The empirical relationship between energy futures prices and exchange rates

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Abstract

This paper investigates the interaction between energy futures prices and exchange rates. Results are presented to show that futures prices for crude oil, heating oil and unleaded gasoline are co-integrated with a trade-weighted index of exchange rates. This is important because it means that there exists a long-run equilibrium relationship between these four variables. Granger causality results for both the long- and short-run are presented. Evidence is also presented that suggests exchange rates transmit exogenous shocks to energy futures prices. © 2000 Elsevier Science B.V. All rights reserved.

1. Introduction

Recently, energy commodity prices have been falling at the same time that the US dollar has been rising. The rising US dollar is in response to increased demand for US currency due in part to several global economic financial crises. Consequently, the relationship between energy commodity price movements and exchange rates is an important and interesting topic to study.

Bloomberg and Harris (1995) have offered some insight into the relationship between commodity prices and exchange rates. In particular, they offered several

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explanations as to why commodity prices began rising in the early 1990s. First, they suggested that commodity prices had rebounded from unusually depressed levels. This rebound in commodity prices may have represented a catching up process as commodity prices return to a more normal pricing situation. Second, they suggested that commodity prices may also have risen in response to the weak dollar. Since commodities are homogeneous and traded internationally, it is most likely that they are subject to the law of one price. This means that commodities have similar prices in each country's home currency. Thus as the US dollar weakens relative to other currencies, ceteris paribus, commodity consumers outside the United States should be willing to pay more dollars for commodity inputs. Consequently, exchange rate movements may be an important stimulus for commodity price changes. Bloomberg and Harris (1995) conduct their analysis by comparing a trade-weighted index of the US dollar with various commodity price indices (like the CRB/Bridge commodity price index). They find that the correlation between exchange rates and commodity price indices increased after 1986. For example, Bloomberg and Harris (1995) found that the correlation between the CRB/Bridge commodity price index and a trade-weighted index of the US dollar was −0.19 over the period 1970−1986 and −0.34 over the period 1987−1994. The correlation between the Journal of Commerce (JOC) commodity price index and a trade-weighted index of the US dollar was −0.02 over the period 1970−1986 and −0.37 over the period 1987−1994.

Of related interest is the paper by Pindyck and Rotemberg (1990) who investigated the co-movements of prices between various unrelated commodities. Pindyck and Rotemberg (1990) offer numerous statistical tests which confirm that the prices of several unrelated commodities, such as wheat, copper, cotton, gold, crude oil, lumber, and cocoa have a persistent tendency to move together. They suggest that one possible explanation for this '...is that commodity price movements are to some extent the result of herd behaviour'. Particularly interesting for my analysis is that Pindyck and Rotemberg (1990) also find that an equally weighted index of the dollar value of British pounds, German marks, and Japanese Yen negatively and significantly impacts the price of crude oil in both ordinary least squares regressions and latent variable models.[1]

In this paper I further investigate the interaction between energy futures prices and exchange rates. Previous work by Serlitis (1994) has shown that futures prices for crude oil, heating oil and unleaded gasoline are co-integrated.[2] In this paper I use Johansen's (Johansen, 1988, 1991) econometric technique to show that futures prices for crude oil, heating oil and unleaded gasoline are co-integrated with a

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[1] The relationship between oil prices and exchange rate movements has also been discussed by Golub (1983), Krugman (1983a,b), and Zhou (1995).

[2] Since the prices in different countries are also affected by transport costs, and transport costs are in turn affected by oil prices, there may be a special reason for prices of oil products to move together. Short-term overcapacity in transport and other short-term factors may upset the long-run relationship. I thank an anonymous reviewer for pointing this out.

trade-weighted index of exchange rates. This is important because it means that there exists a long-run equilibrium relationship between these four variables. A vector error correction model (VECM) is built to investigate Granger causal relationships. Results from the VECM indicate that movements in exchange rates precede movements in heating oil futures prices in both the short- and the long-run while movements in exchange rates precede movements in crude oil futures prices in the short-run. Stability of the VECM is investigated by (1) using recursive estimation of the error correction terms; and (2) testing for co-integration in models with regime shifts.

This paper is organized as follows. Section 2 presents the econometric methodology used. Section 3 reports standard unit root tests for testing the null hypothesis of a unit root against the alternative hypothesis of stationarity. Section 3 also reports results from testing the null hypothesis of no co-integration. Section 4 presents Granger causality results for both the long-run and the short-run. Section 5 discusses the stability of the VECM while Section 6 concludes.

2. Econometric methodology

Consider a congruent statistical system of unrestricted forms represented by Eq. (1):

$$x_t = \sum_{\tau=1}^{p} \Pi_\tau x_{t-\tau} + \varepsilon_t, \quad \varepsilon_t \sim \text{IN}(0,\Omega), \ t = 1,\ldots,T \tag{1}$$

In Eq. (1), x_t is a $(n \times 1)$ vector of $I(1)$ variables. Letting $\Delta x_t = x_t - x_{t-1}$, a convenient reparameterization of Eq. (1) is given by Eq. (2):

$$\Delta x_t = \sum_{\tau=1}^{p-1} \Pi_\tau^* \Delta x_{t-\tau} + \Pi^* x_{t-p} + \varepsilon_t, \tag{2}$$

where both Π_τ^* and Π^* are of dimension $n \times n$. This is the vector autoregressive (VAR) approach that Johansen (1988, 1991) and Johansen and Juselius (1990) used to investigate the co-integration properties of a system. Johansen and Juselius (1990) provide a full maximum likelihood procedure for estimation and testing within this framework. The lag length, p, is chosen to ensure that the errors are independent and identically distributed. Since ε_t is stationary, the rank, r, of the 'long-run' matrix Π^* determines how many linear combinations of x_t are stationary. If $r = n$, all x_t are stationary, and if $r = 0$ so that $\Pi^* = 0$, Δx_t is stationary and all linear combinations of $x_t \sim I(1)$. For $0 < r < n$ there exists r co-integrating vectors, meaning r stationary linear combinations of x_t. In this case, Π^* can be factored as, $\alpha\beta'$, where both α and β are $n \times r$ matrices. The co-integrating vectors of β are the error correction mechanism in the system while α contains the adjustment parameters. This result is known as Granger's Representation Theorem and can be found in Engle and Granger (1987).

The co-integrating rank, r, can be formally tested with two statistics. The first is the maximum eigenvalue test. Denoting the estimated eigenvalues as, λ_i^*, $i = 1, 2, \ldots, n$, the maximum eigenvalue test is given by

$$\lambda_{\max} = -T\ln(1 - \lambda_{r+1}^*) \qquad (3)$$

where the appropriate null hypothesis is $r = g$ co-integrating vectors against the alternative hypothesis that $r \leq g + 1$. The second statistic is the trace statistic and is computed as

$$\text{Trace} = -\sum_{i=r+1}^{n} T\ln(1 - \lambda_i^*) \qquad (4)$$

where the null hypothesis is $r = g$ and the alternative hypothesis is $r \leq n$.

As proven in Engle and Granger (1987) for the case of $I(1)$ variables, error correction and co-integration are equivalent representations. Consequently, if the variables do share a common stochastic trend then they are co-integrated and a vector error correction model (VECM) can be built. Granger (1986, 1988) also shows that co-integration implies that causality, in the Granger (1969) sense, must exist in at least one direction. The finding of co-integration among the variables in x_t is important because it means that there exists a long-run equilibrium relationship between these n variables. When the variables are co-integrated, short-run deviations from the long-run equilibrium will feed back onto the changes in the dependent variable in order to restore the long-run equilibrium. The coefficients on the lagged error correction term represent short-run adjustments to a long-run equilibrium. For any particular equation in the VECM, a statistically significant coefficient on the lagged error correction term means that a long-run Granger causal relationship exists. Conversely, a statistically insignificant coefficient on the lagged error correction term means that the dependent variable in the associated equation responds only to short-run shocks to the system (as measured by the lagged first differences of the n variables). Short-run Granger causal relationships can be tested using likelihood ratio tests on the coefficients of the lagged first differences of the explanatory variables in the VECM. Long-run Granger causal relationships can be tested using t-tests on the coefficients of the lagged error correction terms.

3. Integration properties of the data and systems co-integrating analysis

The natural logarithms of, crude oil futures prices, heating oil (#2) futures prices, unleaded gasoline futures prices, and the trade-weighted US exchange rate are denoted as c, h, g and e, respectively. The data are monthly and cover the period 1987:1−1997:9. These petroleum futures trade on the New York Mercantile Exchange and the futures price data are from the Prophet Information Services (1997) data bank. Following Fama and French (1987) a continuous series for each

commodity is constructed using the nearby futures contract with at least 1 month to maturity on the first trading day of each month. The trade-weighted US exchange rate is from the Federal Reserve Bank of Saint Louis (1998) data bank.

Fig. 1, which plots the actual values of the four series, indicates that none of the series exhibit trends. Consequently, all unit root test regressions were run with a constant but no trend term. Table 1 reports the results from Dickey and Fuller (1979) and Phillips and Perron (1988) unit root tests. The null hypothesis is a unit root in the series ($\alpha = 1$) while the alternative hypothesis is a stationary series ($\alpha \neq 1$). The results from augmented Dickey and Fuller (1979) regressions suggest that at the 1% level of significance, none of the series are stationary in levels. Both the Dickey and Fuller test results and the Phillips and Perron (1988) test results

Table 1

Unit root tests (augmented Dickey and Fuller (1979) test regression)[a] $\Delta x_t = \mu + \alpha x_{t-1} + \sum\limits_{j=1}^{k} c_j \Delta x_{t-j}$
$+ u_t$, $t = 1987{:}1\text{−}1997{:}9$

Variable	Test statistic	k
In levels		
c	−2.35	5
h	−1.89	7
g	−3.03**	2
e	−2.61*	5
In first differences		
c	−5.27***	7
h	−5.05***	8
g	−3.98***	10
e	−3.64***	8

Phillips and Perron (1988) test regression
$\Delta_{xt} = \mu + \alpha x_{t-1} + u_t$, $t = 1987{:}1\text{−}1997{:}9$

Variable	Test statistic	l
In levels		
c	−3.25**	4
h	−2.13***	4
g	−2.73*	4
e	−2.83*	4
In first differences		
c	−8.83***	4
h	−15.11***	4
g	−16.71***	4
e	−7.01***	4

[a]*Notes.* Critical values from Hamilton (1994). ***, **, * denotes a statistic is significant at the 1%, 5% and 10% level of significance, respectfully. The parameter, l, is a truncated lag parameter used in the non-parametric correction for serial correlation and is set according to the sample size. The parameter k is set using Perron's (Perron, 1997)) t-sig criterion.
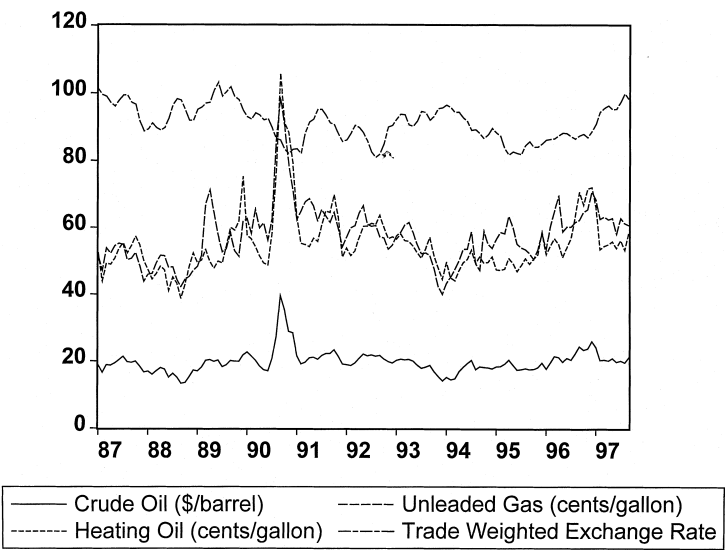
Fig. 1.   Energy futures prices and exchange rates.

indicate that each variable is stationary in first differences at the 1% level of significance. Since differencing the data once produces stationarity, I conclude that each of the series c, h, g, and e is integrated of order 1 ($I(1)$). As discussed in Johansen and Juselius (1990), while unit root tests are useful guides to the possibility of finding a co-integrating relationship they are not sufficient tests for co-integration. Consequently, it is more useful to test for a common stochastic trend among the time series under consideration. If the variables do share a

Table 2

Johansen tests for co-integration[a] $\Delta x_t = \sum_{\tau=1}^{p-1} \Pi_\tau^* \Delta x_{t-\tau}\ \Pi^* x_{t-p} + \varepsilon_t,\ -\Pi^* = \alpha\beta',\ x_t' = (c_t, h_t, e_t, g_t)$

Sample period: 1987:1−1997:9, $p = 6$

Eigenvalues	0.2138	0.0996	0.0642	0.0262
Hypothesis	$r = 0$	$r \leq 1$	$r \leq 2$	$r \leq 3$
Trace test	56.541**	25.510	11.980	3.424
λ_{max} test	31.031**	13.531	8.556	3.424
Co-integrating equation				
$z_t = c_t - 1.863\ h_t + 0.373\ e_t + 0.796\ g_t - 0.417$				

   [a]*Notes.* Critical values from Hamilton (1994). ***, **, * denotes a statistic is significant at the 1%, 5% and 10% level of significance, respectively.

common stochastic trend then they are co-integrated and an error correction model can be built.

Table 2 reports the values for the λ_{max} and the Trace test statistics. Block exogeneity tests in combination with tests for white noise residuals determined the appropriate lag length as $p = 6$. Both the Trace and λ_{max} test statistics indicate that a co-integration rank of 1 is present. The fact that these three energy futures prices are co-integrated with a trade-weighted exchange rate is important for several reasons. First, this finding of co-integration is important because it means that there exists a long-run equilibrium relationship between these four variables. Second, the existence of co-integration implies that there must be Granger causality between some of the variables. Third, co-integration implies that modelling must be done using a VECM rather than a VAR in first differences.

Using Eq. (2), a VECM model can be represented as

$$\Delta x_t = \sum_{\tau=1}^{p-1} A_\tau \Delta x_{t-\tau} + \alpha z_{t-1} + \varepsilon_t, \quad x'_t = (c_t, h_t, e_t, g_t) \tag{5}$$

where the error correction term (ect) is given by

$$z_t = c_t - 1.863 \ h_t + 0.373 \ e_t + 0.796 \ g_t - 0.417 \tag{6}$$

Eq. (6) can be rearranged so that c_t appears on the left-hand side. Since the variables are measured in natural logarithms, the coefficients are elasticities.

$$c_t = 1.863 \ h_t - 0.373 \ e_t - 0.796 \ g_t + 0.417 + z_t \tag{6a}$$

Eq. (6a) indicates that in the long-run equilibrium, a 1% increase in exchange rates lowers crude oil futures prices by 0.373%. This is consistent with Pindyck and Rotemberg (1990) who found that an equally weighted index of the dollar value of British pounds, German marks, and Japanese Yen negatively and significantly

Table 3

VECM results[a] $\Delta x_t = \sum_{\tau=1}^{p-1} A_\tau \Delta x_{t-\tau} + \alpha z_{t-1} + \varepsilon_t, \ x_t' = (c_t, h_t, e_t, g_t) \ z_t = c_t - 1.863 \ h_t + 0.373 \ e_t + 0.796 \ g_t - 0.417$

Sample period: 1987:1−1997:9, $p = 6$				
	Dependent variable:			
	Δc_t	Δh_t	Δe_t	Δg_t
α	0.119	0.396	−0.021	−0.017
p-value	0.278	$6.54e - 4$	0.498	0.881
$Q(6)$	0.570	0.610	0.999	0.853
$Q(12)$	0.239	0.832	0.999	0.240
$Q(24)$	0.583	0.932	0.952	0.420

[a]*Notes*. $Q(6)$, $Q(12)$ and $Q(24)$ are Ljung−Box tests for serial correlation.

impacts the price of crude oil in both ordinary least squares regressions and latent variable models.

The Johansen procedure can be rather sensitive to residual serial correlation. Consequently, the bottom of Table 3 presents Ljung−Box Q statistics for testing white noise residuals. The reported Q statistics indicate that the residuals from the VECM are white noise and therefore the VECM specification with $p = 6$ is adequate.

4. Granger causality tests

The upper portion of Table 4 reports probability values for testing short-run Granger causality. These tests can be computed from Eq. (5) by testing the null hypothesis $A_{ij\tau} = 0$, $\tau = 1, 2, \ldots, p$ in the VECM. These hypothesis are tested using likelihood ratio tests of the form $(T - c)(\log |\Sigma_R| - \log |\Sigma_U|)$ where T is the number of observations, c is a small sample correction equal to the largest number of parameters estimated in any one of the equations from the VECM and log $|\Sigma_R|$ and log $|\Sigma_U|$ are the natural logarithms of the determinants of the variance−co-variance matrices from the restricted and unrestricted VECMs, respectively. This statistic is asymptotically distributed as chi-squared with degrees of freedom equal to the number of coefficient restrictions. The null hypothesis of no Granger causality running from the change in the natural logarithm of exchange rates to the change in the natural logarithm of crude oil futures prices can be rejected at the 10% level of significance. The null hypothesis of no Granger causality running from the change in the natural logarithm of exchange rates to the change in the natural logarithm of heating oil futures prices can be rejected at the 5% level of significance. The results in Table 4 also indicate that in the short-run, exchange

Table 4
Granger causality tests

Sample period: 1987:1−1997:9 (*P*-values are shown).				
From:	To (dependent variable):			
	Δc_t	Δh_t	Δe_t	Δg_t
Tests of no short-run Granger causality				
Δc_t	0.019	0.487	0.856	0.040
Δh_t	0.143	0.324	0.945	0.133
Δe_t	0.099	0.043	0.002	0.316
Δg_t	0.182	0.206	0.687	0.346
Tests of no long-run Granger causality				
z_{t-1}	0.278	$6.54e - 4$	0.498	0.881
Joint tests of no long-run and no short-run Granger causality				
	0.024	$1.71e - 3$	0.792	0.170

rates are primarily driven by their own past values while short-run movements in unleaded gasoline futures prices are primarily driven by movements in crude oil futures prices.

The null hypothesis for testing long-run causal relationships is $\alpha_i = 0$, $i = 1, \ldots, 4$. From Table 4 it is clear that the error correction term is statistically significant at the 1% level in the heating oil equation. These results indicate that the futures prices of heating oil adjust to clear any deviations from long-term disequilibrium.

The fact that the exchange rate equation contains no evidence of either long-run Granger causality or short-run Granger causality suggests that exchange rates are exogenous to the system. The results from formal joint tests of no long-run causality and no short-run causality for each variable are shown in the lower portion of Table 4. The joint hypothesis of no long-run causality and no short-run causality can be rejected at the 5% (1%) level of significance for crude oil futures prices (heating oil futures prices) but cannot be rejected for either exchange rates or unleaded gasoline futures prices. These results suggest that exchange rates transmit exogenous shocks to the system. These results are supportive of the conjecture by Bloomberg and Harris (1995) that recent movements in commodity prices may be a response to movements in the dollar.

5. Stability of the VECM

The estimation period for this study covers the somewhat turbulent time of the 1990 Gulf crisis. Consequently, it is important to check the VECM for structural breaks. One way of checking parameter stability is to estimate the VECM recursively and plot the recursively estimated coefficients, with associated standard error bands, of the error correction terms. Fig. 2A−D show the results from doing this. The plots shown in Fig. 2A−D suggest that a structural stability problem probably does not exist since, for each equation, the size of the estimated coefficients on the error correction terms are fairly similar over the entire sample period.

As a further test of parameter stability, I compute Gregory and Hansen (1996) tests for co-integration in models with regime shifts. These tests are designed to test the null hypothesis of no co-integration against the alternative of co-integration in the presence of a possible regime shift. Gregory and Hansen (1996) define four models.
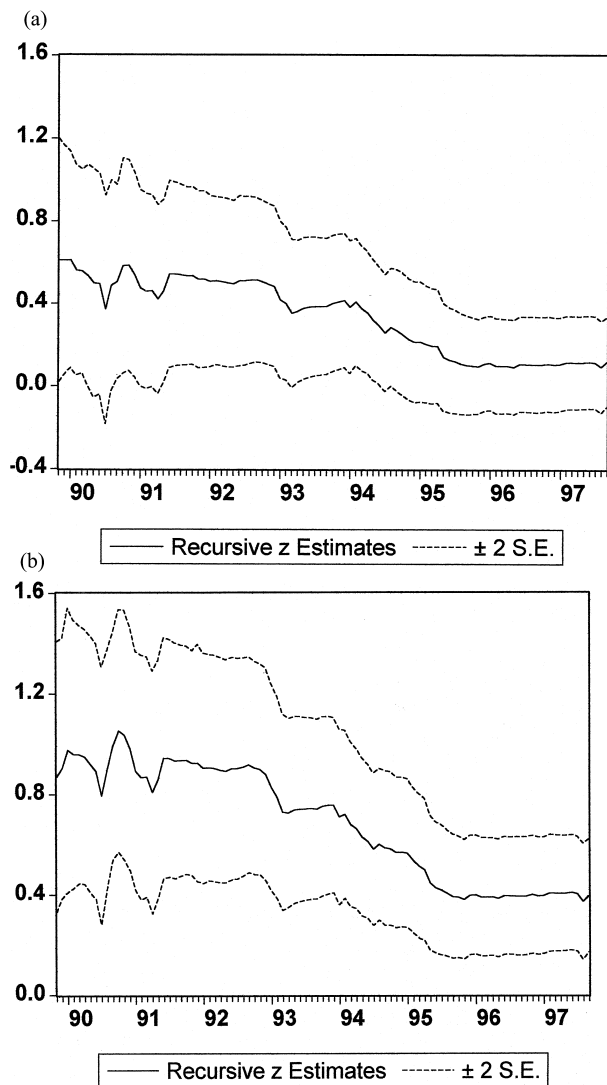
Model 1: standard co-integration.

$$y_{1t} = \mu_1 + \alpha'_1 y_{2t} + e_t, \quad t = 1, \ldots, T$$

where

$$y_{1t} = c_t \quad \text{and} \quad y_{2t} = (h_t, e_t, g_t)$$

Structural change can occur at an unknown point in time and is modelled using a dummy variable. The next three models allow for possible regime shifts in the

(a)

(b)

co-integration equation.

$$\varphi_{t\tau} = 1 \text{ it } t > T^B \quad \text{and 0 otherwise where } T^B \text{ is the break date.}$$

Model 2: level shift (C).

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \alpha'_1 y_{2t} + e_t, \quad t = 1, \dots, T$$
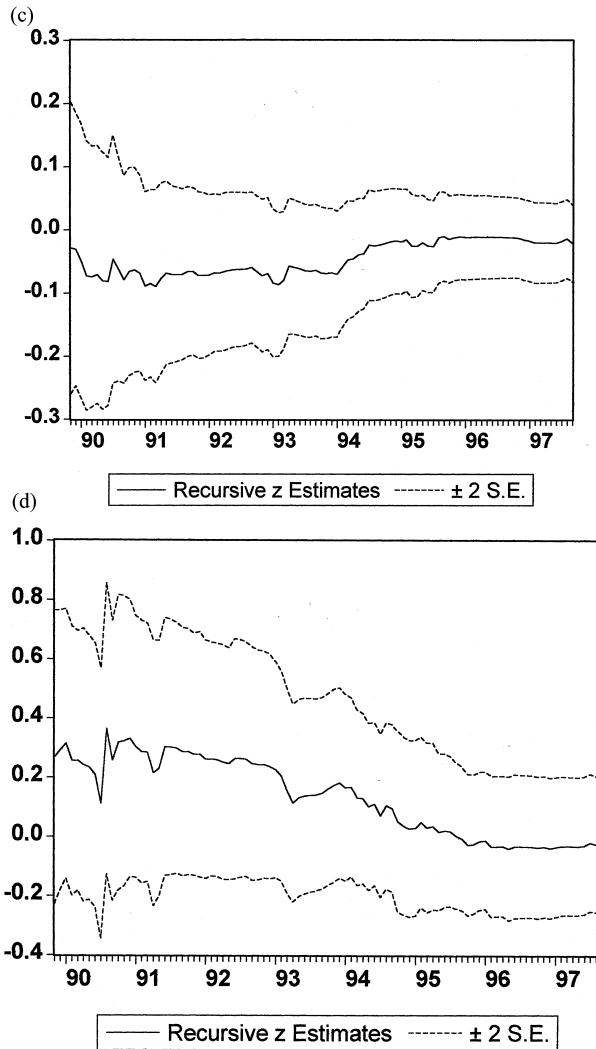
(c)

(d)

Fig. 2. (a) Recursive estimates from crude oil equation. (b) Recursive estimates from heating oil equation. (c) Recursive estimates from exchange rate equation. (d) Recursive estimates from gasoline equation.

Model 3: level shift with trend (C/T).

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \beta t + \alpha'_1 y_{2t} + e_t, \quad t = 1, \dots, T$$

Model 4: regime shift (C/S).

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \alpha'_1 y_{2t} + \alpha'_2 y_{2t} \varphi_{t\tau} + e_t, \qquad t = 1,\ldots,T$$

The computation of the test statistics are straightforward. For each break point, T^B, estimate one of the models 2 through 4 (depending upon the alternative hypothesis under consideration) by OLS yielding the residuals $e^*_{t\tau}$. From these residuals compute the usual augmented Dickey and Fuller test statistic ($\text{ADF}(\tau) = \text{tstat}(e^*_{t-1\tau})$) and define ADF^* as the smallest value of the $\text{ADF}(\tau)$. The numerical value of ADF^* can be compared to the critical values tabulated in Gregory and Hansen (1996).

Both the ADF and ADF^* statistics test the null hypothesis of no co-integration. Gregory and Hansen (1996) offer some suggestions as to how to use the ADF^* statistic. Rejection of the null hypothesis of no co-integration by either the conventional ADF (or any other standard test for co-integration) or the ADF^* statistic indicates some long-run equilibrium relationship in the data. If the ADF statistic does not reject but the ADF^* does reject then structural change may be important. If both the ADF and the ADF^* reject then no information of structural change is provided since the ADF^* is powerful against conventional co-integration.

Table 5 reports ADF^* statistics for testing regime shifts in each of models 2 through 4. The lag length for the ADF^* statistics are chosen using Perron's (Perron, 1997) t-sig criteria. For illustration purposes consider Eq. (7).

$$e^*_{t\tau} = \alpha e^*_{t-1\tau} + \sum_{j=1}^{k} c_j \Delta e^*_{t-j\tau} + u_t \tag{7}$$

Working backwards from 13 lag lengths, the first value of k is chosen such that the t statistic on c_k is greater than 1.6 in absolute value and the t statistic on c_l for $l > k$ is less than 1.6 in absolute value. Table 5, which presents the results from testing models 2–4, shows clear evidence against any structural changes in the co-integrating equation. For Model 1, the conventional ADF statistic is -3.89 (which is significant at the 1% level of significance). Since the ADF, Trace and $\lambda_{\max}$ test statistics (from Table 2) each rejects the null hypothesis of no co-integration, while the ADF^* statistic doesn't, it can be concluded that the VECM is a structurally stable econometric model.

6. Concluding remarks

The main findings of this study can be summarized as follows. Results are presented showing that futures prices for crude oil, heating oil and unleaded gasoline are co-integrated with a trade-weighted index of exchange rates. This is important because it means that: (1) there exists a long-run equilibrium relationship between these four variables; (2) there must be Granger causality between some of the variables; and (3) co-integration implies that modelling must be done using a VECM rather than a VAR in first differences. Granger causality results for both the long- and short-run are presented. Movements in exchange rates precede

Table 5
Tests for co-integration in models with regime shifts[a]

Model	ADF*	Break date	l
C	−3.46	1988:09	11
1%, 5% and 10% critical values are: −5.77, −5.28 and −5.02, respectively			
C/T	−3.67	1988:08	11
1%, 5% and 10% critical values are: −6.05, −5.57 and −5.33, respectively			
C/S	−3.63	1988:10	11
1%, 5% and 10% critical values are: −6.51, −6.00 and −5.75, respectively			

[a]*Notes*. Critical values from Gregory and Hansen (1996). The parameter l is set using Perron's (Perron, 1997)) t-sig criterion.

movements in heating oil futures prices in both the long- and short-run while movements in exchange rates precede movements in crude oil futures prices in the short-run. In the long-run equilibrium, a 1% increase in exchange rates lowers crude oil futures prices by 0.373%.

Evidence is also presented that suggests exchange rates transmit exogenous shocks to energy futures prices. This is supportive of the conjecture by Bloomberg and Harris (1995) that recent movements in commodity prices may be a response to movements in the US dollar.

Acknowledgements

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