the relative price impacts of these breaks (Carter and Smith, 2007).

We identified three structural breaks in the Brent-WTI spread: (1) February 25, 2005; (2) December 15, 2010; and (3) April 4, 2013. While the break in 2010 is consistent with the previous studies, the break in 2005 is three years earlier than what has been reported in Buyuksahin et al. (2013). However, the sample periods between the two studies differ substantially, which would have a nontrivial effect in estimating long-run equilibrium relationships among time-series data. Our results showed that while the price impacts of the 2005 and 2010 breaks were larger for Brent compared to WTI, the price impact of the 2010 break was about sixfold that of the 2005 break for Brent, and about thirteenfold for WTI.

2. Methodology

2.1. The relative price of a substitute (RPS) method

The RPS method proposed by Carter and Smith (2007) is based on the premise that there is a long-run equilibrium relationship between the prices of two commodities that are close substitutes in demand or supply, implying a stable relative price relationship between the commodities. Such a relative price relationship prior to an external event takes the form:

$$\ln P_{1t} - \ln P_{2t} = \mu + \beta' X_t + u_t \quad (1)$$

where u_t is a stationary random variable and X_t denotes supply and demand shifters which are only needed in the analysis if the relative price (the log price difference) is not stationary. Any deviation from this long-run relationship caused by the event will push prices to adjust back to the equilibrium, and thus the cointegrating relationship can be used to measure the price impact of the event. Specifically, a shift in the parameter μ during the event indicates a change in preferences or technology. As Carter and Smith (2007) note that a change in preferences and technology could also shift the parameter β, in which case the price impact of the event would be nonstationary and a function of X_t.

Once a structural break in the relative price is found, then the RPS method carries out model parameter estimation by a vector error correction model (VECM) using the observations until the break date. Then, the estimated model parameters are used to perform out-of-sample forecasts after the break date to compare the actual prices observed (in the case of the event) and the expected prices from the model (in the case of no event). The average forecast error during the event period shows the price impact of the event.

2.2. The Bai-Perron structural break test

Following Carter and Smith (2007), we also focus only on the case of a constant price impact, and test only for a shift in the parameter μ in equation (1) using the structural break test developed by Bai and Perron (1998). The Bai-Perron test is especially useful in the case of unknown and multiple break points. The test searches for the number and location of the breaks simultaneously. The procedure begins with testing the null hypothesis of zero breaks against one break. If the null is rejected, then the first break is taken as given and a second structural break is tested against one break, and this process continues until one cannot reject the null hypothesis of no further breaks. The Bai-Perron test implements sup-F tests; that is, the test statistic is the maximum F-statistic over all possible break points. Thus, the Bai-Perron test statistic is the maximum value of the Chow (1960) test. In addition to this sequential procedure, Bai and Perron (1998) also suggest performing a double maximum test, which aims to test of no structural break against an unknown number of breaks given some upper bound M. There are two versions of the double maximum tests based on how the weights of the M breaks are defined. While UDmax sets all weights equal to unity, WDmax defines the weights as a function of degrees of freedom and the significance level of the test. The Bai-Perron tests apply to stationary data; and hence test for a shift in parameter μ in equation (1) when there are no X_t variables. Therefore, we omit X_t in equation (1) in our structural break tests.

2.3. Rolling cointegration test

We also implement a rolling cointegration test (Swanson, 1998; Brada, Kutan, and Zhou, 2005) which identifies structural breaks by recursively employing the trace tests proposed by Johansen and Juselius (1990). Specifically, the trace test statistics are computed on a rolling time frame with a fixed and sufficiently large window length, n. For example, the first trace test statistic is computed from the first observation to the n^{th} observation. Then, it is recursively calculated from the second to the $n+1$, from the third to the $n+2$ observations, etc. The test statistics are then divided by an appropriate critical value to obtain a series of normalized trace statistics. A normalized trace statistic greater than one implies that we should reject the null hypothesis of no cointegration and conclude that the two series are cointegrated using the previous n observations.

3. Data and Preliminary Analysis

We study the relative price relationship between the two benchmark crude products using the WTI crude oil futures contracts traded at the New York Mercantile Exchange (NYMEX) ---now, known as the CME Group--- and the Brent crude oil futures contracts traded at the Intercontinental Exchange (ICE) from December 17, 1993 to April 12, 2016. Both commodities' futures contracts have expiry dates in every calendar month of the year. A single price series for each commodity is constructed by rolling over the first nearby contract on the first day of the month the trading ceases (i.e. the month preceding the contract month).

Figure 1 presents the natural logarithm of the nearby futures prices during our sample period. It can be seen from figure 1(a) that before 2005, futures prices of these two products had been traded with a small price differential, with Brent slightly discounted compared to WTI. From 2005 to 2010, the spread was nearly zero. From 2011 to 2013, however, the price relationship reversed with WTI consistently traded at a discount against Brent. In the most recent period, the spread decreased but WTI still traded at lower prices. The price relationship did not seem to change after the lift of the U.S. export ban of raw crude oil on December 19, 2015, despite the quality advantages of WTI and no barriers in WTI exports.

4. Empirical Results

4.1. Identifying structural breaks

The Bai-Perron test results presented in table 1 suggest that there are three structural breaks in the Brent-WTI spread measured as $\ln P_{Bt} - \ln P_{Wt}$, where $\ln P_{Bt}$ and $\ln P_{Wt}$ are the natural logarithms of the nearby Brent and WTI futures price, respectively, on day t. In chronological order, the first break is found in February 2005, the second is in December 2010, and the third is in April 2013. These breaks are also depicted in figure 1(b). The relatively wide 95% confidence intervals in table 1 imply that the Brent-WTI spread experienced gradual changes rather than sudden shifts, especially for the third break. In addition, both UDmax and WDmax tests reject the null of no break against the alternative of an unknown number of breaks given an upper bound of eight breaks.

For robustness check, we also conduct the Chow test by taking these three structural breaks as given. The Chow test results indicate that the break dates suggested by the Bai-Perron test are statistically significant. Furthermore, figure 1(b) shows that the average Brent-WTI spread during the four sub-periods split by the three structural breaks are -0.071, -0.002, 0.162 and 0.075, respectively. The t-tests with unequal variances suggest that these average spreads are significantly different from each other.

Results from the rolling cointegration tests obtained with n = 1,000 are presented in figure 2. The first three vertical dashed lines represent the three structural breaks found with the Bai-Perron test. Before February 2005, the trace test statistic was always greater than one, indicating that log prices of Brent and WTI have a cointegrating relationship. After including the observations following the first break, the test statistic became less than one for about a year. There were large fluctuations in the trace test statistic in the following next five years. Once the observations following the second break were included in the rolling cointegration test, the trace test statistic fell below one for a relatively long period. A similar pattern was observed again after the third break, confirming the robustness of the three structural breaks suggested by the Bai-Perron test.

4.2. Pre-break price relationship between Brent and WTI

In order to measure the price impact of these three structural breaks on WTI and Brent prices, a stable relationship between the futures prices before each break should be established. To this end, table 2 presents the results of unit root and cointegration tests for price series along with their mean and standard deviation for each sub-period separated by the three breaks. We further divide the last period of June 2014-April 2016 into two to represent the periods before and after the lift of the United States' export ban of crude oil in December 2015.

During the first two sample periods December 17, 1993-October 31, 2004 (period 1) and June 1, 2005-July 31, 2010 (period 2), both price series have unit roots and they are cointegrated, while the price spread, or the relative price, is stationary. In period 3 (February 1, 2011-October 31, 2012), however, while the WTI price series contains a unit root, the Brent futures price is stationary when an augmented Dickey-Fuller test with a drift is implemented. This results in a failure of finding an equilibrium relationship between the two prices, as the two series should be integrated of the same order to have a cointegrating relationship. Also, the spread is found to be non-stationary during this sample period. In period 4 from June 1, 2014 to April 12, 2016, while both price series contain a unit root, the spread is found to be stationary in the case with a drift. The Johansen cointegration test results also show that the two prices are not cointegrated in this sample period. It is interesting that after December 19, 2015, when the U.S. lifted the crude oil export ban, the average spread decreased from 0.081 to 0.046. However, there is still no evidence that the two price series became cointegrated. This might be due not having enough observations after the export ban lift (only 78 observations).

4.3. Estimating the price impact of breaks

Because Brent and WTI prices are found to be cointegrated only in the first two sample periods, we only focus on the first two structural breaks to estimate their price impact. We estimate an error-correction model (ECM) proposed by Engle and Granger (1987) for forecasting, given as:

$$\begin{aligned} \Delta \ln P_{Bt} &= \alpha_B z_{t-1} + \gamma_B(L)\Delta \ln P_{B,t-1} + \delta_B(L)\Delta \ln P_{W,t-1} + \varepsilon_{Bt} \\ \Delta \ln P_{Wt} &= \alpha_W z_{t-1} + \gamma_W(L)\Delta \ln P_{B,t-1} + \delta_W(L)\Delta \ln P_{W,t-1} + \varepsilon_{Wt} \end{aligned} \quad (2)$$

where $\gamma_B(L)$, $\delta_B(L)$, $\gamma_W(L)$, and $\delta_W(L)$ are polynomials in the lag operator and $z_t = \ln P_{Bt} - \ln P_{Wt} - \mu$ is the error-correction term. In other words, we constrain the cointegrating vector to be (1, -1) and add a constant in the error correction relationship. The parameters α_B and α_W measure the response of Brent and WTI prices to deviations from the long-run trend. The closer these parameters are to zero, the longer it takes for the price series to revert to their long-run trend after a shock. The half-life, a widely used measure of persistence, is defined as the number of periods required for the impulse response to a unit shock to dissipate by half. It is computed as the largest time T such that $IR(T-1) \geq 0.5$ and $IR(T-1) < 0.5$, where $IR(T)$ denotes the impulse response at time T (Steinsson, 2008).

Table 3 presents the ECM estimates from equation (2) for periods 1 and 2 (December 17, 1993-October 31, 2004 and June 1, 2005-July 31, 2010) which correspond to the two periods during which the Brent and WTI prices were cointegrated. In period 1, the estimated value of the error-correction parameter for WTI futures, α_W, is 0.033, indicating that on average the daily WTI futures price adjusts to correct 3.3% of any deviation from long-run trend. Thus, the half-life of a typical shock is 27.5 days, indicating that the two prices can deviate from their long-run relationship for about five-and-a-half weeks. The error-correction parameter for Brent, α_B, is -0.020, but it is statistically insignificant, suggesting that the primary adjustment to restore the long-run equilibrium occurs in the WTI price.

Using the ECM parameter estimates for period 1, we forecast the log prices of Brent and WTI futures and their differences from November 1, 2004 to May 31, 2005 which contains the 95% confidence interval of the first structural break as shown in table 1. Figure 3 shows the forecast errors for Brent and WTI and the log relative price (i.e. the spread). It is evident that the forecast error of the spread kept increasing very slowly during the break, despite a slight drop in the beginning of the time period. The mean forecast errors of the log prices of Brent and WTI after the break are 0.057 and 0.017, respectively. This indicates that the price impact of this structural break was much larger for Brent futures prices (5.7%) compared to WTI prices (1.7%).

In period 2, the estimated value of the error-correction parameter for WTI futures is 0.063. This value indicates that on average the daily WTI price adjusts to correct 6.3% of any deviation from the long-run trend, with a 14.5 days half-life of a typical shock. Thus, the two prices can deviate from their long-run relationship for about three weeks. This is almost half the time found in period 1, showing that the adjustment process was faster after 2004. Similar to the results for the first period, the error-correction parameter for Brent is not significantly different from zero, indicating that it the WTI price that reacts to restore the long-run equilibrium relationship between these two commodities.

Using the ECM results for period 2, we forecast the log prices of Brent and WTI futures and their differences from August 2, 2010 to May 16, 2011 which includes the 95% confidence interval of the second structural break as shown in table 1. Figure 4 shows the forecast errors for the three series. Several events took place during the time period surrounding the second break, each could affect the price of Brent and WTI futures differently (Buyuksahin et al., 2013). As figure 4 depicts, the forecast error of the Brent-WTI spread started to slightly increase in November 2010 when the weight of Brent in the S&P GSCI commodity index is announced to increase. After the start of the Tunisian revolution on December 20, 2010, the rate of the increase in forecast error enlarged significantly. This might be due to a possible combined effect with the commodity index weight adjustment becoming effective in January, 2011. The forecast error in spread reached its peak on February 25, 2011, when the Libyan crisis started. The crises resulted in a large withdrawal of sweet crude oil from the market (Buyuksahin et al., 2013), and thus both the WTI and Brent prices were significantly higher than what was expected (i.e. positive forecast errors), but the forecast error in the spread decreased from 0.168 to 0.047. Following the Fukushima-Daiichi nuclear disaster on March 14, 2011 there was instant decrease in the forecast errors for both Brent and WTI prices followed by an increase most likely due to the increasing demand for fossil fuel in Japan. However, the forecast error of the spread was stable despite the nuclear disaster, showing that its impact on the prices of Brent and WTI were similar. The mean of forecast errors of the log prices of Brent and WTI after the second structural break are 0.325 and 0.217, respectively. Similar to the first break, the price impact of the second break is larger for Brent compared to WTI. Further, the price impact of the structural break in December 2010 is substantially larger than that of the break in February 2005 for both commodities. For Brent, the price impact of the second break is 32.5% compared to 5.7% after the first break. For WTI, the price impact is 21.7% compared to 1.7%.

5. Conclusions

While Brent and WTI are still considered as the two benchmarks for light crude oil in global markets, their price relationship has changed over time, especially since mid-2000. The fluctuations in Brent-WTI spread and any deviations from its long-term trend have profound economic implications and therefore have drawn considerable attention from policy makers, media, practitioners, and academicians. However, the literature is still limited in understanding the real price impacts of deviations of the Brent-WTI spread from its equilibrium.

By using the Bai-Perron test, which allows for multiple breaks with unknown dates, we identified three structural breaks in the spread during the sample period of December, 1993 to April, 2016. While the prices of Brent and WTI were found to be cointegrated in the time periods before and after the first break in February 2005, they were not contintegrated after the second break in December 2010. Using Carter and Smith's (2007) relative price of a substitute method, we exploited the stable equilibrium price relationship found from 1993 to 2010 to measure the real price impacts of the two structural shifts in the spread. While the price impact of the 2005 structural break was 5.5% for Brent, the impact of the 2010 break was 31.7%. Similarly, the price impacts of the structural breaks in 2005 and 2010 for WTI were 1.7% and 21.7%. These results suggest that the events that led to a break in the Brent-WTI spread in 2010, affected both commodities' prices substantially more compared to the earlier break.

Because the spread became non-stationary process following the structural break in 2010, after which WTI has traded at sizeable discount against Brent, we were not able to apply the RPS method to measure the price impact of the third structural break found in April, 2013. The non-stationarity of the spread after 2010 is consistent with the findings in Cheng, Huang, and Yi (2015) and suggests a persistence change in the spread, at least until 2013. Because of the limitation of the sample period, our results are somewhat limited to examine the full impact of the lift of the U.S. crude oil export ban in December 2015. Based on the existing data we have not found significant evidence that the export ban lift had an instant impact neither on the level of the spread nor on the cointegrating relationship between Brent and WTI prices. This leaves an interesting research question which can be answered as more recent price data are observed in the future.

References

Bai, J. and P. Perron. 1998. "Estimating and testing linear models with multiple structural changes." *Econometrica* 66:47-78.

Bai, J. and P. Perron. 2003. "Computation and analysis of multiple structural change models." *Journal of Applied Econometrics* 18:1-22.

Bentzen, J.B. 2007. "Does OPEC influence crude oil prices? Testing for co-movements and causality between regional crude oil prices." *Applied Economics* 39:1375-1385.

Brada, J. C., A.M. Kutan, and S. Zhou. 2005. "Real and monetary convergence between the European Union's core and recent member countries: a rolling cointegration approach." *Journal of Banking and Finance* 29: 249-270.

Buyuksahin, B., T.K. Lee, J.T. Moser, and M.A. Robe. 2013. "Physical markets, paper markets and the WTI-Brent spread." *The Energy Journal* 34:129-151.

Carter, C. and A. Smith. 2007. "Estimating the market effect of a food scare: The case of genetically modified StarLink corn." *The Review of Economics and Statistics* 89:522-533.

Chen, W., Z. Huang, and Y. Yi. 2015. "Is there a structural change in the persistence of WTI-Brent oil price spreads in the post-2010 period?" *Economic Modelling* 50:64-71.

Chow, G.C. 1960. "Tests for equality between sets of coefficients in two linear regressions." *Econometrica* 28:591-605.

Fattouh, B. 2007. "WTI benchmark temporarily breaks down: Is it really a big deal?" *Oxford Energy Comment*, Oxford Institute for Energy Studies, April.

Fattouh, B. 2010. "The dynamics of crude oil price differentials." *Energy Economics* 32:334-342.

Johansen, S. and K. Juselius. 1990. “Maximum likelihood estimation and inference on cointegration-with applications to the demand for money.” *Oxford Bulletin of Economics and Statistics* 52:169-210.

Liu, W., E. Schultz, and J. Swieringa. 2015. “Price dynamics in global crude oil markets.” *The Journal of Futures Markets* 35:148-163.

Swanson, N.R. 1998. “Money and output viewed through a rolling window.” *Journal of Monetary Economics* 41:455-473.

Worstall, T. 2015. “The economic effects of lifting the crude export ban: WTI/Brent spread disappears.” Available at http://www.forbes.com/sites/timworstall/2015/12/28/the-economic-effects-of-lifting-the-crude-export-ban-wtibrent-spread-disappears/.

Table 1. Bai-Perron Tests for Breaks in the Brent-WTI Spread

Test	Test Statistic	5% critical value	Date of maximal F-statistic	95% confidence interval	Conclusion
Sup-F(1|0)	63.25	9.63	12/14/2010	08/19/2010-12/27/2010	At least 1 break
Sup-F(2|1)	158.36	11.14	02/25/2005	11/04/2004-05/31/2005	At least 2 breaks
Sup-F(3|2)	20.72	12.16	04/04/2013	11/01/2012-05/22/2014	At least 3 breaks
Sup-F(4|3)	11.21	12.83	07/09/1999	-	3 breaks
UDmax	148.31	10.17	-	-	# breaks ∈ {1, 2, 3, 4, 5, 6, 7, 8}
WDmax	181.93	10.91	-	-	# breaks ∈ {1, 2, 3, 4, 5, 6, 7, 8}

Notes: Sample period is 12/17/1993-04/12/2016. Maximum number of breaks is set to eight and minimum regime size is set to 5% of the sample. Robust standard errors with AR(1) pre-whitening are used for all tests (Bai and Perron, 1998).

Table 2. Pre-Break Unit Root and Cointegration Tests

		Period 1: 12/17/1993-10/31/2004				
Obs. = 2,667	Type	Test statistic	5% critical value	Conclusion	Mean	Std. dev.
Augmented Dickey-Fuller Tests						
	None	1.09	-1.95	Unit Root		
$\ln P_{Bt}$	Drift	-0.93	-2.86	Unit Root	3.05	0.32
	Drift+trend$^{\Delta}$	-2.06	-3.41	Unit Root		
	None	0.98	-1.95	Unit Root		
$\ln P_{Wt}$	Drift	-1.16	-2.86	Unit Root	3.12	0.31
	Drift+trend$^{\Delta}$	-2.30	-3.41	Unit Root		
	None	-2.27**	-1.95	Stationary		
$\ln(P_{Bt}/P_{Wt})$	Drift$^{\Delta}$	-5.52***	-2.86	Stationary	-0.07	0.036
	Drift+trend	-5.60***	-3.41	Stationary		
Johansen cointegration test						
	None	36.50***	17.95	Cointegration		
$\ln P_{Bt}$ and $\ln P_{Wt}$	Drift	37.96***	19.96	Cointegration		
	Trend	40.43***	25.32	Cointegration		
		Period 2: 06/01/2005-07/31/2010				
Obs.= 1,298	Type	Test statistic	5% critical value	Conclusion	Mean	Std. dev.
Augmented Dickey-Fuller Tests						
	None	0.33	-1.95	Unit Root		
$\ln P_{Bt}$	Drift$^{\Delta}$	-1.80	-2.86	Unit Root	4.27	0.25
	Drift+trend	-1.76	-3.41	Unit Root		
	None	0.29	-1.95	Unit Root		
$\ln P_{Wt}$	Drift$^{\Delta}$	-1.80	-2.86	Unit Root	4.27	0.26
	Drift+trend	-1.78	-3.41	Unit Root		
	None$^{\Delta}$	-4.51***	-1.95	Stationary		
$\ln(P_{Bt}/P_{Wt})$	Drift	-4.51***	-2.86	Stationary	-0.00	0.04
	Drift+trend	-4.55***	-3.41	Stationary		
Johansen cointegration test						
	None	31.65***	17.95	Cointegration		
$\ln P_{Bt}$ and $\ln P_{Wt}$	Drift	31.87***	19.96	Cointegration		
	Trend	33.49***	25.32	Cointegration		

Table 2. Cont'd.

Period 3: 02/01/2011-10/31/2012						
Obs. = 443	Type	Test statistic	5% critical value	Conclusion	Mean	Std. dev.
Augmented Dickey-Fuller Tests						
	None	0.13	-1.95	Unit Root		
$\ln P_{Bt}$	Drift$^{\Delta}$	-2.89**	-2.86	Stationary	4.72	0.06
	Drift+trend	-3.04	-3.41	Unit Root		
	None	-0.17	-1.95	Unit Root		
$\ln P_{Wt}$	Drift$^{\Delta}$	-2.30	-2.86	Unit Root	4.56	0.08
	Drift+trend	-2.51	-3.41	Unit Root		
	None	-0.16	-1.95	Unit Root		
$\ln(P_{Bt}/P_{Wt})$	Drift$^{\Delta}$	-2.27	-2.86	Unit Root	0.16	0.05
	Drift+trend	-2.34	-3.41	Unit Root		
Johansen cointegration test						
	None	14.05	17.95	Not cointegrated		
$\ln P_{Bt}$ and $\ln P_{Wt}$	Drift	14.31	19.96	Not cointegrated		
	Trend	15.42	25.32	Not cointegrated		

Period 4: 06/01/2014-04/12/2016						
Obs. = 470	Type	Test statistic	5% critical value	Conclusion	Mean	Std. dev.
Augmented Dickey-Fuller Tests						
	None	-1.68*	-1.95	Unit Root		
$\ln P_{Bt}$	Drift$^{\Delta}$	-1.48	-2.86	Unit Root	4.09	0.36
	Drift+trend	-1.52	-3.41	Unit Root		
	None	-1.66*	-1.95	Unit Root		
$\ln P_{Wt}$	Drift$^{\Delta}$	-1.55	-2.86	Unit Root	4.01	0.36
	Drift+trend	-1.45	-3.41	Unit Root		
	None	-1.43	-1.95	Unit Root		
$\ln(P_{Bt}/P_{Wt})$	Drift$^{\Delta}$	-3.05**	-2.87	Stationary	0.08	0.04
	Drift+trend	-3.07	-3.42	Unit Root		
Johansen cointegration test						
	None	12.86	17.95	Not cointegrated		
$\ln P_{Bt}$ and $\ln P_{Wt}$	Drift	15.84	19.96	Not cointegrated		
	Trend	15.26	25.32	Not cointegrated		

Table 2. Cont'd.

Period 4a: 06/01/2014-12/18/2015						
Obs. = 392	Type	Test statistic	5% critical value	Conclusion	Mean	Std. dev.
Augmented Dickey-Fuller Tests						
	None	-2.37	-1.95	Stationary		
$\ln P_{Bt}$	Drift$^{\Delta}$	-0.48	-2.86	Unit Root	4.18	0.31
	Drift+trend	-1.64	-3.41	Unit Root		
	None	-2.34	-1.95	Stationary		
$\ln P_{Wt}$	Drift$^{\Delta}$	-0.64	-2.86	Unit Root	4.10	0.32
	Drift+trend	-1.60	-3.41	Unit Root		
	None	-1.12	-1.95	Unit Root		
$\ln(P_{Bt}/P_{Wt})$	Drift$^{\Delta}$	-2.65*	-2.86	Unit Root	0.08	0.04
	Drift+trend	-2.63	-3.41	Unit Root		
Johansen cointegration test						
	None	9.95	17.95	Not cointegrated		
$\ln P_{Bt}$ and $\ln P_{Wt}$	Drift	16.31	19.96	Not cointegrated		
	Trend	13.60	25.32	Not cointegrated		

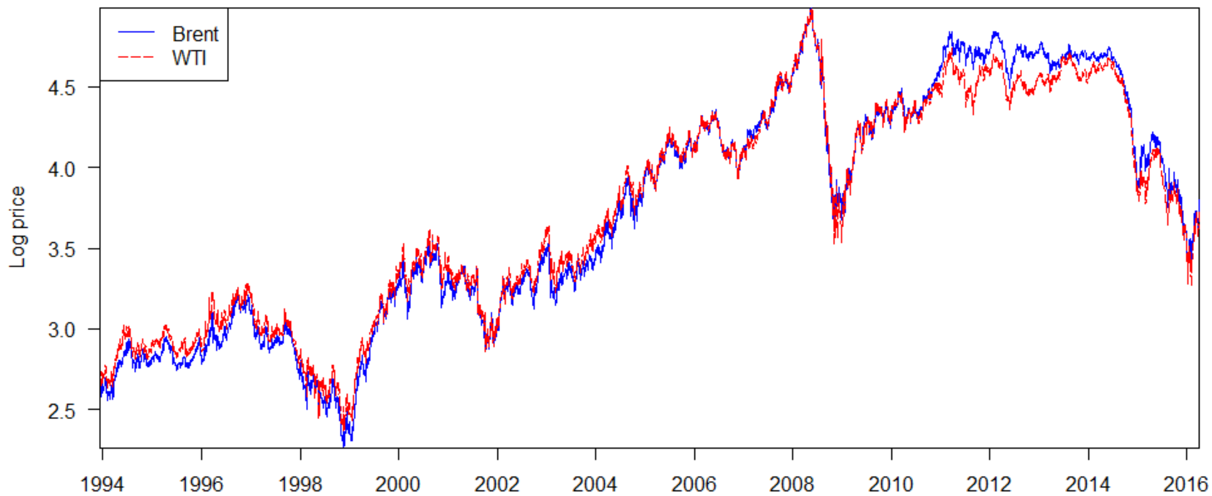
Period 4b: 12/19/2015-04/12/2016						
Obs. = 78	Type	Test statistic	5% critical value	Conclusion	Mean	Std. dev.
Augmented Dickey-Fuller Tests						
	None	0.73	-1.95	Unit Root		
$\ln P_{Bt}$	Drift$^{\Delta}$	-0.48	-2.91	Unit Root	3.58	0.11
	Drift+trend	-1.82	-3.48	Unit Root		
	None	0.41	-1.95	Unit Root		
$\ln P_{Wt}$	Drift$^{\Delta}$	-0.98	-2.91	Unit Root	3.53	0.12
	Drift+trend	-1.84	-3.48	Unit Root		
	None	-0.55	-1.95	Unit Root		
$\ln(P_{Bt}/P_{Wt})$	Drift$^{\Delta}$	-1.73	-2.91	Unit Root	0.05	0.04
	Drift+trend	-1.55	-3.48	Unit Root		
Johansen cointegration test						
	None	6.00	17.95	Not cointegrated		
$\ln P_{Bt}$ and $\ln P_{Wt}$	Drift	6.94	19.96	Not cointegrated		
	Trend	10.97	25.32	Not cointegrated		

Notes: $^{\Delta}$ indicates the most appropriate type of augmented Dickey-Fuller test selected by the significance of drift and trend terms. *, **, and *** denote statistical significance at the 10%, 5%, and 1% levels, respectively. The null hypothesis of the augmented Dickey-Fuller test is the existence of a unit root. The number of augmenting lags is determined by AIC and by the finding of no serial correlation in the residuals. For the Johansen cointegration test, only the results with the null hypothesis of no cointegrating relationship are presented. The number of lags in cointegration tests is determined by AIC.

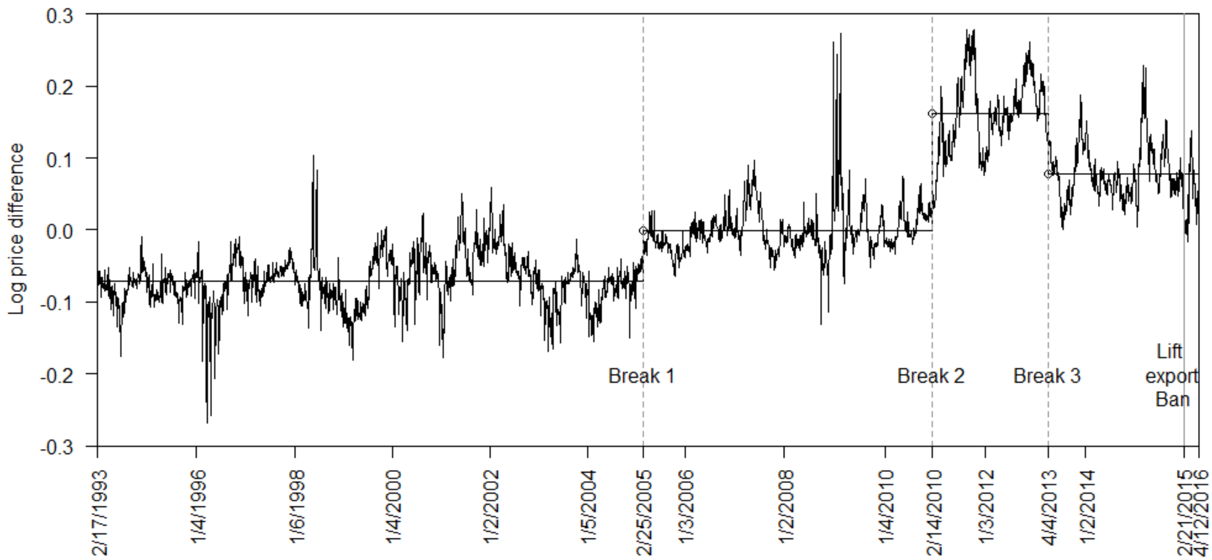
Table 3. Pre-Break Error Correction Model Results

	Period 1: 12/17/1993-10/31/2004			**Period 2: 06/01/2005-07/31/2010**		
Parameter		Brent	WTI		Brent	WTI
μ	-0.070^{***}			-0.002		
	(0.005)			(0.005)		
α		-0.020	0.033^{**}		-0.010	0.063^{***}
		(0.013)	(0.014)		(0.020)	(0.022)
γ_1		-0.085^{**}	0.034		-0.157^{***}	-0.161^{***}
		(0.035)	(0.038)		(0.055)	(0.062)
γ_2		0.019	0.067^{*}		0.052	0.041
		(0.035)	(0.038)		(0.055)	(0.062)
γ_3		0.020	0.118^{***}		-0.155^{***}	-0.110^{*}
		(0.035)	(0.037)		(0.055)	(0.062)
γ_4		-0.049	0.046			
		(0.035)	(0.037)			
γ_5		0.001	0.102^{***}			
		(0.035)	(0.037)			
δ_1		0.075^{**}	-0.011		0.099^{**}	0.096^{*}
		(0.033)	(0.036)		(0.050)	(0.056)
δ_2		-0.064^{*}	-0.123^{***}		-0.081	-0.088
		(0.033)	(0.035)		(0.050)	(0.056)
δ_3		-0.010	-0.088^{**}		0.187^{***}	0.148^{***}
		(0.033)	(0.035)		(0.050)	(0.056)
δ_4		0.028	-0.040			
		(0.033)	(0.035)			
δ_5		0.003	-0.081^{**}			
		(0.033)	(0.035)			
Diagnostics						
RMSE		0.022	0.023		0.024	0.027
Log likelihood	14198.790			6723.052		
Autocorrelation	6.065			5.373		
[p-value]	[0.194]			[0.251]		

Notes: *, **, and *** denote statistical significance at the 10%, 5%, and 1% levels, respectively. Standard errors are presented in parentheses. The number of lags is determined by AIC and also by no serial correlation in the residuals. RMSE is the root mean squared error. Autocorrelation test is conducted via the LM test for first-order serial correlation.



(a) Log prices



(b) Log price difference

Figure 1. Brent and WTI futures prices

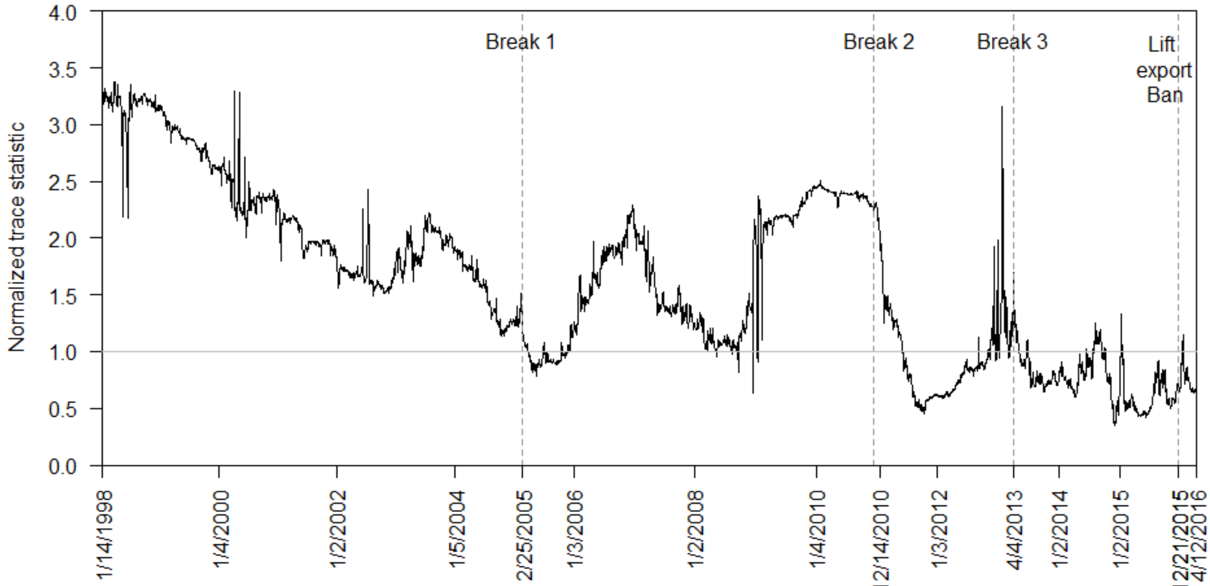


Figure 2. Rolling cointegration test

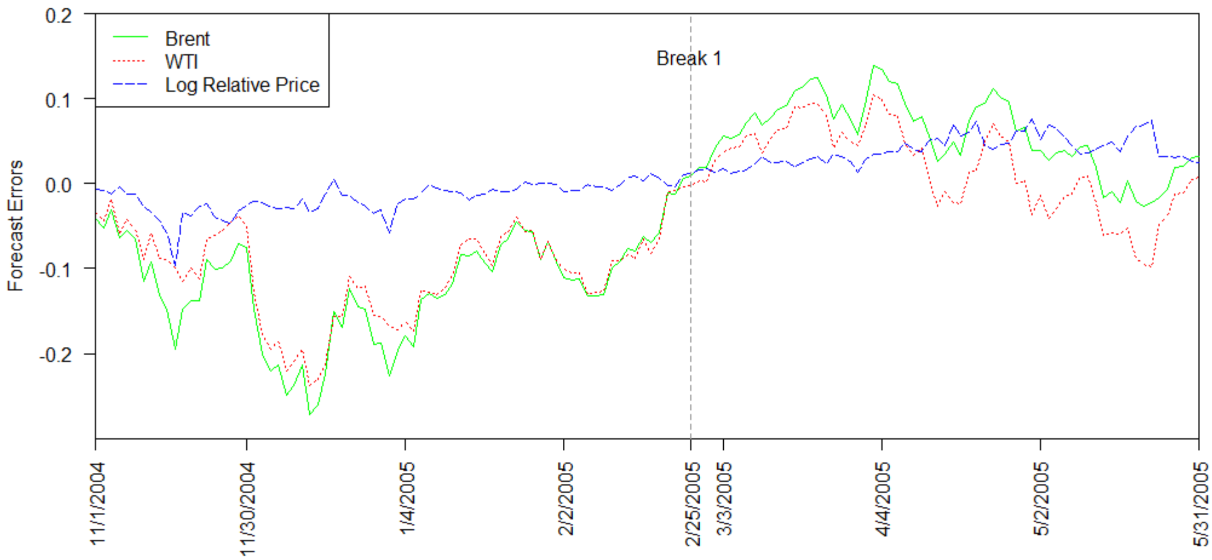


Figure 3. Forecast errors during the first break (February 25, 2005)

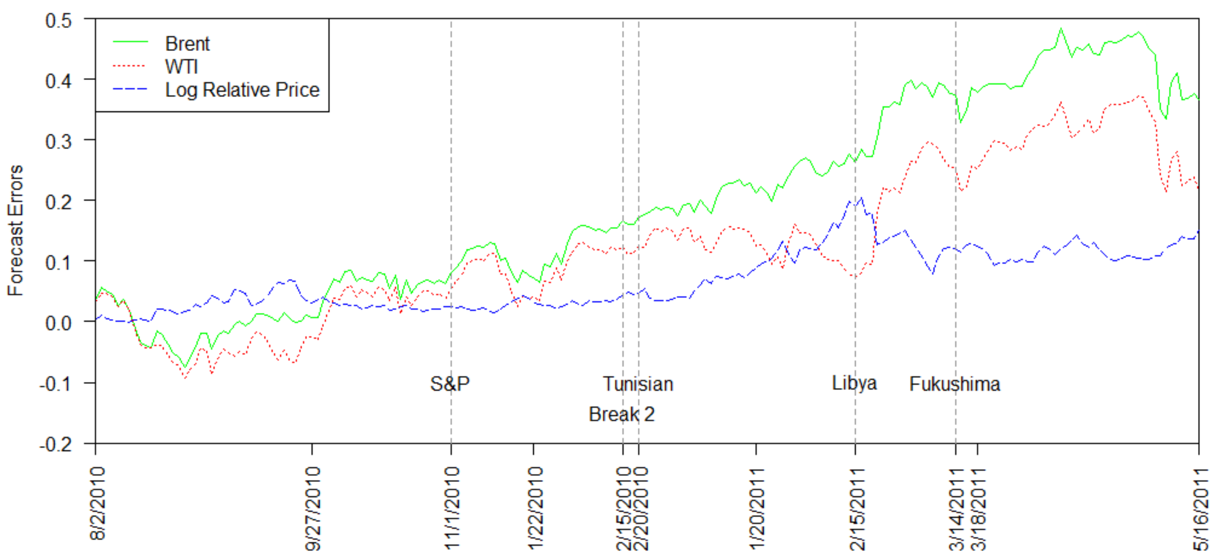


Figure 4. Forecast errors during the second break (December 15, 2010)