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A VARIANCE RATIO TEST OF RANDOM WALK IN ENERGY SPOT MARKETS

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Abstract

This study tests the random walk hypothesis in the spot prices of the petroleum products markets. Under the variance ratio test, a less restrictive random walk process namely the martingale process is examined over the period 1998-2008. The variance ratio methodology is capable of providing information regarding the linearity of multi-period variances, serial correlation as well as possible conditional heteroscedastic effect in the selected spot markets. Due to the long spanning daily data, CUSUM and Andrews tests of structural change are conducted to avoid any possible misleading statistical inferences caused by the unstable parameter in the spot markets. Our empirical findings can be summarized as follows: (1) All the energy markets reject the independent and identically distributed random walk; (2) The WTI crude oil spot prices evidence the presence of autocorrelation and conditional heteroscedastic increments; (3) The Brent crude oil and New York Harbour conventional gasoline spot prices provide strong evidence of conditional heteroscedastic increments martingale process. As a conclusion, although the energy resources returns are martingales, the heteroscedastic increments can still be used to measure the market risk and to earn a risk-adjusted abnormal return in the energy spot markets.

Keywords: spot markets, martingale process, variance ratio test.

JEL Classification: C13,C22,C53,G32,Q40.

1. Introduction

The crude oil and its product price dynamics are proven to be concomitant with the movements of global macroeconomics and financial markets. These phenomena are evidenced in the recent studies by Narayan and Smyth (2007), Cologni and Manera (2008), Gronwald (2008), Miller and Ratti (2009), Park and Ratti (2008), Nandha and Faff (2008). Knowing the importance of these energy commodities, there are ample studies investigating the underlying stochastic processes (Askari and Krichene,2008;Maslyuk and Smyth,2008;Maslyuk and Smyth, 2009; Narayan et al.,2008) in the spot prices. The deep understanding on the behaviour of spot prices provided significant contributions to model and forecast the prices movements especially to financial practitioners and energy researchers.

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Weak-form efficient market hypothesis (EMH) is one of the most frequently debatable issues in asset pricing. According to Fama (1965,1970), the essence of the EMH indicated that asset prices at any time fully reflected all available information and therefore the information is not helpful to provide any abnormal returns. In short, given only limited historical price and return, the best forecast one can make of the future price (P_{t+1}) is the current price (P_t). Since the expected price changes is zero ($E[P_{t+1} - P_t]=0$), the price changes are unpredictable given the information at time t. To be more specific, a random walk (RW) is well represented by the weak-form EMH. However, the identically and independently distributed (iid) RW is too restrictive for most of the financial markets. Although under the presence of RW_{iid}, it is still not sufficient to conclude that the market is weakly inefficient due to the possible misspecification in the RW_{iid}. Nowadays global financial markets are characterized by less restrictive non-identically and dependently (niid) RW with heteroscedastic price increments. Hence, a more general martingale process is used to denote both the RW_{iid} and RW_{niid} in this study. Once a spot price is proven to be martingale, then the return is unpredictable which strongly supports the weak-form EMH.

There are ample studies which documented the RW_{iid} test as the indicator to imply the presence of weak-form EMH in the energy markets. Serletis and rangel-Ruiz (2004) reported the daily WTI price from 1991 to 2001 as RW_{iid}. Coimbra and Esteves (2004) tested the Brent's crude oil spot and futures prices with the non-rejection of RW_{iid}. Hutchison (2004) is not able to reject the RW_{iid} in the total energy production in the Europe and UK over the quarterly data from 1960 to 1987. Based on the monthly crude oil and refined petroleum products in the United States, Kaufmann and Laskowski (2005) reported the presence of RW_{iid}. Due to the high volatile and high intensity jumps (Askari and Krichene,2008), many energy markets experienced structural changes which might alter the nature of market efficiency tests. This is because under the structural change the conventional ADF and PP tests suffered from loss of power in the statistical inferences. After the inclusion of one or multiple structural breaks in the energy markets, some researchers (Lee et al.,2006; Lee and Lee,2009;Postali and Picchetti,2006) found that their empirical studies have rejected the RW_{iid}. To the author's best knowledge, most of the energy market EMH tests under the structural change are using the weekly, monthly, quarterly and annual data. Therefore it is worth to run the daily energy spot prices EMH test under the structural break.

The purpose of this study is to examine martingale hypothesis of two crude oil and one gasoline spot prices under the consideration of structural break. A long spanning high frequency daily data of ten years (1998-2008) is used to make constructive statistical inference. A variance ratio test is conducted to analyze the variance linearity, autocorrelation and finally the martingale hypothesis. Overall, the empirical results are in favour of less restrictive RW with heteroscedastic price increments in all the studied spot markets. The strong evidence of martingale processes in the spot markets implied that market practitioners and investors are unable to make abnormal return over time. However, the heteroscedastic price increments provided predictable components that might allow market participants to earn a risk-adjusted abnormal return in the energy spot markets.

2. Data Source

According to the Energy Information Administration (http://www.eia.doe.gov), the spot price is defined as the price for a one-time open market transaction for immediate delivery of a specific quantity of product at a specific location where the commodity is purchased "on the spot" at current market rates. Therefore, the fluctuation of the spot prices provided demand and supply information of a particular energy resource across the globe. This study selected a few important energy resources such as the daily spot prices of West Texas Intermediate (WTI) Cushing[1], Brent[2] and New York Harbour[3] (NYH) conventional gasoline from the Energy Information Administration over the period 2nd January 1998 to 31st December 2008 with a total of 2713, 2759 and 2714 observations respectively.

To ensure possible spurious statistics inferences due to structural change in the energy financial markets, two structural change tests namely the Andrews (1993) test and CUSUM plot (Brown et al., 1975) are used to detect the potential break point in the selected spot markets. These tests focused on the possible structural break that is caused by the instable parameters across the studied price level (p_t) and percentage compounded returns ($r_t = \log\left(p_t / p_{t-1}\right) \times 100\%$). For Andrews test, the null hypothesis of no structural break at the x% trimmed data is based on the stability of the drift and first order autoregressive models as follows:

$$y_t = c_{I,0} + c_{I,0} y_{t-1} + \varepsilon_{I,t} \qquad t = 1,\ldots,T_1$$

$$y_t = c_{II,0} + c_{II,0} y_{t-1} + \varepsilon_{II,t} \qquad t = T_1+1,\ldots,T \qquad \ldots (1)$$

where $\varepsilon_t \sim N(0, \sigma_t^2)$, $T_1 < T$ and the y represented either the price or return series. The maximum Chow's F-statistic is defined as

$$Max \text{ F-statistic} = \max_{1 \le t \le T_i} \frac{\left(\hat{\varepsilon}'\hat{\varepsilon} - (\hat{\varepsilon}_I'\hat{\varepsilon}_I + \hat{\varepsilon}_{II}'\hat{\varepsilon}_{II})\right)/r}{(\hat{\varepsilon}_I'\hat{\varepsilon}_I + \hat{\varepsilon}_{II}'\hat{\varepsilon}_{II})/(T_{i+1} - 2r)}; \qquad \ldots (2)$$

where $\hat{\varepsilon}$ is the residual in a certain regime, r is the number of parameters and Ti denoted the break location until sub-sample T_i+1. The decision is based on the Hansen (1995) p-values. All the tests are run under 15% (x) trimmed data except for the WTI price level (x=5) where the initial CUSUM plot indicated possible break point at 12th Nov 1998.

First, the preliminary CUSUM plots for all the prices and returns series lay between the 5% significance lines which indicated the cumulative sum of residuals (e_t) are stable across the period except for the WTI price level. The CUSUM plots are illustrated in **Figure 1** and **Figure 2**. Second, the Andrews test is presented in **Table 1,** where the tests failed to reject the null hypothesis of no break point for all the spot markets including the price level and return series. It is important to note that the possible nonlinearity (second moment or above) in the returns is not taken into account under this test.

Table 1. Andrews Structural Break Test

Spot market	*Price series*		*Return series*	
	Max-F statistic	*Possible break point*	*Max-F statistic*	*Possible break point*
WTI Cude oil	9.0377 (0.2079)	12th Nov 1998	6.1709 (0.3897)	25th Mar 2003
Brent Crude oil	3.1646 (0.8641)	25th Aug 2003	3.0695 (0.8782)	15th Nov 2001
NYH Gasoline	4.2323 (0.6870)	17th Dec 2004	8.3958 (0.1769)	11th Jan 2007

Note: * and ** denoted 5% and 1% significance levels.

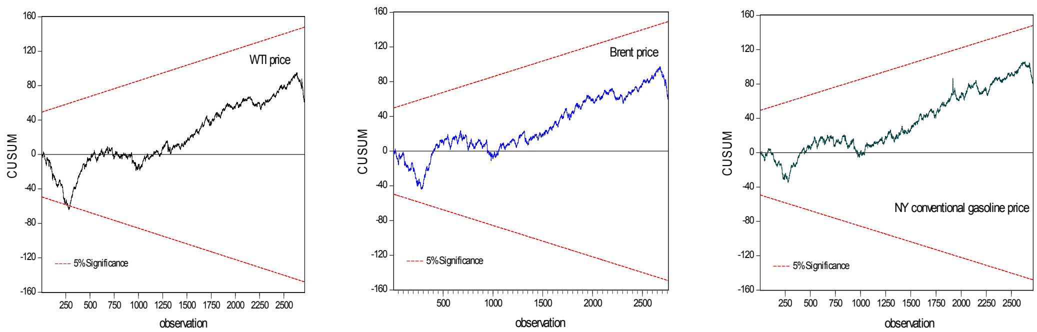

Figure 1. CUSUM Tests for Spot Market Prices

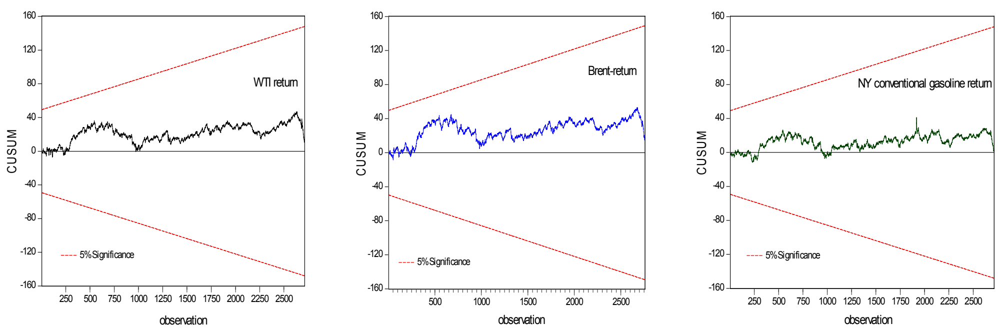

Figure 2. CUSUM Tests for Spot Market Returns

3. Methodology

In this study, two types of random walks are considered in the variance ratio test. The first type of random walk (RW_{iid}) refers to the most restrictive independent and identically distributed (iid) increments with the following specification:

Price: $p_t = \omega + p_{t-1} + \varepsilon_t$

Price changes: $r_t = p_t - p_{t-1} = \omega + \varepsilon_t$ … (3)

where p_t and ω represented the natural logarithm of spot price and possible arbitrary drift respectively. The increments ε_t are iid Gaussian random variables with mean zero and variance σ_0^2. In other words, the ε_t is restricted with no serial correlation in any form (including linear or nonlinear) among the lead and lag increments.

However, the energy spot markets often indicated non-gaussian and possible correlation volatility which deviated from the RW_{iid}. In order to cater to these empirical stylized facts, a less restrictive random walk by relaxing both the independent and identical property in the RW_{iid} can be formed. This general version of non-iid random walk (RW_{niid}) or sometimes referred as martingale allowed the possible correlation in the non-linear form of increments where the autocovariances $\gamma(\varepsilon_t^2, \varepsilon_{t-h}^2) \neq 0$ even though the linear form $\gamma(\varepsilon_t, \varepsilon_{t-h}) = 0$ for all the $h \neq 0$. In short, the RW_{iid} provides the homoscedastic increments whereas the RW_{niid} is endowed with heteroscedastic increments.

3.1 Variance Ratio Test

The core of this test relied on the linearity property of the increments of RW_{iid}. Although the RW_{niid} encompassed heteroscedastic increments, the variance of sum has to be the same as the sum of variance in the variance ratio test. In short, under the RW_{iid}, the variance of k-price changes ($r_t + r_{t-1} + ... + r_{t-k+1}$) must equal the sum of *k* time the variance of r_t (**Appendix A**).

3.1.1 RW_{iid} test

Let a sample size of nk+1 observations (p_0, p_1, …., p_{nk}) at equally spaced intervals, where k is any integer greater than 1 and *nk* is the number of observations of p_t. The unbiased estimators (Lo and MacKinlay,1988) using overlapping k^{th} differences of p_t can be expressed as follows:

$$\hat{\sigma}^2_{(1)} = \frac{1}{nk-1}\sum_{t=1}^{nk}\left(p_t - p_{t-1} - \hat{\mu}\right)^2 \quad \dots (4)$$

$$\hat{\sigma}^2_{(k)} = \frac{n-1}{kn(nk-k+1)}\sum_{t=k}^{nk}\left(p_t - p_{t-k} - k\hat{\mu}\right)^2 \quad \dots (5)$$

where $\hat{\mu} = \frac{1}{nk}\sum_{t=1}^{nk} r_t$ and $\hat{\sigma}^2_{(1)}$ is the unbiased sample variance. Under the normal limiting distribution (Stuart and Ord,1987) and further expression in autocorrelation function (**Appendix B**), the RW_{iid} estimated variance ratio $\hat{VR}(k)$ is

$$\sqrt{nk}\left(\frac{\hat{\sigma}^2_{(k)}}{\hat{\sigma}^2_{(1)}} - 1\right) = \sqrt{nk}\left(\hat{VR}(k) - 1\right) \sim N\left(0, \frac{2(2k-1)(k-1)}{3k}\right). \quad \dots (6)$$

If the null hypothesis of homoskedastic increments of RW_{iid} is not rejected, the associated test statistic has an asymptotic standard normal distribution as follows:

$$Z(k) = \frac{\hat{VR}(k) - 1}{\hat{\sigma}_0(k)} \sim N(0,1) \qquad \dots (7)$$

where $\hat{\sigma}_0(k) = \left[\frac{2(2k-1)(k-1)}{3k(nk)}\right]^{1/2}$ is the asymptotic variance for $\hat{VR}(k) - 1$.

3.1.2 RW_{niid} test

The energy spot market that rejected the RWiid normally has the VR(*k*) either statistically more or less than unity. Under this condition, the *k*-period returns variance is

$$VR^*(k) = 1 + \frac{2}{k}\sum_{h=1}^{k-1}(k-h)\rho_{(h)} \qquad \dots (8)$$

with any nonzero $\rho_{(h)}$ and the (k-h) is the multiplier for autocorrelation lag-*i* and decreased as the lag increased. Under the four conditions introduced by Lo and MacKinlay (1988) and Campbell et al. (1997), the RW_{niid} hold uncorrelated increments but allowed for heteroscedastic increments in the deterministic variance and time-varying variance (Engle,1982). The null hypothesis H_0^* for RW_{niid} however still remained as unity, as long as the returns are uncorrelated as the number of observations increased without bound.

To perform the test, one needed to obtain the asymptotic standard error $\hat{\sigma}_0^*(k)$ from the autocorrelation form of $VR^*(k)$ in **equation 8** as follows:

$$\hat{\sigma}_0^*(k) = \left[\frac{4}{nk}\sum_{t=1}^{k-1}\left(1 - \frac{h}{k}\right)^2 \hat{\varphi}_h\right]^{1/2} \qquad \dots (9)$$

$$\text{where } \hat{\varphi}_h = \frac{nk\sum_{t=h+1}^{nk}(p_t - p_{t-1} - \hat{\mu})^2(p_{t-h} - p_{t-h-1} - \hat{\mu})^2}{\left[\sum_{t=1}^{nk}(p_t - p_{t-1} - \hat{\mu})^2\right]}. \qquad \dots (10)$$

In a similar way, the standardized heteroscedastic RW_{niid} test can be obtained as

$$Z^*(k) = \frac{\hat{VR}^*(k) - 1}{\hat{\sigma}_0^*(k)} \sim N(0,1). \qquad \dots (11)$$

4. Empirical Result

A quick statistical overview is reported at **Table 2**. First, the normality test indicated strong deviation from the skewness and kurtosis from a normal distribution (skewness=0 and kurtosis=3) for all the spot prices. Therefore, the Jacque-Bera[4] statistics rejected the hypothesis of normal distribution at the 1% significance level. In addition, the **Figure 3** shows similar results in the quantile-quantile plots with heavier tails than a normal distribution.

Second, the serial correlation Ljung-Box5 Q(12) statistics indicated strong rejection at 1% significance level for the price level. However, after the first differencing, the price changes (return) shown only rejection at 5% level for the WTI and failed to reject for both the Brent and NYH. Third, the RWiid unit root tests failed to reject the presence of unit root at 5% level for the WTI and Brent under the ADF and PP6 tests. However, the NYH indicated contrary result with no unit root.

Table 2. Preliminary Statistical Analysis

Spot market	*Normality test (return)*			*Serial correlation test*		*RWiid - Unit root test (price)*		
	Skewness	*Kurtosis*	*JB statistic*	*LB Q(12) price*	*LB Q(12) return*	*ADF test*	*Lag*	*PP test*
WTI	-0.3040	7.6137	2448.06 **	32108** (0.000)	23.454 * (0.024)	-2.8550 (0.1777)	2	-2.9160 (0.1574)
Brent	-0.1999	6.9807	1839.39 **	32648** (0.000)	6.8858 (0.865)	-2.7281 (0.2251)	0	-2.9136 (0.1582)
NYB	0.0284	6.5264	1406.61 **	31923** (0.000)	18.037 (0.115)	-3.4722 * (0.0426)	3	-3.4461 * (0.0456)

Note: The superscript a, b, c denoted the 10%, 5% and 1% significance level respectively.

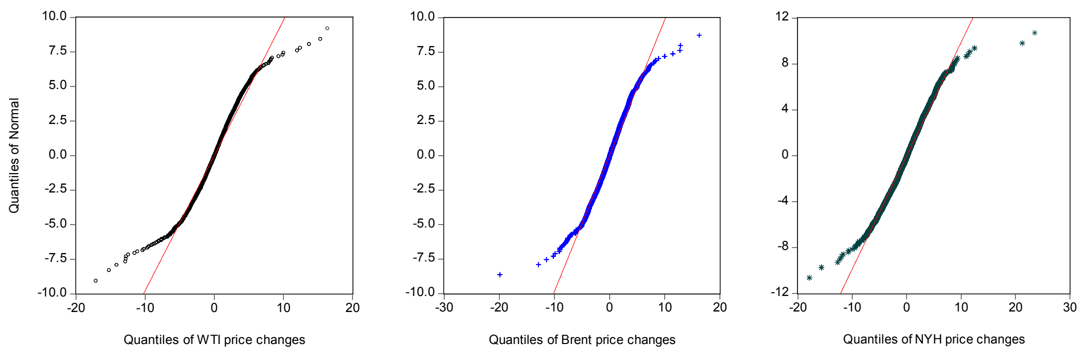

Figure 3. Normality Tests for Spot Market Price Changes (Returns)

As a result, the preliminary analysis summarized that all the return series deviated from the normal distribution and indicated strong correlation at the price level however weakly serial correlation after the first differencing. For RW tests, a more comprehensive variance ratio test will be conducted to conclude the existence of RW.

4.1 Variance Ratio Test

The variance ratio (VR) provided some intuitive statistical information regarding the studied series. For example, when VR(k) < 1, the negative autocorrelations in the return series demonstrated a mean-reverting process, while the positively correlated (VR(k)>1) return exhibited trends. In short, this test is most likely to reject the RW_{iid} when the VR deviated significantly from unity.

The rejection of homoskedasatic RW_{iid} is based on whether the absolute $Z(q)$ is greater than the one-sided 10%, 5% and 1% standard normal distribution critical values (1.64, 1.96 and 2.57). This rejection confirmed the presence of either heteroskedastic or/and serial correlation in the spot markets. In the following test of heteroskedastic RW_{niid}, the rejection of heteroskedastic random walk led to the existence of autocorrelation in the spot price series.

Table 3. Two Sided Variance Ratio Tests

WTI crude oil	*Variance ratio*	*RWiid*	*RWniid*	*autocorrelation*
k	VR(k)	Z(k)	Z*(k)	$\hat{\rho}(1) = VR(2) - 1$
2	0.9744	-0.9425	-0.8844	-0.0256
4	0.8994	-3.7018 c	-1.9213 a	
8	0.8564	-5.2855 c	-1.8323 a	
16	0.7827	-8.0014 c	-1.9969 b	
32	0.7323	-9.8577 c	-1.7998 a	
Brent crude oil		*RWiid*	*RWniid*	*autocorrelation*
k	VR(k)	Z(k)	Z*(k)	$\hat{\rho}(1) = VR(2) - 1$
2	1.0180	0.6695	0.8145	0.0180
4	1.0343	1.2752	0.8217	
8	1.0511	1.8987 a	0.7478	
16	1.0951	3.5337 c	0.9557	
32	1.0373	1.3866	0.2698	
NYH conventional gasoil		*RWiid*	*RWniid*	*autocorrelation*
k	VR(k)	Z(k)	Z*(k)	$\hat{\rho}(1) = VR(2) - 1$
2	1.0404	1.4882	1.1789	0.0404
4	1.0333	1.2274	0.5568	
8	0.9681	-1.1725	-0.3636	
16	0.9397	-2.2201 b	-0.5155	
32	0.9361	-2.3517 b	-0.4183	

Note: [a], [b] and [c] denoted the 10%, 5% and 1% level of significance.

Table 3 reported the variance ratio test for various *k* consecutive returns for all the spot markets. First, the VR(*k*) for WTI crude oil indicated the tendency of deviation from unity (0.9744 to 0.7323) as the *k*-period returns increased from k=2 to k=32 for the homoscedastic RW_{iid} test. This happened when the linear function of autocorrelation $\rho_{(h)}$ is not all zero in the VR (**equation 8**). The stronger deviations from unity indicated more significant autocorrelated items in the autocorrelation function. All the VR(*k*) are less than unity implying that the WTI has mean reversion in prices. In other words, the prices are negatively correlated, for example, the $\rho_{(1)}$ is -0.0256 in **Table 3**. Based on the normal critical value, only the VR(2) failed to reject the homoscedastic RW_{iid} while the higher VRs rejected the null hypothesis at 1% significance level. Therefore, the variances of the WTI increments are not linear in all the sampling intervals. Due to this, we further conducted the heteroscedastic RW_{niid} and found that only the VR(2) failed to reject the heteroscedastic RW_{niid}. However for the higher VR(k) (*k*=4,8,16,32), they are rejected at 5 or 10% significance levels which implied the autocorrelation dominated price series.

Second, the Brent and NYH prices indicated less deviation from unity compared to WTI across the VR(k) series. All the Brent VRs exceeded unity which implied permanent trend in price series. However, the NYH series indicated mixture of positive correlation for 2-,4- and 8-period consecutive returns series and negatively correlated for 16- and 32-period series. Under the VR test, the Brent series rejected the homoscedastic RW_{iid} for VR(8) and VR(16) whereas the NYB at VR(16) and VR(32) respectively. For heteroscedastic RW_{niid} test, both the series failed to reject the presence of heteroscedastic RW_{niid} at 1% significance level.

5. Conclusion and Discussion

This paper examined whether three selected energy spot market returns followed a random walk process. A long spanning data of ten years (1998-2008) have been used for this analysis. As a summary, there are two important findings that might attract the interest of investors and energy researches.

First, the structural change is not identified in the long spanning (1998-2008) price and return series under the CUSUM plots and Andrews tests. This finding is somewhat contrary to most of the lower frequencies (weekly, quarterly, annually) data. These included weekly WTI and Brent crude oil (Maslyuk and Smyth,2008) with two structural breaks (both in 1999 and 2001) over the period 1991-2004; monthly US real oil prices over the period 1947-1996 (Sadorsky,1999) with a structural break; annual data from 1861 to 1999 (Postali and Picchetti 2006) documented two endogenous structural break in the international oil prices. The long period interval data (*eg*. monthly data) normally will smooth out the possible daily fluctuations and incurred possible instability in the model parameter estimations. Although the detections of structural breaks are most likely to happen in a long period interval the selection of data frequency (daily, weekly, month or annually data) may result in somewhat different results in the oil product spot markets.

Second, the variance ratio tests provided not only the RW_{iid} (like ADF and PP unit root) information but also the variance linearity, serial correlation as well as possible heteroscedastic effect in the oil product spot markets. This information has strong implications in the weak-form market efficiency hypothesis where the stock prices at any time fully reflected all available market information. The prices for oil products rejected the homoscedastic RW_{iid} processes under various periods of variance ratio tests. On the other hand, it is found that the spot prices are in favour of heteroscedastic or autorcorrelation increments under the heteroscedastic RW_{niid} tests. In other words, overall there are evidences that all the spot prices followed a less restrictive RW, the heteroscedastic RW_{niid} (martingale process). As a conclusion, all the prices or returns in the three spot energy markets are unpredictable and therefore the market participants are unable to make abnormal returns in future. However, the conditional heteroscedastic increments provided predictability components in the conditional variances which can be used to deal with risk management and portfolio analysis in the selected energy spot markets.

Footnotes:

1. A crude stream is produced in Texas and southern Oklahoma which served as a reference or "marker" for pricing a number of other crude streams and which is traded in the domestic spot market at Cushing, Oklahoma (Unit: USD per barrel). Crude oil is refined to produce a wide array of petroleum products, including heating oils; gasoline, diesel and jet fuels; lubricants; asphalt; ethane, propane, and butane; and many other products used for their energy or chemical content.
2. A blended crude stream is produced in the North Sea region which served as a reference or "marker" for pricing a number of other crude streams (Unit: USD per barrel).
3. The location is specified in either spot or futures contracts for delivery of a product in New York Harbor. Finished motor gasoline is not included in the oxygenated or reformulated gasoline categories; excluded reformulated gasoline blendstock for oxygenate blending (RBOB) as well as other blendstock.

4. The Jarque-Bera statistic measured the deviations of skewness and kurtosis from the normal distribution:

$$JB = \frac{T}{6}\left[\text{Skewness}^2 + \left(\frac{\text{Kurtosis} - 3}{2} \right)^2 \right]; \quad \ldots (12)$$

5. The Ljung-Box *Q*-statistic (*h*) indicated there is no autocorrelation up to lag *h*:

$$LB - Q = T(T+2)\sum_{i=1}^{h} \frac{\rho_{(i)}^2}{T - i}; \quad \ldots (13)$$

6. Both the augmented Dickey-Fuller (ADF) and Philips-Perron (PP) one sided unit root tests with exogenous regressors which consisted of a constant and trend under the null hypothesis of the presence of unit root. The decision is based on the MacKinnon (1996) 5% critical value t_α ($t_{0.05\%,WTI}$ = –3.41146, $t_{0.05\%,Brent}$ = –3.41143 and $t_{0.05\%,NYH}$ = –3.41146 respectively);

References

Andrews D.W. K. (1993), “Tests for Parameter Instability and Structural Change With Unknown Change Point”, *Econometrica,* 61(4); 821–856.

Askari H., Krichene, N. (2008), “Oil price dynamics (2002-2006)”, *Energy Economics,* 30; 2134-2153.

Brown R. L., Durbin J. and Evans J. M. (1975), “Techniques for Testing the Constancy of Regression Relationships Over Time”, *Journal of the Royal Statistical Society, Series B*, 37; 149–192.

Campbell J.Y., Lo A.W., MacKinlay A.C. (1997), *The econometrics of financial markets*, Princeton University Press: New Jersey.

Coimbra C., Esteves P.S. (2004), “Oil price assumptions in macroeconomic forecasts: should we follow futures market expectations?”, *OPEC Review*, 28; 87–106.

Cologni A., Manera M. (2008), “Oil prices, inflation and interest rates in a structural cointegrated VAR model for the G-7 countries”, *Energy Economics*, 30(3); 856–888.

Engle R. (1982), “Autoregressive Conditional Heteroscedasticity with Estimates of the Variance of United Kingdom Inflation”, *Econometrica,* 50; 987–1007.

Fama E. (1965), “Random Walks in Stock Market Prices”, *Financial Analysts Journal*, 21(5); 55–59.

Fama E. (1970), “Efficient Capital Markets: A Review of Theory and Empirical Work”, *The Journal of Finance*, 25(2); 383–417.

Gronwald M. (2008), “Large oil shocks and the US economy: infrequent incidents with large effects”, *Energy Journal*, 29-1; 151–171.

Hutchison M.M. (2004), “Manufacturing sector resiliency to energy booms: empirical evidence from Norway, the Netherlands, and the United Kingdom”, *Oxford Economic Papers*, 46; 311–329.

Kaufmann R.K., Laskowski C. (2005), “Causes for an asymmetric relation between the price of crude oil and refined petroleum products”, *Energy Policy*, 33; 1587–96.

Lee C.C., Lee J.D. (2009), “Energy prices, multiple structural breaks, and efficient market hypothesis”, *Applied Energy*, 86; 466-479.

Lee J., List J.A., Strazicich M.C. (2006), “Non-renewable resource prices: deterministic or stochastic trends?”, *Journal of Environmental Economics and Management*, 51; 354–370.

Lo A.W., MacKinlay A.C. (1988), “Stock market prices do not follow random walks: evidence from a simple specification test”, *Review of Financial studies*, 1; 41-66.

Maslyuk S., Smyth R. (2008), “Unit root properties of crude oil spot and futures prices”, *Energy Policy*, 36; 2591-2600.

Maslyuk S., Smyth R. (2009), “Nonlinear Unit root properties of crude oil spot and futures prices”, *Energy Policy*, 36; 2591-2600.

Miller J.I., Ratti R.A. (2009), “Crude Oil and Stock Markets: Stability, Instability, and Bubbles”, *Energy Economics*, doi:10.1016/j.eneco.2009.01.009.

Nandha M. Faff R. (2008), “Does oil move equity prices? A global view”, *Energy Economics*, 30; 986-997.

Narayan P.K., Narayan S., Smyth R. (2008), “Are oil shocks permanent or temporary? Panel data evidence from crude oil and NGL production in 60 countries”, *Energy Economics*, 30(3); 919–936.

Narayan P.K., Smyth R. (2007), “A panel cointegration analysis of the demand for oil in the Middle East”, *Energy Policy*, 35(12); 6258-6265.

Park J., Ratti R.A. (2008), “Oil price shocks and stock markets in the U.S. and 13 European countries”, *Energy Economics*, 30; 2587-2608.

Postali F., Picchetti P. (2006), “Geometric Brownian Motion and structural breaks in oil prices: a quantitative analysis”, *Energy Economics*, 28; 506–522.

Sadorsky P. (1999), “Oil price shocks and stock market activity”, *Energy Economics*, 5; 449–469.

Serletis A., Rangel-Ruiz R. (2004), “Testing for common features in North American energy markets”, *Energy Economics*, 26; 401–414.

Stuart A., Ord K. (1993), *Kendall’s advanced theory of statistics, Vols. I-III*, Oxford University Press: New York.

Appendix A

Some intuitive variance ratio statistics under a general population values can be obtained under the linearity property. The variance for *k*-returns can be derived using the induction method as follow:

For *i*=2,

$$V(r_t + r_{t+1}) = V(r_t) + V(r_{t+1}) + 2\gamma(r_t, r_{t+1})$$

$$= 2V(r_t) + 2\gamma_{(1)}; \qquad \text{… (A1)}$$

For *i*=3,

$$V\left((r_t + r_{t+1}) + r_{t+2}\right) = 2V(r_t) + 2\gamma_{(1)} + V(r_{t+2}) + 2\gamma(r_t + r_{t+1}, r_{t+2})$$

$$= 3V(r_t) + 2\gamma_{(1)} + 2\gamma(r_t, r_{t+2}) + 2\gamma(r_{t+1}, r_{t+2})$$

$$= 3V(r_t) + 2\left(2\gamma_{(1)} + \gamma_{(2)}\right); \qquad \text{… (A2)}$$

For *i*=4,

$$V\left((r_t + r_{t+1} + r_{t+2}) + r_{t+3}\right) = 3V(r_t) + 4\gamma_{(1)} + 2\gamma_{(2)} + V(r_{t+3}) + 2\gamma(r_t + r_{t+1} + r_{t+2}, r_{t+3})$$

$$= 4V(r_t) + 2\left(3\gamma_{(1)} + 2\gamma_{(2)} + \gamma_{(3)}\right). \qquad \text{… (A3)}$$

Similarly, for i=*k*-period returns,

$$V\left(r_t + \dots + r_{t+k-1}\right) = kV(r_t) + 2\left((k-1)\gamma_{(1)} + \dots + \gamma_{(k-1)}\right)$$

$$= kV(r_t) + 2\sum_{i=1}^{k-1}(k-i)\gamma_{(i)}. \qquad \text{… (A4)}$$

Under the RW_{iid}, the variance ratio (VR) *k*-period returns for any integer $k \geq 2$ is

$$VR(k) = \frac{V\left(r_t + \dots + r_{t+k-1}\right)}{kV(r_t)}.$$

$$= 1 + \frac{2}{k}\sum_{i=1}^{k-1}(k-i)\rho_{(i)}$$

$$=1 \qquad \text{… (A5)}$$

where all the autocorrelations $\rho_{(i)} = 0$. Therefore the null hypothesis of the RW_{iid} assumed that the $\varepsilon_t \overset{iid}{\sim} N(0, \sigma_0^2)$.

Appendix B

Using overlapping k^{th} differences of p_t , the autocorrelation form of variance can be expressed as follows:

$$\hat{\sigma}^2_{(k)}$$

$$= \frac{1}{nk^2}\sum_{t=k}^{nk}\left(p_t - p_{t-k} - k\hat{\mu}\right)^2$$

$$= \frac{1}{nk^2}\sum_{t=k}^{nk}\left(\sum_{w=1}^{k}\hat{\varepsilon}_{t-w+1}\right)^2$$

$$= \frac{1}{nk^2}\sum_{t=k}^{nk}\left(\sum_{w=1}^{k}\hat{\varepsilon}^2_{t-w+1} + 2\sum_{w=1}^{k-1}\hat{\varepsilon}_{t-w+1}\hat{\varepsilon}_{t-w} + \ldots + 2\hat{\varepsilon}_t\hat{\varepsilon}_{t-k+1}\right)$$

$$= \frac{1}{nk^2}\left[k\sum_{t=1}^{nk}\hat{\varepsilon}^2_t + 2(k-1)\sum_{t=2}^{nk}\hat{\varepsilon}_t\hat{\varepsilon}_{t-1} + 2(k-2)\sum_{t=2}^{nk}\hat{\varepsilon}_t\hat{\varepsilon}_{t-2} + \ldots + 2\sum_{t=2}^{nk}\hat{\varepsilon}_t\hat{\varepsilon}_{t-k+1} + \text{other terms}\right]$$

$$= \hat{\gamma}_{(0)} + \sum_{h=1}^{k-1}\frac{2}{k}(k-h)\hat{\gamma}_{(h)} + \text{other terms} \qquad \ldots \text{(B1)}$$

where the 'other terms' represented a quantity that is of an order smaller than $1/\sqrt{n}$ in probability. Under the standard limit theorem, both the estimators $\hat{\sigma}^2_{(1)}$ and $\hat{\sigma}^2_{(k)}$ (estimator for the unknown population $\hat{\sigma}^2_0$) followed the Gaussian limiting distribution (Lo and MacKinlay,1988):

$$\sqrt{nk}\left(\hat{\sigma}^2_{(1)} - \hat{\sigma}^2_0\right) \sim N\left(0, 2\hat{\sigma}^4_0\right)$$

$$\sqrt{nk}\left(\hat{\sigma}^2_{(k)} - \hat{\sigma}^2_0\right) \sim N\left(0, 2\hat{\sigma}^4_0\left[\frac{2k^2+1}{3k}\right]\right) \qquad \ldots \text{(B2)}$$

According to Hausman's (1978):

$$\sqrt{nk}\left(\frac{\hat{\sigma}^2_{(k)}}{\hat{\sigma}^2_{(1)}} - 1\right) \sim N\left(0, 2\left[\frac{2k^2+1}{3k} - 1\right]\right) \sim N\left(0, \frac{2(2k-1)(k-1)}{3k}\right) \qquad \ldots \text{(B3)}$$

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