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Extracting Long-run Information from Energy Prices-The role of Exogeneity

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Abstract

This article considers cointegration analysis to detect key features of long-run structure in the gasoline market. The main purpose of this study is to investigate possible long-run price leadership in the US gasoline market and the characteristics relevant to a competitive market using the vector autoregressive model. After examining the stationarity and cointegration of the weekly gasoline prices in eight different regions of the US we research long-run price leadership and parallel pricing in the framework of the cointegrated vector auto-regression (VAR). The problem is considered on extended data on 901 weekly gasoline prices for eight regions of the US. The discovery of a single common trend has been observed for a smaller number of regions, but when the system is estimated across the US it is found that the single common trend cannot be sustained. In addition to this failure the tests of exogeneity suggest that the extent to which regional gasoline prices in the long-run respond to each other is limited.

Keywords: Arbitrage, Law of one price, Cointegration, Error Correction Model, Long-run relationship, Weak exogeneity, Price dispersion.
JEL Classification: C22, C13, R11, D4

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1. Introduction

Gasoline is one of the products with the highest price variation in the world and the current dramatic changes in gasoline prices significantly affect the consumer and business behaviour in the market. The gasoline price is significantly influenced by innovation, technological progress, and political instability in the global economy. The core process in the production of gasoline the oil field to the gas station pump is observed in terms of four main steps: oil exploration, refining, distributing the refined oils to the different companies and regions, selling the product. The price of the gasoline at the pump includes a considerable amount of tax which is one of the vital revenue streams for the government.

The gasoline market has generally been considered as competitive, because the product is homogeneous, there are strict rules as to what can be added to fuel, consumers are less influenced by branding, there are many suppliers and consumers and a significant amount of price related information is commonly available. Nevertheless pump prices at the gas station do differ in terms of location, local tax levels and services provided by the outlet.

Observing the process that gives rise to equilibrium in a market can confirm the appropriateness of the structure and the completeness of a market. Price disequilibria in the long-run between neighbouring regions would affect regional activity and consumers might react radically towards high price differentials by moving job and/or house to reduce travel costs, by the purchase of more fuel efficient vehicles etc., but the persistent price differential suggest discrimination and identifies the possibility of some market power and informational inefficiency.

In this article we discuss further the developments in the literature previously summarized in Hendry and Juselius (2001), Hunter and Burke (2012), Hunter and Tabaghdehi (2013) among others. That information on price can be provided efficiently to customers and that consumers can monitor retail gasoline prices is one of the main concerns for the global economy. To this end government intervention and regulation may be required to control price discrepancy and improve market structure.

In section I we introduced the gasoline market and econometric model used in the analysis. In section II we reviewed of essential literature on stationarity, error correction, and exogeneity. Section III identified the data for the empirical analysis. Section IV we reviewed the price analysis and cointegration. In part V we test for weak exogeneity, long-run exclusion, and strict exogeneity to investigate the nature of parallel pricing in the gasoline market. Finally, in Section VI we conclude.

2. Review of Essential Literature

The “law of one price” implies that in any market that the price of goods identical in terms of quality and specification must tend to be the same for an efficient market regardless of where they are traded. The law of one price can be reformulated in the case of transport and transaction cost. When prices at different locations differ as a result of transport and transaction cost, the arbitrage opportunity will close the price gap after a short period of time. The error correction model (ECM) is one of the appropriate mechanisms for analyzing the law of one price in the long-run (Johansen, 1995). Vanya and Walls (1999), Hendry and Juselius (2001) and La Cour and Møllgaard (2002) identified that cointegration analysis is a practical mechanism for measuring competition in markets.

Forni (2004) suggests the use of stationarity tests of the log price differential to determine a broad (stationary) as compared with a narrow (non-stationary) market. A broad market identifies that the market is competitive. To this end Giulietti et al (2010), and Hunter and Tabaghdehi (2013) have also applied univariate and panel stationarity as compared with non-stationary tests to study gasoline market competitiveness.

Following Forni (2004) we applied a range of time series methods to test whether log price differentials were stationary as a monitoring procedure to examine the potential for an anti-trust case in the gasoline market, and investigate the structure of the gasoline market without any need to normalize on a specific price in the long-run or condition the problem relative to a specific price seen to be exogenous. To address the issue of exogeneity and the interrelatedness of prices in this article we consider the VAR in error correction form.

Analysing price properties may be effective in testing for "market definition" when the persistence of the volatility is reduced by this transformation of the data. If volatility is quite persistent (the largest eigen value – spectral radius of the ARCH[1] polynomial exceeds .85) then the Johansen test may only converge to the asymptotic distribution for sample sizes in the range 600-1000.[2]

The approach in this study is appropriate for more extensive data sets of the variety we have here. It is hoped that it will be possible to verify the competitive behaviour in the market from the long-run decomposition of prices. Consequently we use the conditional ECM and VAR approach for testing cointegration, to improve and develop the long-run relationships and consider the potential for arbitrage correction in the gasoline market.

Any specific patterns of pricing behaviour in the market that can give rise to profitable opportunities from arbitrage cannot survive for long and over time they will dissipate as others seek them out (Fama, 1998). However, in the US gasoline market the

[1] The notion of Autoregressive Conditional Heteroscedasticity (ARCH) relates to Engle (1982) and was first applied to price data for the UK.
[2] The interested reader is directed to the simulation results obtained by Rahbek et al (2002).

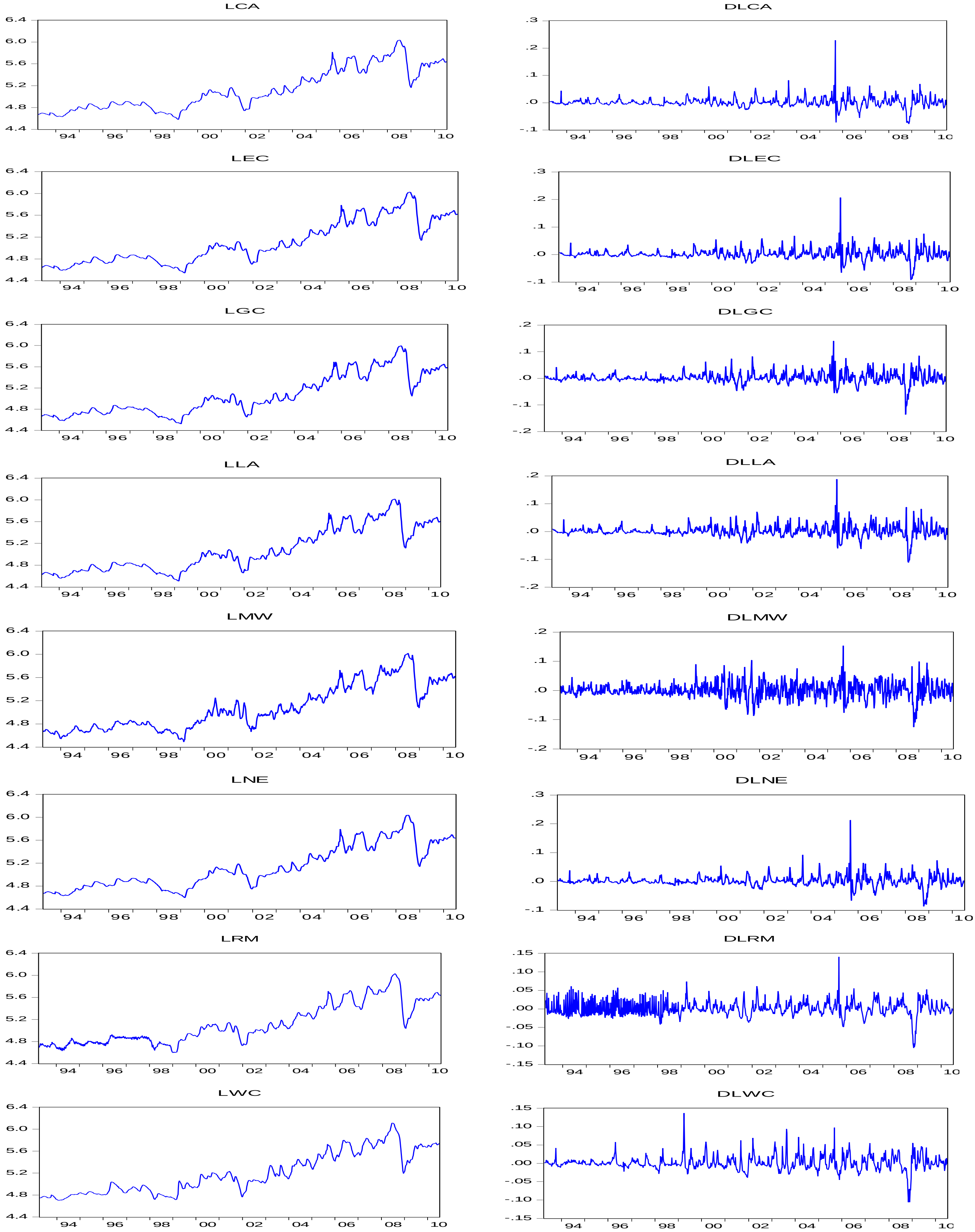
specific pattern of regional price differentiation may be constantly affecting market efficiency. Hence we employ the cointegration methodology of Johansen (1995) to test empirically the definition of the market and the nature of integration of the price series. However energy storability makes it suitable for price arbitrage and hedging. When considering the price of gasoline in the different region of the US it is possible to observe opportunities for location arbitrage. Consequently to tackle arbitrage opportunities in a market-oriented industry to address market power there needs to be some form of regulation (Küpper and Willems, 2010). However, poor regulation in the gasoline market would distort competition.

3. Time-series properties of the data

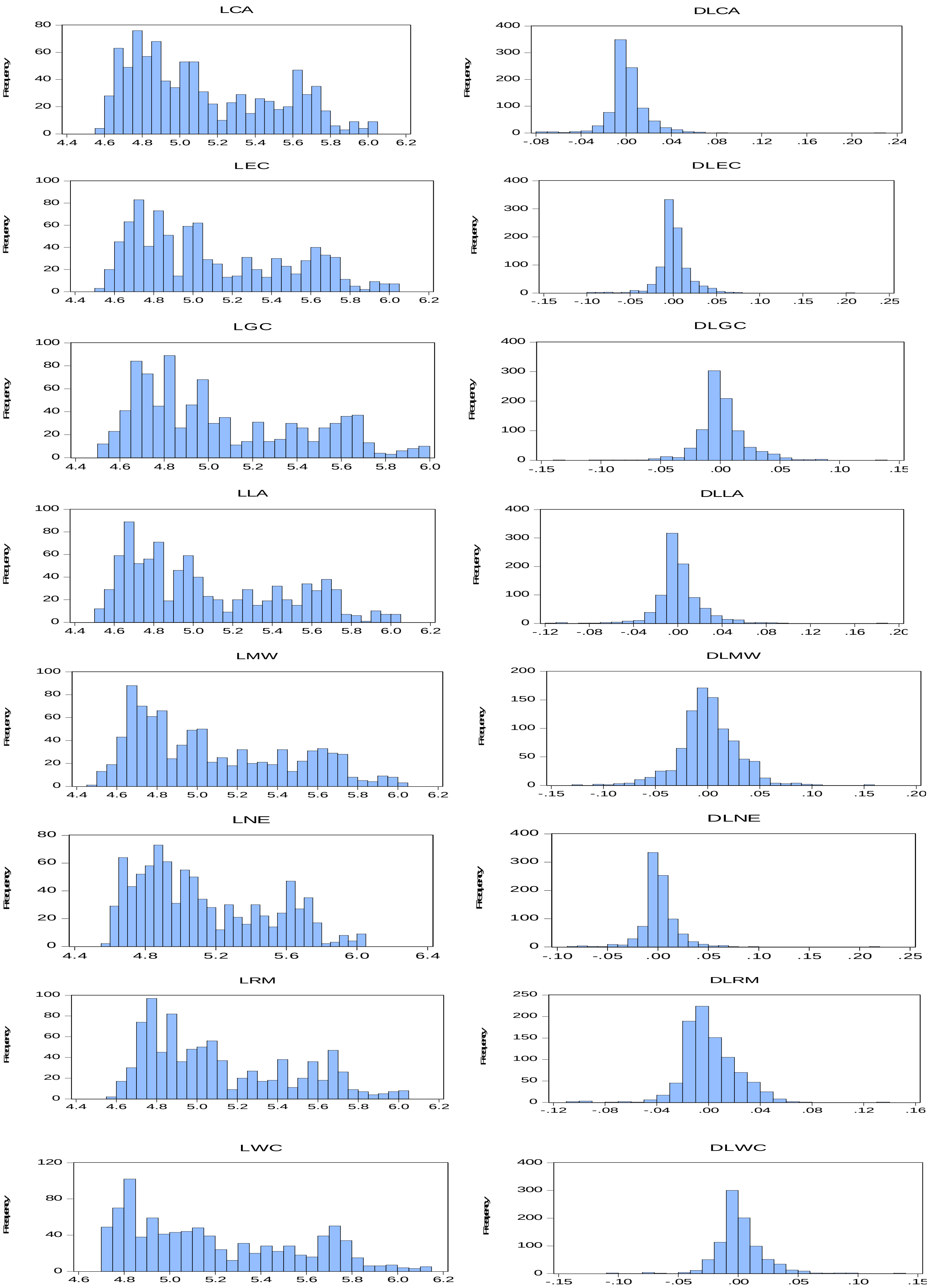
In this section we consider the time series properties of a data set consisting of weekly gasoline prices across eight different regions in the US (West Coast (WC), Central Atlantic (CA), East Coast (EC), Gulf Coast (GC), Lower Atlantic (LA), Midwest (MW), New England (NE), Rocky Mountains (RM)) from May 1993 to May 2010.[3] Considering regional gasoline infrastructure across the US we test cointegration on eight different regions and nine price differentials. The data in (log) levels and (log) differences are graphed in the plots in figure 1 and the frequency distributions of the data are graphed in figure 2. From figure 1, the price level drifts upwards, whereas the price differences appear to move randomly around a fixed mean. While, the frequency distributions of the price level in figure 2 suggests non-stationarity whereas the frequency distribution of the differences suggest the series are closer to normality.

[3] The data have been obtained from the energy information administration website (www.eia.doe.gov).

Plot 1- Gasoline price at eight US locations in (log) levels and (log) differences



Plot 2- Frequency distribution of Gasoline price at eight US locations in (log) levels and (log) differences



4. Price analysis, Cointegration, and Arbitrage Correction in gasoline market

Time series might be non-stationary as a result of technological progress, economic evolution, crises, changes in the consumers' preference and behaviour, policy or regime changes, and organizational or institutional improvement. However, regressions based on stochastic non-stationary series simply as a result of cumulating the events or shocks of the past may give rise to 'nonsense regression', and this can cause significant problems in forecasting and inference (Hendry and Juselius, 2000).

Following Hosken and Taylor (2004), and Kurita (2008) we analysed the cointegration and exogeneity properties of regional gasoline prices in the US, but using all the regional data across the US mainland except for Alaska.

De Vanya and Walls (1999), Hendry and Juselius (2001), and Forni (2004) proposed that finding a single cointegrating vector suggests that the market is efficient. Forni (2004) suggested that a broad market implies *(N-1)* cointegrating vectors for *N* price series while a narrow market implies less than *(N-1)* cointegrating relations. However with *(N-1)* cointegrating vectors we might observe long-run causality (cointegrating exogeneity, Hunter, 1990) or detect that one of the prices is weakly exogenous (WE) for all the cointegrating vectors (Johansen, 1992).

Buccarossi (2006) and Forni (2004) suggest that when comparison is made with the prices of two companies or regions then competitive behaviour is consistent with parallel pricing and this can be confirmed by testing for the stationarity of log price proportions. Forni (2004) suggested that a sequence of stationarity tests might be used to measure market definition. Such an approach has merit when the data is limited by the extent of the time series, but Hunter and Burke (2007) suggest that univariate time series analysis does not provide a formal mechanism by which it may be confirmed that there are *N-1* such relations. They show that this may be better tested in a multivariate context and that it is possible to distinguish between a case where arbitrage holds and all the series follow a common stochastic trend and the case where there is aggressive price leadership or a single variable is WE for the matrix of cointegrating vectors (β).

In a bivariate case of gas prices conditioned on a WE oil price, Hendry and Juselius (2001) find that competition implies a common trend driving prices across markets and this can be generalized to a multi-price framework. We pay particular attention to the role of the common trend and exogeneity in explaining the competitive structure using a large time series sample.

In this study to determine potential long-run equilibrium relations on US gasoline prices in different regions, first for comparison with the stationarity testing of Forni (2004) we develop a cointegration analysis based on a bivariate model:

$$p_{at} = \mu_0 + \beta p_{bt} + u_t, \qquad (1)$$

where p_{at} and p_{bt} are prices of gasoline in two different regions of the US, and u_t is a random disturbance term. Here μ_0 represents the log of the proportionality coefficient. $\mu_0 = 0$ when the prices in different regions are identical, and $\mu_0 \neq 0$ if there is a fixed transportation cost and fixed quality differences in the product related to different regions. However with a perfectly integrated market the price reflects all available information and traders cannot benefit consistently from any arbitrage opportunities. Equation [1] is a cointegrating regression where β explains the nature of the relation between the regional prices. The hypothesis related to parallel pricing implies that $\beta = 1$ is tested. $\beta = 0$ implies no relation between the prices of the two regions, and otherwise ($\beta = 1$) specifies that regional prices are proportional and it confirms the law of one price

(LOP), although the observed may differ from 1 by an arbitrary constant(c) where $| \beta - 1| \leq c$. In the case of perfect integration c is close to zero.

According to Engle and Granger (1987) the linear combination of two non-stationary series of p_{at} and p_{bt} can be transformed to stationarity:

$$\eta_t = p_{at} - \beta p_{bt} \qquad (2)$$

and $\eta_t = \mu_0 + u_t \sim I(0)$. The latter embodies the notion of cointegration that two (or more) I(1) series, here p_{at} and p_{bt}, give rise to a relation that is stationary. Therefore when η_t represents a residual from a regression, then when this combination is stationary there is a long-run relationship between p_{at} and p_{bt} otherwise this relates to a nonsense regression. Consequently for the price of any homogeneous good in an identical market a cointegrating relation is necessary as arbitrage would clear mispricing in the long-run.

One difficulty with the Engle and Granger (1987) test is the nonstandard nature of the statistical inference and that it does not provide a direct test of the law of one price (Forni, 2004). However, the methodology developed by Johansen (1995) can be applied to test the LOP in a VAR and the potential for price leadership.[4] When the gasoline prices of different regions in the US are identical, then the associated market will be in equilibrium, otherwise there would arbitrage opportunities across all regions. This trading mechanism will be inclined to equalize the prices in the long term by raising prices in the low-price regions and lowering prices in the high-price regions.

In empirical modelling multivariate time series analysis is applied to estimate long-run equilibrium relations. The ECM provides one method to investigate the nature of adjustment across prices to determine long-run equilibrium, see Patterson (2000).

[4] The Engle-Granger cointegration approach is based on the assumption that the univariate processes are I(1) and the vector moving average (VMA) process has a partial over difference that relates to a reduced rank in VMA polynomial. Whereas the Johansen cointegration approach assumes that the parsimonious time series representation is given by a VAR process. The Johansen methodology is based on the maximum likelihood approach and the Johansen test is a likelihood ratio test statistic clearly explained along with the full test methodology in Johansen (1995).

We investigate long-run equilibrium in the US gasoline market using the error correction model that is also termed arbitrage correction by Burke and Hunter (2012). The hypothesis underlying this argument relates to the possibility that the sequence of regional gasoline prices that deviate from equilibrium give rise to an arbitrage opportunity that is correcting in the long-run when there are *N-1* arbitrage correction terms across *N* markets (Burke and Hunter, 2011).

According to Kremers, Ericsson, and Dolado (1992) the ECM is a good model to detect long-run behaviour. The single equation ECM is a starting point for modelling, which binds the cointegration relationship in the long-run and as a result of super consistency (Ericson and MacKinnon, 2002) the approach is robust to specific lag lengths and model dynamics. However the ECM might not accurately identify the suitable long run relationship in the presence of structural change as this may result in finding inconsistent cointegrating relations that are poor in terms of prediction (Clements and Hendry, 1995).

To further investigate the short-run dynamics of the relationship in gasoline prices of different regions in the US we employ a vector error correction model (VECM) specification. For example, Bachmeier and Griffin (2006) found that the prices of crude oil in different geographical regions of the world are cointegrated. De Vanya and Walls (1999) using the ECM, identified cointegration between eleven regions of the US in relation to electricity prices.

The first step of the Engle and Granger (1987) method identifies equilibrium relations from a cointegrating regression that gives rise to an error correction term estimated from the OLS regression residual:

$$\hat{\eta}_t = e_t = p_{at} - m_o - bp_{bt} \qquad (3)$$

We may test whether these series are stationary by applying the Dickey-Fuller test to these residuals:

$$\Delta\hat{\eta}_t = \gamma\, \hat{\eta}_{t-1} + v_t \qquad (4)$$

$$\Delta\hat{\eta}_t = \gamma_0\, \Delta p_{bt} + \gamma\, \hat{\eta}_{t-1} + \epsilon_t \qquad (5)$$

where: $\epsilon_t = \gamma_0\, \Delta p_{bt} + v_t$ then $\beta = \beta_0$ and $\gamma_0 = \beta - \beta_0$.

It is also possible to have cointegration when $\gamma < 0$, but this may not be consistent with efficiency as $\mu_o \neq 0$ and $\beta \neq 1$. In the long-run the prices are set to their long-run average values $p_{at} = \check{p}_{at}$, $p_{bt} = \check{p}_{bt}$. If the long-run is appropriately defined in terms of these long run average prices so that:

$$\check{p}_{at} = \mu_0 + \beta\check{p}_{bt}.$$

Where the μ_0 and β are long-run parameters and for efficiency in the market and to avoid persistent long-run profits being exploited from arbitrage possibilities we require $\mu_0=0$, $\beta=1$ and then:

$$p_{at} = p_{bt} + \eta_t \text{ or } \eta_t = p_{at} - p_{bt}\,;\ \hat{\eta}_t = \eta_t.$$

It follows that the ECM gives rise to a long-run relation restricted to the same form as the Dickey-Fuller model used to test stationarity (Dickey and Fuller, 1979). It is shown in Kremers et al (1992) that the Dickey Fuller (DF) test that is applied by Forni (2004) is a special case of a pure ECM (see Davidson et al, 1978). Therefore:

$$\Delta\,(p_{at} - p_{bt}) = \gamma\,(p_{at-1} - p_{bt-1}) + v_t. \qquad (6)$$

Equation (6) is a restricted version of the model applied at the second step of the Engle-Granger approach where the lagged equilibrium error is defined by Hendry (1995) in this more general case as an equilibrium correction model (EqCM). Here we follow the pure ECM approach where $(p_{at-1} - p_{bt-1}) = \eta_t \sim I(0)$ indicates that the ECM defines the equilibrium error or when $\eta_t \sim I(1)$ this is not an equilibrium error. The γ in equation (6) is a short-run parameter and specifies how quickly the disequilibrium will be removed from the system or the speed at which arbitrage occurs[5]. Therefore the larger the absolute value of γ the more quickly any disequilibrium or mispricing will be removed. The null hypothesis H_0: $\gamma = 0$ tests the significance of the error correction coefficient, when compared with the one sided alternative of H_A: $\gamma < 0$.[6] The acceptance of H_A is evidence supporting cointegration and market efficiency.

The error correction representation exists if p_{at} and p_{bt} are cointegrated. Furthermore, with *N* price variables adapting the results in Smith and Hunter (1985) to the non-stationary case, there are *1/2N(N-1)* non-trivial combinations of error or cross arbitrage correction terms between all the prices. Such relations are termed coherent by Smith and Hunter (1985) when the slope coefficients are the same and for pure arbitrage that is unity. The zero intercept restriction is not critical to the argument though it gives rise to the same error correction applying in the long-run for all these combinations. It follows from Smith and Hunter (1985) in relation to the cross arbitrage for exchange rates that in the coherent case when *N-1* stationary relations are found, then by simple algebraic manipulation and the stationarity of the primary relations the remaining *½(N-1)(N-2)* should also be stationary. Non-coherence implies that different stationary or some non-stationary combinations may arise and as a result some of the long-run relations may include all the prices.

The results for the augmented Dickey Fuller (ADF) test and ECM estimations are compared in Table 1.[7] Acceptance of the alternative hypothesis underlining the ADF tests imply that the price proportions related to eight combinations are stationary based on a one sided test at the 5% level. Significant results indicate that the series

[5] ***γ %*** of the disequilibrium at time ***t-1*** is removed in period ***t***.
[6] **γ** >0 implies that variables are moving in the wrong direction to correct for disequilibrium.
[7] All estimations are undertaken using Oxmetrics Professional (Doornik and Hendry, 2009).

move in proportion to each other in the long-run, but any rejection of the alternative may arise as a result of the bivariate analysis of the problem.

[Table 1 goes here]

In the case of the ECM, testing for cointegration follows from an analysis of each single equation in turn via the significance of the error correction term. In all but one case the error correction terms are significant, this one exception may arise due to a lack of cointegration or weak exogeneity,[8] or the cointegrating relation cannot be identified from a single error correction term in a single equation dynamic model. In the case of the ADF test this may arise, because as was considered above this model imposes efficiency on both the short-run and the long-run relations. This is given support by the observation that this coefficient is significant in the error correction model for the GC and the LA.

Based on Dickey Fuller inference (Patterson, 2000) the coefficient on the error correction term is not significant in two cases that relate to the RM and the WC relative to the GC. However, according to Kremers, Ericsson and Dolado (1992), the error correction test is asymptotically normal, but converges at a slower rate than is usual with conventional inference (Ericsson and MacKinnon, 2002). Assuming such convergence and normal inference the only insignificant case would relate to the WC. The latter may arise for three reasons, the most obvious when comparison is made with the ADF tests would be that the model is over-parameterised or the test inefficient as a result of the number of lag terms included in the model, this relation may also be inefficient or the RM model may not contain an error correction term as this price is WE as it forces, but is not forced by the rest of the US market. The rejection of cointegration may also be a function of the bivariate nature of these models.

[8] Single equations may suggest more WE variables than can arise when the test is applied at the level of the system.

In further investigating the system we follow Boswijk (1992), Hunter and Simpson (1996), and Bauwens and Hunter (2000) and apply restrictions on **α**, **β** (dimensioned N×r), and **α** as well as **β** to study the exogeneity structure of the data and identify potentially WE variables.

Following Forni (2004) the test of stationarity jointly tests $\mu_0=0$, $\beta=1$ and $\gamma<0$, finding stationarity implies both cointegration and a relation with $\mu_0=0$, $\beta=1$. The VAR under the assumption of normality of the errors and based on the notion that there is drift in the individual price series implies an unrestricted intercept to determine the number of long-run relations or cointegrating rank (*r*). It can be shown (Hunter and Burke, 2007) that when all the series are I(1), then there are *r=N-1* cointegration relations in the competitive case and a single common trend.

The following equation is a VECM that is a re-parameterisation of a VAR:

$$\mathbf{\Gamma}(\mathbf{L})\,\Delta\,\mathbf{p}_t = \mathbf{\Pi p}_{t-1} + \boldsymbol{\mu} + \boldsymbol{\varepsilon}_t.$$

Where $\mathbf{\Gamma(L)} = (\mathbf{I} - \Gamma_1 L - \ldots - \Gamma_{k-1} L^{k-1})$, Γ_i are $N\times N$ matrices and **I** and *N* dimensioned identity matrix. The hypothesis that relates to the cointegrating rank is:

$$H_1(r)\text{: } \mathbf{\Pi} = \alpha\beta'.$$

Using the Johansen trace test we identify the number of cointegrating vectors (*r)* and the number of common trend when there is less than *N-1* cointegrating relations. The results on the Johansen trace test for eight regional gasoline prices in the US are presented in Table 2. The results related to this test indicate that there are *r=4* cointegrating vectors for a test applied at the 5% level. This implying that there are *N-r=4* stochastic trends, this is not consistent with the results that arise when cointegration is tested based on the single equation tests of stationarity.[9] When *r<N-1* there are more stochastic trends than might be anticipated by a single competitive

[9] Result can be provided in request.

market implying that the market is partitioned or that there may be features of the data not characterised by the model.

[Table 2 goes here]

Further analysis is required to interrogate the nature of the inter-relations that may impact price behaviour. Each long-run relation will be forced by up to four trends so there may be four different price relations that are driving the system in the long-run.

There may also be the type of separation in the market place related to cointegrating exogeneity and quasi-diagonality (Hunter, 1992) or weak exogeneity (Johansen, 1992). In the first instance gas prices in different parts of the US may respond to a different stochastic trend or in some parts of the US there may be relations linked to all the trends and in others to a subset of trends. Up to four variables may also be WE implying that they are not affected by the long-run price behaviour in the other segments of the market.[10] Such segmentation may be consistent with price differentiation and these anomalies are indicative of collusive agreements or when long-run causality can be detected there is potential for leadership by some of the major gasoline supplier's.

In the case where we have a large amount of data the VAR would appear appropriate. Here we progress the VECM as the cointegrating rank is only properly testable in the VAR when r>1.

5. Exogeneity and causality analysis- Test of weak exogeneity and parallel pricing

Granger (1969) devised a means to test for causality in the context of stationary series, while the concept of cointegrating exogeneity was developed by Hunter (1990) to handle causality between non-stationary variables in the long-run. Giannini and

[10] See Chapter 5 of Burke and Hunter (2005) for further discussion of weak exogeneity related to sub-blocks of the cointegrating vectors.

Mosconi (1992) tested Granger Causality subject to CE. Testing for causality has been found useful by Horowitz (1981), Ravallion (1986), Slade (1986), and Gordon, Hobbs, and Kerr (1993) for defining market boundaries. Here, subject to the finding on rank, the focus will be on exogeneity restrictions and long-run exclusion.

Analysing single equations from the VAR, econometrically and theoretically is less restrictive. At one level the ADF test imposes a common factor restriction that relates to efficiency being imposed on the short-run relations, thus causing arbitrage to be imposed on the short-run parameters. Hence by estimating the VAR and relating this to the ECM, we can determine whether there is market segmentation and the nature of arbitrage across the system. Following Hendry and Juselius (2001) we consider the conventional VECM, but with eight inter-related market prices.

The VECM model (7) applied here is based on a VAR(k) where Δp_t is stationary the error term is stationary and based on the previous analysis the *N* variables give rise to *r* long-run relations where $0<r<N$. It would seem clear from the analysis conducted in this study thus far that the finding of no cointegration ($r=0$) can be rejected. However, a generous or more careful interpretation of all the above analysis might suggest *N-1* stationary relations subject to finding WE variables, but a stricter reading of the ADF tests might imply $r=N-2$, the error correction models somewhere between *N-2* and *N-3* and the Johansen test *N-4*.

Following De Vanya and Walls (1999) we consider cointegration as a system that may relate to the more general case of equilibrium price targeting (Burke and Hunter, 2011). Cointegration across the system gives rise to a set of long-run relations that are tested jointly. Furthermore, the finding of weak exogeneity can distinguish between parallel pricing and aggressive price leadership (Hunter and Burke, 2007 and Kurita, 2008).

Irrespective of r, when the series are cointegrated there is a restricted long-run parameter matrix:

$$\mathbf{\Pi}=\boldsymbol{\alpha}\boldsymbol{\beta}'.$$

These can be identified in turn by setting $\boldsymbol{\alpha}'=[\mathbf{A}\ \mathbf{I_r}]$ or $\boldsymbol{\beta}'=[\mathbf{I_r}\ \mathbf{B}]$ and then we either find the $\boldsymbol{\beta}$ specifying the long-run relations, or we identify all the elements of $\boldsymbol{\alpha}$ which shows the adjustment to each cointegrating relation in the short-run. In the I(0) case r=N and identification of the long-run follows from $\boldsymbol{\alpha}$=diag[α_{ii}] and fixing the diagonal elements of $\boldsymbol{\beta}$ to unity, it then follows that each equation has an identified error correction or cointegration term. In this paper the 8 × 8 matrices express the long-run relationship in eight gasoline prices of eight different geographical areas in US. Therefore:

$$\Pi=\alpha\beta'=\begin{bmatrix}\alpha_{1,1} & 0 & \cdots & 0\\ 0 & \alpha_{1,1} & \cdots & 0\\ \vdots & \vdots & \ddots & \vdots\\ 0 & 0 & \cdots & \alpha_{N,N}\end{bmatrix}\begin{bmatrix}\beta_{.1}\\ \beta_{.12}\\ \vdots\\ \beta_{.N}\end{bmatrix}$$

where $\boldsymbol{\beta}_{.i}=[\beta_{1i}\quad \dots \quad 1 \quad \dots \quad \beta_{Ni}]$. The existence of cointegration in a VAR system specifies that the stochastic trends are integrated in the system, where in the N-dimensional I(1) cointegrated VAR with $\boldsymbol{r}$ cointegrating relations there are N-r common stochastic trends. In a VECM with N data series and $\boldsymbol{r}$ cointegrating vectors there can be up to $(N$-$r)$ weakly exogenous variables (Johansen, 1995). In this study for the multivariate case with eight price series and $r=4$ the corresponding unrestricted model is specified as follows:

$$\begin{bmatrix}\Delta p_{1t}\\ \vdots\\ \Delta p_{8t}\end{bmatrix}=\begin{bmatrix}\alpha_{1,1} & \cdots & \alpha_{1,4}\\ \vdots & \ddots & \vdots\\ \alpha_{8,1} & \cdots & \alpha_{8,4}\end{bmatrix}\begin{bmatrix}1 & \cdots & \beta_{7,1} & \beta_{8,1}\\ \vdots & \ddots & \vdots & \vdots\\ \beta_{1,4} & \cdots & \beta_{7,4} & \beta_{8,4}\end{bmatrix}\begin{bmatrix}p_{1t-1}\\ \vdots\\ p_{8t-1}\end{bmatrix}+\begin{bmatrix}\gamma_{1,1} & \cdots & \gamma_{1,8}\\ \vdots & \ddots & \vdots\\ \gamma_{8,1} & \cdots & \gamma_{8,8}\end{bmatrix}\begin{bmatrix}\Delta p_{1t-1}\\ \vdots\\ \Delta p_{8t-1}\end{bmatrix}+\begin{bmatrix}\varepsilon_{1t}\\ \vdots\\ \varepsilon_{8t}\end{bmatrix}.$$

Hence for studying the gasoline market structure and identifying the number of long-run relations, it is necessary to impose some further restrictions on the VAR model above. Following Forni (2004), De Vanya and Walls (1999), and Hendry and Juselius (2001) we analysed the competitiveness of the gasoline market and a broad market definition by investigating a potential single trend among the series.

Following, Johansen (1992), Hunter and Simpson (1995), Bauwens and Hunter (2000), and Burke and Hunter (2012) weak exogeneity in the long-run has been identified by imposing restriction on $\alpha_{ij}= 0$, for i=1, … ,8 and j=1, …, 4. For weak exogeneity and long-run exclusion there are r restrictions on $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ for each variable excluded while for strict exogeneity, there are *2r* restrictions on $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ for each variable excluded from the loadings and the cointegrating relations. The restrictions are tested by further likelihood ratio test statistics, which conditional on *r* are distributed $\chi^2(i)$ with *i=r* and *2r* respectively**.**

Moreover long-run exclusion (Juselius, 1995) can be tested by imposing restrictions in $\beta_{ji} =0$, for j=1, ... , 8 and i=1, …, 4 . A further component of the process used to identify is to select the most appropriate normalisation of the data by imposing the restriction below:

$$\beta_{ii}=1, \text{ for } i=1, \ldots, 4$$

$$\beta_{ij}= 0 \text{ , for } \begin{cases} i = 1, \ldots, 4 \\ j = 1, \ldots, 4 \text{ .} \\ i \neq j \end{cases}$$

It is important not to normalise on a variable that is weakly exogenous and/or long-run excluded. Parallel pricing is tested by imposing restrictions of the form in the case of the first long-run relation as $\boldsymbol{\beta}_{.1}' = \begin{bmatrix} 1 & 0 & \cdots & -1 \end{bmatrix}$ and subsequently for $\boldsymbol{\beta}_{.i}'$ the i^{th} term is set to unity and all the other up to N^{th} can be set to zero to confirm a long-run correspondence between the price series.

In Table 3, tests of cointegration are derived from the VAR model and the result of imposed restrictions on $\boldsymbol{\alpha}$ or $\boldsymbol{\beta}$ or both $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ at the same time are presented accordingly. The sample includes 901 observation and the results relate to weak exogeneity tests, long-run exclusion test and strict exogeneity with *k=20* lags in the estimation. The first block of results in Table 3 relate to a weak exogeneity test conditional on *r=4* and from the p-values it can be determined that the log price of the GC, LA, EC and MW are potentially WE for $\boldsymbol{\beta}$. However, there can be no more than *N-r=4* WE variables. Ordering by the weak exogeneity test suggests the most likely WE variables are the Gulf Coast and Lower Atlantic. This suggests that the gasoline prices of GC and LA determines the other regions prices, consistent with Burke and Hunter (2012).

However as the GC price changes will directly affect the other region's prices, then everything can be conditioned on the GC. Following Juselius (1995) the next section of the table tests long-run exclusion. These results are strongly significant in all regions indicating the appropriateness of the rank condition and the likely robustness of propositions on the cointegrating vectors. We could order the system using the test on long-run exclusion[11] as it is not appropriate to normalise on a variable that may be the long-run excluded (LE) as is explained in Boswijk(1996). Here finding a variable is not LE implies that it may interact with all the other variable in the long-run and that variable must be present in at least one cointegrating vector.

Next in Table 3 following the normalization and weak exogeneity, the system is conditioned on LA log prices as the variable that is most likely to be weakly exogenous. Next, testing for the normalization and weak exogeneity conditioned on the log price of GC. The same values of the χ^2 test apply from the normalization and weak exogeneity test and the weak exogeneity test confirms that the normalisation is innocuous.[12]

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[11] Long-run exclusion tests in estimation exclude potentially redundant variables from the cointegration space.

[12] With *r<N-1*, identification may arise in a range of different ways including via what is termed the normalisation rule in Boswijk (1996), but given the result that none of the variables can be LE from the long-run relations gives a further basis for long-run arbitrage not to apply.

[Table 3 goes here]

In the final section of Table 3 the system might be reordered based on the imposition of strict exogeneity by testing whether a variable can be LE and WE at the same time. However, the conclusion in Hunter and Simpson (1996) does not seem appropriate as a mechanism to reorder and condition the system as none of the series are strictly exogenous.

The result indicates that long-run arbitrage may be driven by GC and LA and this is consistent with the findings of Burke and Hunter (2012). To this end regional gasoline pricing may not be consistent with a fully functioning gasoline market in the US. There may be geographical or structural reasons for this to occur. To further investigate market structure it would be useful to study US company gasoline prices and search for WE price series with that data (Burke and Hunter, 2011). A difficulty associated with analysing company price series, is that they are volatile and these data sets are not as large.[13]

6. Conclusion

For non-stationary variables, the Johansen methodology of cointegration and exogeneity testing appears an appropriate approach to investigate market performance. Here a range of tests for the integrated nature of the market have been employed to analyse US gasoline prices. The empirical findings indicate that gasoline prices of different regions are cointegrated and this suggests that the market may not be distinct. Forni (2004) found with a very modest regional data set for Italian milk prices that stationarity tests such as that of Dickey and Fuller (1979) can provide an effective way of defining the dimensions of a market, especially when there is a limit to the number of

[13] Company data were analysed, but these results are preliminary. The findings suggest $r=N-2$, but with a smaller sample and volatile price data they are viewed as tentative and for reasons of space and consistency with the above discussion they are not reported here.

time series observations. One problem with that approach is that the long-run restrictions are also binding on the short-run. For this reason preference is given to the test based on the ECM. Furthermore, the ECM as part of an *N* dimensioned system with *N* error correction terms can be coherently defined (Boswijk, 1992). Further from the findings in Kremers, Ericsson and Dolado (1992) the test based on the error correction term in a dynamic model should be more powerful than the ADF test.

However, the single equation methods do not bind the reduced rank restriction across the whole set of prices. This suggests that when there is a large data set available that the VECM is to be preferred. In particular in the presence of relatively strong ARCH behaviour the simulations presented in Rahbek et al (2002) imply that testing may only be reliable with data sets in the range 600-1000 observations. Here even though there is some evidence of ARCH we feel confident in an analysis based on a sample of 901 observations with a clear finding that the cointegrating rank (r) is less than N-1. This is also not inconsistent with a strict analysis of the single equation results. Unfortunately these findings combined with the results on long-run exclusion call into question the existence of long-run arbitrage pricing across the eight US regions investigated here.

Hunter and Burke (2007), and Kurita (2008) suggest that even where there are N-1 cointegrating relations that this may be inconsistent with an efficient market when one of the prices is weakly exogenous. In that case a single variable drives the stochastic trend and as a result the long-run can be appropriately conditioned on that price; the most likely candidates being the LA and the GC. Furthermore, the tests here reveal that at least four of the regional prices can be considered weakly exogenous casting further doubt on the extent to which prices are responding to each other in the long-run. To this end the observed market behaviour in the long-run could be due to the geographical conditions and this may be further reflected in the ownership of regional refinery capacity.

Considering the empirical results we are suggesting a change in the regulation of the gasoline market to enhance competition. This could relate to tax breaks to extend the refinery and distribution capacity of smaller firms. A similar conclusion to Forni (2004) arises as the failure to find a “Broad Market” in the long-run suggests that the anti-trust authorities resist further concentration in the industry via merger or acquisition. The availability and accessibility of market information to the consumer could also affect price responsiveness in this market. Similar conclusions may also be pertinent to other countries such as the UK.

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Table 1- Summary of ADF tests, ECM test of regional price proportion. (With intercept and no trend)

Log price differential (q)[14]	ADF (q)/ OLS t-statistic	ECM (q)/ OLS t-statistic
$P_{NE\text{-}MW}$ (25)	**-3.81 ***	-14.48 **| P_{MW}
$P_{MW\text{-}CA}$ (25)	**-4.93 ***	-8.70 **| P_{CA}
$P_{MW\text{-}EC}$ (25)	**-4.72 ***	-10.15 **| P_{EC}
$P_{LA\text{-}GC}$ (23)	-2.22	-5.63 **| P_{GC}
$P_{RM\text{-}WC}$ (16)	**-5.81 ***	-6.62**| P_{WC}
$P_{MW\text{-}GC}$ (20)	**-3.36***	-8.46 **| P_{GC}
$P_{GC\text{-}RM}$ (16)	**-5.21***	-1.22| P_{RM}
$P_{GC\text{-}WC}$ (20)	**-3.78****	-2.65 ** | P_{WC}
$P_{MW\text{-}RM}$ (24)	**-4.43***	-3.76 **| P_{RM}

Note**:** Critical value at 1% is -3.44, at 5% is -2.87 computed in Professionl Oxmetrics Professional (Doornik and Hendry, 2009). ***** Significant at the 95% confidence level and ** significant at the 99% confidence level

Table2: Johansen trace test for cointegratione

H_0 : rank ≤	Trace test	P-value
rank =0	226.673	[0.0000] **
rank =1	159.485	[0.0001] **
rank =2	115.337	[0.0012] **
rank =3	76.017	[0.0147] *
rank =4	48.471	[0.0437] *
rank =5	28.207	[0.0754]
rank =6	11.631	[0.1755]
rank =7	1.1499	[0.2836]

Note: * significant at the 5% level and ** significant at the 1% level.

[14] q is the lag order of each series which had been selected by using same process as the previous study via inspection of the correlogram.

Table 3- Test of cointegration, WE, LE, SE and Parallel Pricing of US Gasoline Price 1993-2010

Hypothesis		Null ($r \leq 4$)	Statistics [p-value]
(WE)| r=4	P_{CA}	$\alpha_{1i}= 0$, for i=1, …, 4	$\chi^2(4) = 15.266$ [0.0042] **
	P_{EC}	$\alpha_{2i}= 0$, for i=1, …, 4	$\chi^2(4) = 7.0753$ [0.1320]
	P_{GC}	$\alpha_{3i}= 0$, for i=1, …, 4	$\chi^2(4) =1.7056$ [0.7897]
	P_{LA}	$\alpha_{4i}= 0$, for i=1, …, 4	$\chi^2(4) =3.5660$ [0.4679]
	P_{MW}	$\alpha_{5i}= 0$, for i=1, …, 4	$\chi^2(4) =8.7465$ [0.0678]
	P_{NE}	$\alpha_{6i}= 0$, for i=1, …, 4	$\chi^2(4) =10.803$ [0.0289] *
	P_{RM}	$\alpha_{7i}= 0$, for i=1, …, 4	$\chi^2(4) =32.663$ [0.0000] **
	P_{WC}	$\alpha_{8i}= 0$, for i=1, …, 4	$\chi^2(4) = 19.215$ [0.0007] **
(LE)| r=4	P_{CA}	$\beta_{j1}= 0$, for j=1, …, 4	$\chi^2(4) = 18.758$ [0.0009] **
	P_{EC}	$\beta_{j2}= 0$, for j=1, …, 4	$\chi^2(4) = 10.883$ [0.0279] *
	P_{GC}	$\beta_{j3}= 0$, for j=1, …, 4	$\chi^2(4) = 15.663$ [0.0035] **
	P_{LA}	$\beta_{j4}= 0$, for j=1, …, 4	$\chi^2(4) = 8.9190$ [0.0632]
	P_{MW}	$\beta_{j5}= 0$, for j=1, …, 4	$\chi^2(4) = 24.574$ [0.0001] **
	P_{NE}	$\beta_{j6}= 0$, for j=1, …, 4	$\chi^2(4) = 0.89249$ [0.9256]
	P_{RM}	$\beta_{j7}= 0$, for j=1, …, 4	$\chi^2(4) = 36.858$ [0.0000] **
	P_{WC}	$\beta_{j8}= 0$, for j=1, …, 4	$\chi^2(4) = 26.188$ [0.0000] **
Normalization (N) + (WE) P_{GC} | r=4		$\beta_{ii}=1$, for i=1, …, 4 $\beta_{ij}= 0$, for $\begin{cases} i = 1, \dots, 4 \\ j = 1, \dots, 4 \\ i \neq j \end{cases}$ $\alpha_{3i}= 0$, for i=1, …, 4	$\chi^2(4) = 1.7056$[0.7897]
Normalization (N) + (WE) P_{LA} | r=4		$\beta_{ii}=1$, for i=1, …, 4 $\beta_{ij}= 0$, for $\begin{cases} i = 1, \dots, 4 \\ j = 1, \dots, 4 \\ i \neq j \end{cases}$ $\alpha_{4i}= 0$, for i=1, …, 4	$\chi^2(4) = 3.5698$[0.4673]
SE = (LE) + (WE)| r=4	P_{CA}	$\alpha_{1i}= 0$, for i=1, …, 4 $\beta_{j1}= 0$, for j=1, …, 4	$\chi^2(8) = 53.103$ [0.0000] **
	P_{EC}	$\alpha_{2i}= 0$, for i=1, …, 4 $\beta_{j2}= 0$, for j=1, …, 4	$\chi^2(8) = 35.607$ [0.0012] **
	P_{GC}	$\alpha_{3i}= 0$, for i=1, …, 4 $\beta_{j3}= 0$, for j=1, …, 4	$\chi^2(8) =44.477$ [0.0000] **
	P_{LA}	$\alpha_{4i}= 0$, for i=1, …, 4 $\beta_{j4}= 0$, for j=1, …, 4	$\chi^2(8) =30.611$ [0.0063] **
	P_{MW}	$\alpha_{5i}= 0$, for i=1, …, 4 $\beta_{j5}= 0$, for j=1, …, 4	$\chi^2(8) =49.294$ [0.0000] **
	P_{NE}	$\alpha_{6i}= 0$, for i=1, …, 4 $\beta_{j6}= 0$, for j=1, …, 4	$\chi^2(8) =33.533$ [0.0024] **
	P_{RM}	$\alpha_{7i}= 0$, for i=1, …, 4 $\beta_{j7}= 0$, for j=1, …, 4	$\chi^2(8) =70.029$ [0.0000] **
	P_{WC}	$\alpha_{8i}= 0$, for i=1, …, 4 $\beta_{j8}= 0$, for j=1, …, 4	$\chi^2(8) = 62.893$ [0.0000] **

Note: Weak Exogeneity (WE), Long-run Exclusion (LE), and Strict Exogeneity (SE). * significant at the 5% level and ** significant at the 1% level.