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 error correction model. After examining the stationarity and cointegration properties of the weekly gasoline prices across eight different regions of the US we consider long-run price leadership and parallel pricing in the framework of the cointegrated vector auto-regression (VAR). In contrast with Kurita (2008) the complete market is considered using data on 901 weekly gasoline prices for the US. The finding 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 cointegrating rank based on a broader range of prices implies two further common trends. One can be associated with at least one weakly exogenous variable and the others to cointegrating exogeneity and the tests of exogeneity suggest that the Gulf Coast, Mid West and West Coast gasoline prices are forcing rather than responding to the other regions in the long-run.

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 key process in the production of gasoline takes the product from the oil field to the gas station pump in four steps: 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 a vital revenue stream 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. However, a persistent price differential suggests discrimination and identifies the possibility of 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 discrepancies and improve market structure.

In section I we introduced the gasoline market and econometric model used in the analysis. In section II we review essential literature on stationarity, error correction, and exogeneity. Section III identifies the data for the empirical analysis. Section IV we review 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 offer our conclusions.

2. Review of the Essential Literature

It would be natural to assume that competitive behaviour ought to be reflected in price movement. La Cour and Møllgaard (2002) focused on the appropriateness of a legal definition that might be used to define anti-trust behaviour. The focus is on the extent to which a firm may be able to operate independently of its competitors. While Forni (2004) has approached the problem from a slightly different manner in terms of categorising a market as broad or narrow and thus defining the extent of a market. Where the breadth of a market within a region is linked to the degree of price responsiveness across a physical entity or the degree to which there is interaction in firms prices. If prices are sticky, then it should be possible without the limitations of law or physical borders within a market to arbitrage the product and this links to the related concept of the law of one price.

In particular, the mechanism by which firms react to pricing decisions has been considered for a considerable time, take for example Markham (1952) or Stigler (1947). Stigler (1947) provides a rationale for price responses to be slow. More recently attention has been paid to breaking down the nature of these price responses both theoretically and empirically. From a theoretical perspective, Buccirossi (2006) considers whether competitive behaviour requires firms to adopt parallel pricing. This corresponds to what has been called the "law of one price" that implies that in any market 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, arbitrage will give rise to price correction and when the market is efficient it might be anticipated that such adjustment should be fast.

Much of the earlier empirical literature is well summarized in a report for the United Kingdom Office of Fair Trading (OFT) by LECg (1999) where the focus was on price correlation and causality. The suggestion being that correlation was an indication of collusion (see Maunder (1972)). Further consideration was made of endogeneity by Slade (1986) again an indication that certain firms pricing decisions were driving the market. More recently, the distinction has been made between the long-run and the short-run. One reason might be that it may be easier to encourage a committee or jury that irregularity in pricing in the long-run is serious enough to lead to legal action as harm to the consumer. Consumer harm has been defined in terms of consumer detriment (see Hunter et al (2001)). Detriment can be measured either directly or indirectly, the direct measures relate to the extent of legal activity and complaint in terms of the delivery and quality of the product delivered. However, monitoring whether the consumer is damaged by corporate inactivity is not as straightforward to determine. Indirect, measures require some notion of cost and this often relates to accounting information only available on an annual basis (Hunter et al 2001) and this is a reason for Forni (2004) emphasising the use of price information as compared with calculating measures of the residual demand curve. Furthermore, short-run behaviour that does not have a persistent affect and where harm may balance against occasional benefits may be less harmful to the consumer.

Forni (2004) suggested a typology of tests across market segments to categorize the nature of the market. As mentioned above, Forni (2004) considers a number of mechanisms to determine whether a market is competitive, but argues that finding the log price proportion to be stationary is an effective and efficient approach to determine what he terms a “broad market”. To this end Forni (2004) has analysed the extent to which the market for milk across Italy can be viewed as being competitive. If the market is not competitive, then prices are not adjusting in a long-run sense or there are regional anomalies and then there are strong anti-trust reasons to limit further concentration in the industry. To this end the market is seen as broad when milk pricing in one part of Italy is reflected in the pricing decisions of all the other regions and breadth is measured by the extent to which inter-regional price proportions are stationary. Forni (2004) emphasises that when the market can be arbitraged, then it is competitive and this relates in the long-run to parallel pricing (Buccirossi (2006)). When compared with analysing a system of prices via a VAR, Forni (2004) argues that the method he applies does not require the log price series to be integrated of order one (I(1)) and the test considers jointly stationarity and parallel pricing. Forni (2004) finds limited evidence for competitive behaviour and a market that is broad. The analysis is extended by Giulietti et al (2010), and Hunter and Tabaghdehi (2013) from the univariate to the panel context to analyse market definition in relation to energy prices in UK and US respectively. While Hosken and Taylor (2004) provide an extended reply to Forni (2004) by use of an example using gasoline prices in the US to evidence some of limitations with the univariate approach.

Whether the analysis relates to the gasoline market or the market for milk, the application of stationarity tests on time series or panel data is used to 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. However, this is bought at a price as ordinarily the conventional stationarity tests bind the same structure to the time series dynamics and as a result impose the long-run arbitrage restriction on the short-run.[1] The VAR in error correction form can be used to address the issue of exogeneity and the interrelatedness of prices in the long-run.

La Cour and Møllgaard (2002) suggested that it is possible to test the proposition commonly used by the European Court of insufficient response to competitors by embedding the testing within a dynamic system. While Hendry and Juselius (2001) suggest that parallel pricing in the gas market, can be analysed in the context of a bivariate VAR conditioned on oil prices. The error correction model (ECM) is one mechanism used to analyze the price adjustment process in a dynamic context and by considering prices in a system the simultaneous finding of parallel pricing can be tested. It is shown by Hunter and Burke (2007) that there are *N-1* interrelated prices that yield stationary proportional prices. This is termed long-run equilibrium price targeting (LEPT) in Burke and Hunter (2011). This permits the law of one price to be analysed in the long-run and prices to be seen to be interrelated. La Cour and Møllgaard (2002) suggested that the argument over the capacity to determine competitiveness in a system relates to the extent to which it is possible to determine whether there is price reaction across all segments of the market place and in addition to the capacity to test exogeneity the VAR permits analysis of price interaction at the level of the market.

[1] Pesaran and Pesaran (2010) implement the ADF test subject to further stationary variables and then simulate the critical values subject to these additional terms, implying the test may be handled without imposing the short-run restriction.

However, finding *N-1* long-run relations that all obey parallel pricing is necessary, but not sufficient for an efficient market. This case then corresponds to LEPT (Burke and Hunter (2011)), which corresponds to all firms following a target price that depends on all the prices across the market. In the case where there are *N-1* long-run relations, then there is a single stochastic trend and this encapsulates all the demand and supply shocks in the market. In essence the single stochastic trend is a random walk and so all past information on prices is embodied in this.

Attention also needs to be paid to long-run exogeneity and the dependencies that this engenders. The capacity to investigate long-run causality and conditioning is a key advantage to estimation of a system over the approach of Forni (2004). The study by LecG (1999) points out that the two primary mechanisms to analyse price relations are defined in terms of correlation and causality. It is this that permits us to distinguish between long-run relations that suggest agents respond to competitors and the case where a particular firm or sector is not responding to the others. A further feature of the systems method is that it permits a distinction between parallel pricing and the case where one or more firms are not impacted by the price shocks that affect the other firms.

Analysing pricing properties may be effective in testing for "market definition" when the persistence in volatility is reduced by consideration of the price proportions. However, when volatility is quite persistent then in a system the largest eigen value (spectral radius) of the ARCH[2] polynomial exceeds .85, because there is evidence (Rahbek et al (2002)) that the Johansen test statistic does not converge at the rate anticipated to the asymptotic distribution unless the sample increases to between 600 and 1000 observations.[3]

[2] The notion of Autoregressive Conditional Heteroscedasticity (ARCH) relates to Engle (1982) and it was viewed in initial studies as a feature of nominal data as it was first applied to price data for the UK.

[3] Burke and Hunter (2014) have found similar results for a trivariate model with spectral radius in excess of .85 and T=680 observations.

Thus with a more extensive data set, then the tests are likely to be better sized in the presence of volatility. This would suggest that the cointegration methodology of Johansen (1995) is appropriate to test empirically the definition of the market and the nature of integration of the price series. This should make it possible to determine competitive behaviour in the market from the long-run decomposition of prices. Consequently, the conditional ECM and VAR can be used to test cointegration, analyse the long-run relations and consider the potential for arbitrage correction in a market.

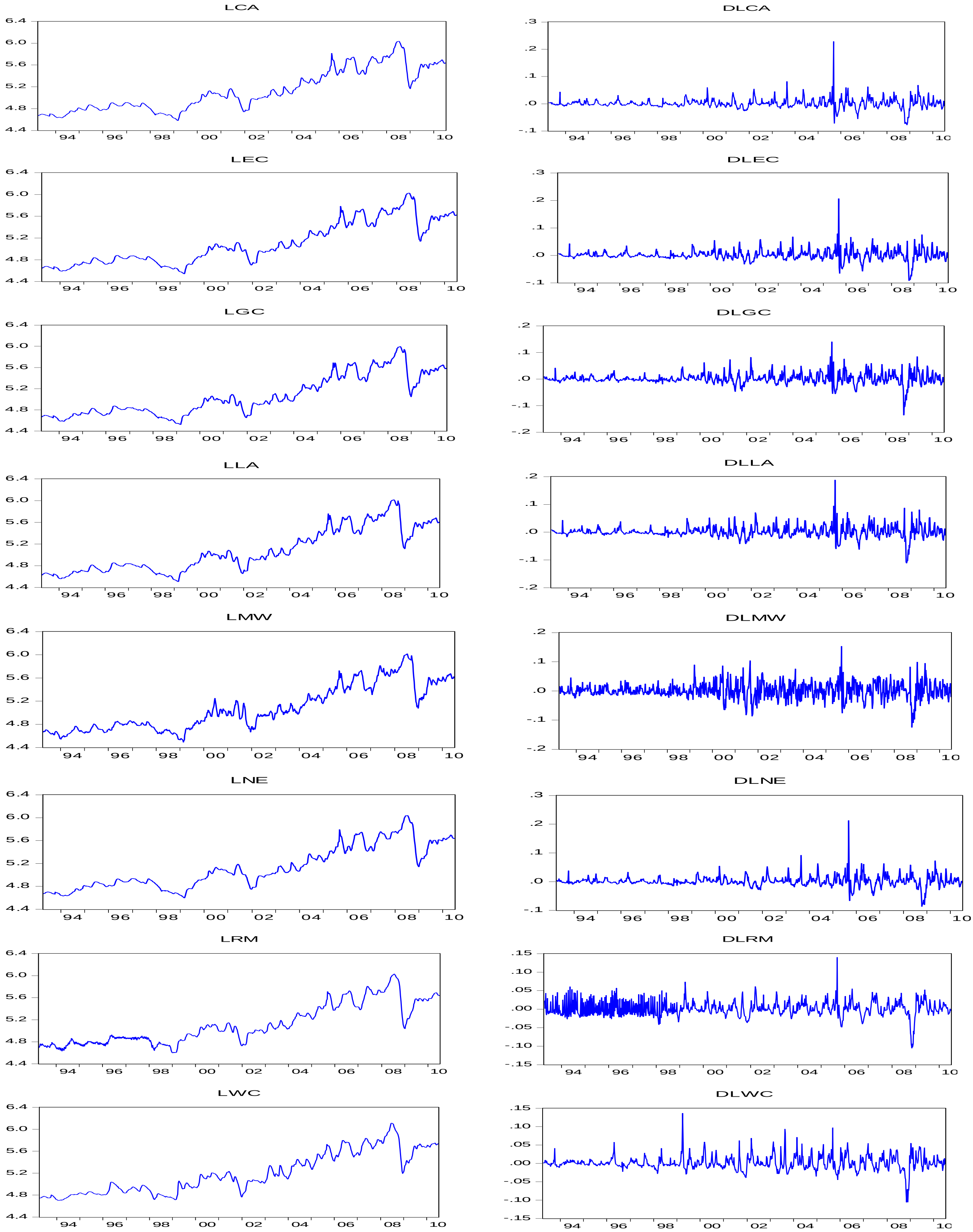
In financial markets it is suggested by Fama (1998) that specific patterns of pricing behaviour in the market can give rise to profitable opportunities from arbitrage that cannot survive for long and over time they will dissipate as others seek them out. In the energy market there is capacity to store the product and this should make it suitable for price arbitrage and hedging. When considering the price of gasoline should the market be efficient, then it should be possible to observe opportunities for location arbitrage. However, in a country the size of the US, this may be limited by the extent to which there may regional or physical barriers. Consequently to tackle arbitrage opportunities in a market-oriented industry to address market power there may need to be some form of regulation (Küpper and Willems, 2010).

3 Time-series properties of the data

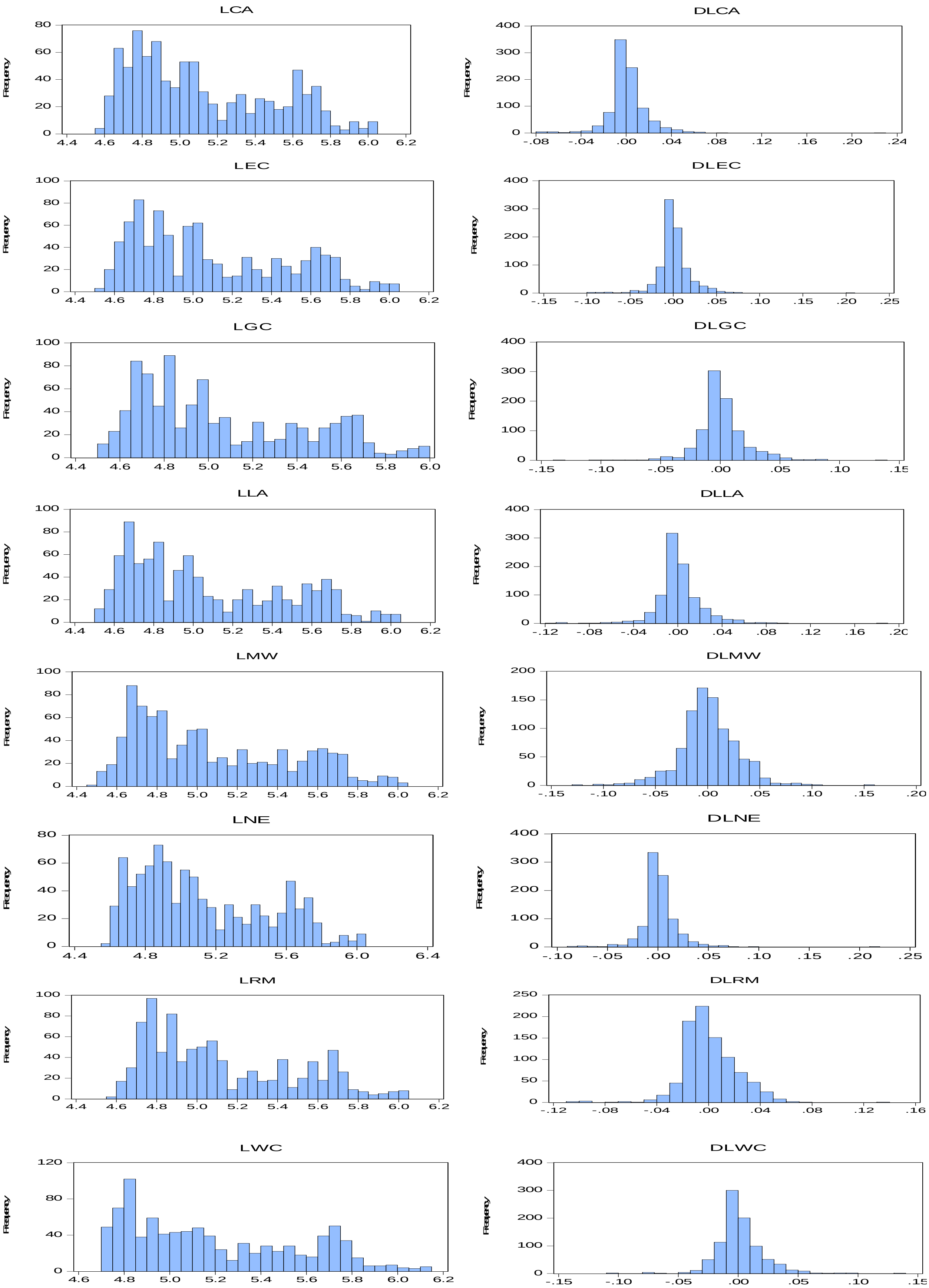
In this section we consider the time series properties of a data set consisting of weekly gasoline prices across eight different regions in the US (West Coast (WC), Central Atlantic (CA), East Coast (EC), Gulf Coast (GC), Lower Atlantic (LA), Midwest (MW), New England (NE), Rocky Mountains (RM)) from May 1993 to May 2010.[4] Considering regional gasoline

[4] 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



infrastructure across the US we test cointegration on eight different regions. 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 and the frequency distribution of the differences suggest the series are closer to normality.

It is also of note that the data are volatile and that there are some large movements. It might be considered that the largest shocks relate to the financial markets crisis in 2008, but that is not the case. As can be observed from the time series plots and the findings in Kurita (2008) log gasoline prices are clearly difference stationary. Hunter and Tabaghdehi (2013) have applied a broad range of stationarity test on the log levels and log differentials of the data analysed here and they are also unable to reject the notion that these series are difference stationary (I(1)). As the largest movements relate to the earlier sample excluding the recent crisis, then it is anticipated the strongly persistent autoregressive behaviour during 2008 is indicative of the powerful movements that can be observed with series following stochastic trends. Burke and Hunter (2012) show that similar data can be readily calibrated and simulated as random walks. It should be noted that in the latter case the simulated data do not have any structural break, but vary in the same way as the actual data for the shorter sample used by Kurita.

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 using regional data across the US mainland, this excludes Alaska that has a significant physical border, Canada.

De Vany and Walls (1999), Hendry and Juselius (2001), and Forni (2004) suggest that finding cointegration between two prices is indicative of an efficient market. Forni (2004) analyses all possible price combinations to determine whether the market is efficient basing the conclusion on a typology of the findings on stationarity tests under both the null of stationarity and non-stationarity. However this is a single equation approach that is not able to bind the findings to a test of all market segments.

Following Hunter and Burke (2007) it is suggested that arbitrage implies that there are *(N-1)* cointegrating relations derived from *N* price series and this is consistent with a broad market. While a narrow market implies fewer than *(N-1)* cointegrating relations. However, the finding of *(N-1)* long-run relations does not negate the possibility that the market is segmented in the sense that some prices do not respond to the other prices in the market. This may arise when the form of long-run causality related to cointegrating exogeneity (Hunter (1990)) is observed or detect that one or more prices is weakly exogenous (WE) for all the cointegrating vectors (Johansen, 1992).

Forni (2004) suggest that when comparison is made between the prices of two regions then competitive behaviour is consistent with parallel pricing when in testing price proportions it is found that they are stationary. However, such an approach has merit when the data is limited by the extent of the time series. 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 using 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 idea is generalized by Hunter and Burke (2007) to a multi-price framework. We pay particular attention to the role of the common trend and exogeneity in explaining the competitive structure. Here there is a large time series sample and this is useful as the series are volatile.

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

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

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 Now $\mu_0 = 0$ when the prices in different regions are identical, and $\mu_0 \neq 0$ if there is a fixed transportation and other characteristics related to different regions. However with a perfectly integrated market the price reflects all available information and traders ought not to benefit consistently from arbitrage opportunities. Equation [1] is a cointegrating regression where b explains the nature of the relation between the regional prices. The hypothesis related to parallel pricing implies that b=*1* is the key hypothesis to be tested as when *b=1,* regional prices respond in proportion to each other and this conforms with the law of one price. Though the observed value may differ from 1 by an arbitrary constant(*c*) where $| b -1|\leq c$. In the case of perfect integration *c* is close to zero.

In contrast with the tests of stationarity of price proportions, the linear combination of two non-stationary series p_{at} and p_{bt} can be transformed to stationarity (Engle and Granger (1987)) when:

$$\eta_t = p_{at} - bp_{bt} \sim I(0). \quad (2)$$

This 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 relation between p_{at} and p_{bt} otherwise the relation is nonsense. Consequently for the price of any homogeneous good in an identical market a cointegrating relation is required as arbitrage should remove 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 law of one price in a VAR and the potential for price leadership. 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 regions.

Here, the ECM provides one method to investigate the nature of adjustment across prices to determine long-run equilibrium, see Patterson (2000). We investigate long-run equilibrium in the US gasoline market using the error correction model. This case in particular is termed arbitrage correction by Burke and Hunter (2012). The hypothesis underlying this argument relates to the possibility that a sequence of regional gasoline prices that deviate from equilibrium give rise to an arbitrage opportunity that is correcting in the long-run. If there are *N-1* arbitrage correction terms across *N* markets, then this also relates to LEPT (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 relations 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.

To further investigate the short-run dynamics of the relations in gasoline prices of different regions in the US we employ a vector error correction model (VECM). For example, Bachmeier and Griffin (2006) found that the prices of crude oil in different geographical regions of the world are cointegrated. While De Vany and Walls (1999) using a VECM, 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 residual:

$$\hat{\eta}_t = e_t = p_{at} - \hat{\mu}_o - \hat{b} p_{bt}. \quad (3)$$

We may test whether these series are stationary by applying the Dickey-Fuller test to these residuals and this relates to the following dynamic model:

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

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

where: $\epsilon_t = \gamma_0 \Delta p_{bt} + v_t$ then $b = b_0$ and $\gamma_0 = b - b_0$. It is also possible to have cointegration as a result of $\gamma < 0$, but this may not be consistent with efficiency as $\mu_o \neq 0$ and $b \neq 1$. In the long-run when the prices are set to their long-run average values $p_{at} = \check{p}_{at}$, $p_{bt} = \check{p}_{bt}$, then:

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

Where the μ_0 and b are long-run parameters and for efficiency in the market we require $\mu_0=0$, $\beta=1$. Therefore:

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

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 term. 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

5 ***γ %*** of the disequilibrium at time ***t-1*** is removed in period ***t***.

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 presented in Table 1.[7] Acceptance of the alternative hypothesis underlining the ADF tests implies 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 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 individual significance of the error correction term. In all but one

[6] $\boldsymbol{\gamma} > 0$ 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).

case the error correction terms are significant, this one exception may arise due to a lack of cointegration, weak exogeneity,[8] or that 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 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 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 arise as a result of inefficiency or the RM model may not contain an error correction term as this price is WE for the long-run relation. In the latter case 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.

In further investigating the system we follow Boswijk (1992), Hunter and Simpson (1996), and Bauwens and Hunter (2000) and apply restrictions on $\boldsymbol{\alpha}$, $\boldsymbol{\beta}$ (dimensioned N×r), and $\boldsymbol{\alpha}$ as well as $\boldsymbol{\beta}$ to study the exogeneity structure of the data and identify potentially WE variables.

[8] I a single equation context one may observe more WE variables than can arise when the rank restriction is applied across the system.

The following equation is the VECM parameterisation of the VAR:

$$\boldsymbol{\Gamma}\,(\mathbf{L})\ \Delta\ \mathbf{p_t} = \boldsymbol{\Pi}\mathbf{p_{t-1}} + \boldsymbol{\mu} + \boldsymbol{\varepsilon_t}.$$

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

$$H_1(r)\colon \boldsymbol{\Pi} = \alpha\beta'.$$

Using the Johansen trace test we identify the number of cointegrating vectors (r) and the number of common trends. The results on the Johansen trace test for eight regional gasoline prices in the US are presented in Table 2. We find that it is possible to accept the null hypothesis that there are $r=5$ cointegrating vectors for a test applied at the 5% level, the alternative is rejected as the test is not significant so $r>5$ cannot be accepted. This also implies that there are $N-r=3$ stochastic trends. This does not correspond with the results that arise when cointegration is tested based on the single equation methods. If $r<N-1$ there are more stochastic trends than might be anticipated by a single competitive market implying that LEPT cannot hold and the market is partitioned.

[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 three trends so there may be up to three different prices 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 three variables may also be WE implying that they are not affected by the long-run price behaviour in the other segments of the market.[9] 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.

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 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) in 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 market efficiency being imposed on the short-run relations, thus causing the arbitrage restriction to be imposed on the short-run parameters. Hence by estimating the VAR the short-run restriction does not bind 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 potentially inter-related market prices.

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

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 there are *r=N-3* long-run relations. However, a generous or more careful interpretation of the results derived from the single equation approach might suggest *N-1* stationary relations subject to finding of a WE variable. A stricter reading of the ADF tests might also suggest *r=N-2*, the error correction models somewhere between *N-2* and *N-3* when compared with the Johansen test where it is *N-3*.

Following De Vany and Walls (1999) we consider cointegration as a system and that may relate to the more general case of LEPT (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 $\boldsymbol{\alpha}$ that gives rise to adjustment to each cointegrating relation in the short-run. Let the *i*th column vector of $\boldsymbol{\beta}$ be denoted $\boldsymbol{\beta}_{.\mathbf{i}}$. Subject to a normalisation on the ith element, then $\boldsymbol{\beta}_{.i}=[\beta_{1i} \quad \dots \quad 1 \quad \dots \quad \beta_{Ni}]'$. The existence of cointegration in a VAR system implies that the stochastic trends are combined as *r* stationary linear combinations; there are *N-r* of these trends and this may give rise to no more than *N-r* weakly exogenous variables (Johansen (1995)). In this study there are eight price series and *r=N-3* the corresponding unrestricted model is specified as follows:

$$\begin{bmatrix}\Delta p_{1t}\\ \vdots \\ \Delta p_{8t}\end{bmatrix}=\begin{bmatrix}\alpha_{1,1} & \cdots & \alpha_{1,5}\\ \vdots & \ddots & \vdots \\ \alpha_{8,1} & \cdots & \alpha_{8,5}\end{bmatrix}\begin{bmatrix}1 & \cdots & \beta_{7,1} & \beta_{8,1}\\ \vdots & \ddots & \vdots & \vdots \\ \beta_{1,5} & \cdots & \beta_{7,5} & \beta_{8,5}\end{bmatrix}\begin{bmatrix}p_{1t-1}\\ \vdots \\ p_{8t-1}\end{bmatrix}+\begin{bmatrix}\gamma_{1,1}(L) & \cdots & \gamma_{1,8}(L)\\ \vdots & \ddots & \vdots \\ \gamma_{8,1}(L) & \cdots & \gamma_{8,8}(L)\end{bmatrix}\begin{bmatrix}\Delta p_{1t-1}\\ \vdots \\ \Delta p_{8t-1}\end{bmatrix}+\begin{bmatrix}\varepsilon_{1t}\\ \vdots \\ \varepsilon_{8t}\end{bmatrix}.$$

Where $\gamma_{i,j}(L)$ for $i, j = 1, \ldots, 8$ is a univariate polynomial of lag order *k*. Hence for studying the gasoline market structure and identifying the number of long-run relations, it is necessary to impose further restrictions on the VAR model.

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 a restriction on a each row vector $\boldsymbol{\alpha}_{i.}$=[0, 0, 0, 0, 0] from **α** in turn (for i=1, … ,8) and that excludes the long-run from each equation in the system. While long-run exclusion (Juselius, 1995) can be tested by imposing restrictions in $\boldsymbol{\beta}_{i.}$=[0, 0, 0, 0, 0] on each row vector of **β** in turn for i=1, ... , 8 and that excludes a variable from all the cointegrating vectors. Weak exogeneity and long-run exclusion impose r restrictions on **α** and **β** for the variable excluded. Strict exogeneity combines the weak exogeneity and long run exclusion restrictions for the i^{th} variable and imposes *2r* restrictions for each variable excluded from **α** and **β**. 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**.**

A further component of the process used to identify is to select the most appropriate normalisation of the data by imposing the restriction below:

$$\beta_{ii}=1, \text{ for } i=1, \ldots, 5$$

$$\beta_{ij}= 0 \text{ , for } \begin{cases} i = 1, \ldots, 5 \\ j = 1, \ldots, 5 \\ i \neq j \end{cases}.$$

Bauwens and Hunter (2000) suggest it is important not to normalise on a variable that is weakly exogenous and Boswijk (1996) suggests the same for long-run exclusion. For parallel pricing let the first column of $\boldsymbol{\beta}$ be tested by imposing restrictions of the form $\boldsymbol{\beta}_{.1} = \begin{bmatrix}1 & 0 & \cdots & -1\end{bmatrix}'$ 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 results related to the imposed restrictions on $\boldsymbol{\alpha}$ or $\boldsymbol{\beta}$ or both $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are presented accordingly. The sample includes 901 observation and the results relate to tests of weak exogeneity, long-run exclusion and strict exogeneity. There are $k=21$ lags in the VAR estimations. The first block of results in Table 3 relate to a weak exogeneity test conditional on $r=5$ and from the p-values it can be determined that the log price of the GC, the LA and the MW are potentially WE for $\boldsymbol{\beta}$. The joint test that all the N-$r=3$ variables are WE for β giving rise to 15 restriction is clearly rejected at the 5% level as the test, 38.227 has a p.value =[0.0008]. However, the null hypothesis cannot be rejected that the GC and the MW price series are WE for β as the test is 14.273 [0.1609], and similarly for the LA and the MW prices as the test is 15.280 [0.1222]. However, this does not hold for the GC and the LA prices. There are good reasons to order the system based on these tests as when the system is normalised this can be seen as a conditioning on the series most likely to be exogenous (Hunter and Simpson (1995)).

[Table 3 goes here]

Prior to further investigation of α, following Juselius (1995) the next section of Table 3 presents tests of long-run exclusion. These test results are significant for all regions indicating the appropriateness of the rank condition and the likely robustness of propositions on the cointegrating vectors. When there are long-run excluded variables, then it would be appropriate to order the system using this test prior to normalisation, because it is not appropriate to normalise on a variable that may be the long-run excluded (LE) and this can be viewed as one of the criteria devised by Boswijk (1996) to identify the long-run. In terms of the indication of anti-competitive behaviour finding a variable that is not LE implies that it may interact with all the other variables in the long-run as that variable must be present in at least one cointegrating vector. The final section in Table 3 relates to strict exogeneity and that combines the weak exogeneity with the long-run exclusion restriction. However, this will not be considered further as none of the price series appear to be strictly exogenous.

Next in Table 3 the system us normalised and conditioned in turn on the two price series that satisfy most readily the weak exogeneity tests that is for GC and LA. If the GC price is viewed as weakly exogenous for β, then the test is 5.1254 with normalisation restrictions. From the normalisation rule (Boswijk (1996)) following normalisation by a different price in each vector is here subject to that variable being neither LE nor WE, imposes *r-1* restrictions to exactly identify each cointegrating vector in β. However, the likelihood is unaltered as the restrictions are not binding and this representation gives rise to a long-run reduced form.

The last variable in the revised system is the GC price and α is restricted to impose weak exogeneity and this price will condition the long-run. Then based on subsequent investigation a further 21 restrictions are imposed on α and then β, and this gives rise to the matrices based on restricted coefficients:

$$\alpha = \begin{bmatrix} -0.252 & .222 & 0.0 & -0.03 & 0.025 \\ -0.189 & 0.223 & 0.0 & -0.018 & 0.021 \\ 0.0 & -0.182 & 0.025 & -0.198 & 0.077 \\ -0.109 & 0.0 & 0.0 & -0.045 & 0.028 \\ -0.187 & 0.314 & 0.0 & 0.0 & 0.014 \\ 0.0 & 0.0 & 0.038 & 0.0 & 0.0 \\ 0.0 & 0.0 & -0.014 & 0.0 & -0.053 \\ 0.0 & 0.0 & 0.0 & 0.0 & 0.0 \end{bmatrix} \text{ and}$$

$$\beta' = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 & 0 & -.371 & -.638 \\ 0 & 1 & 0 & 0 & 0 & 0 & -.429 & -.611 \\ 0 & 0 & 1 & 0 & 0 & -0.933 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 & -2.707 & .66 & 1.134 \\ 0 & 0 & 0 & 0 & 1 & 0 & -.39 & -.671 \end{bmatrix} = \begin{bmatrix} \beta_{.1}' \\ . \\ . \\ . \\ \beta_{.5}' \end{bmatrix}.$$

The system is ordered such that $x_{t-1}' = \begin{bmatrix} p_{CA} & p_{EC} & p_{RM} & p_{NE} & p_{LA} & p_{WC} & p_{MW} & p_{GC} \end{bmatrix}_{t-1}$. These restrictions gives rise to a likelihood ratio test of 20.936 and from the p.value=[.7453] it is not possible to reject them. In addition, the likelihood ratio test statistic related to further 21 over-identifying restrictions is computed as 15.8106 and as the p.value is [0.7802] then these are also not significant.

Following the imposition of the normalization rule the first *r* columns of **Π** reflect α and as a result of this it can be observed that on top of the WE restriction there is a block triangular section that is zero and this is consistent with the MW and the WC prices being CE for $\beta_{.1}$ and $\beta_{.2}$ (the first two cointegrating vectors).[10] The former is consistent with the joint test of WE that implies the prices for the MW and the GC might be considered WE for β. This would imply a system that could be conditioned on both these prices; implying two stochastic trends relating to the GC and the MW price. However, it was decided from inspection of α that as the MW price seemed to depend on $\beta_{.3}$ and $\beta_{.5}$ it was better to consider this as a CE for these two vectors. This seems pertinent as the MW price has a very similar coefficient in $\beta_{.1}$, $\beta_{.2}$ and $\beta_{.5}$. The long-run non-causality related with CE, seems to extend to $\beta_{.4}$. Similarly the block triangularity of **Π** implicit in the structure of the restricted α and β implies that the WC price is cointegrating exogenous for $\beta_{.1}$ and $\beta_{.2}$. However, in this case this is trivial as the terms related to the WC prices are excluded from these equations, but it can be observed from

[10] More strictly a row from α annihilates a column from β or $\pi_{ij}=\alpha_{.i}\,\beta_{j.}=0$ (Hunter and Simpson (1995)) or more generally the necessary condition for cointegrating exogeneity is $\mathbf{\Pi}_{21}=0$, an $N_2\times N_1$ sub-block of **Π** (Hunter (1990)).

α that $\beta_{.2}$ is the only vector that appears in the dynamic equation for the WC price and this implies that it is also CE for $\beta_{.4}$.

Considering the cointegrating vectors in turn, it follows from the restriction on $\beta_{.1}$ and $\beta_{.2}$ that the CA and EC price are driven by the same CE and WE variables the GC and the MW prices. These prices for all intents and purposes have similar coefficients and the two prices would appear to define a sub-block variant of LEPT (Burke and Hunter (2011)) that relates to $\beta_{.1}$ $\beta_{.2}$ and $\beta_{.5}$. Here we will focus on the equations explaining the prices for the CA and the EC that are being forced and as a result the GC and MW prices do not reflect the price related to these regions.

The form of $\beta_{.3}$ appears very close to what has been termed parallel pricing. It should be recalled that this only relates to LEPT when there are *N-1* similar vectors. Here this defines a partitioned market so the RM and WC prices are reflected in each other and none of the other prices are forcing this long-run relation so they share a common stochastic trend.

The cointegrating vector $\beta_{.4}$ relates to the NE price and this is driven by the GC, MW and WC prices. This appears to indicate that the NE price reflects information from across the US. This is given further support as the dynamic equation from the VAR is also impacted by the correction related to $\beta_{.1}$ that explains the CA price in the long-run and $\beta_{.5}$ that explains the LA price in the long-run. Based on the normalisation this may be viewed as the own vector, but the form of $\beta_{.4}$ seems less easy to understand given that it is anticipated that we observe parallel pricing and LEPT. However, this vector can be seen as a combination of three parity relations between the NE and WC price, the WC and MW price, and the WC and GC price. These are combinations that would arise from the tests of stationarity, but are rejected as stationary when it comes to the system. Furthermore, the NE prices are not being reflected in prices for the GC, MW and WC.

The long-run equation explained by $\beta_{.5}$ relates the LA to the MW and the GC prices. The GC price is again the driver and is not impacted by the LA price in the long-run. This is consistent with the investigation of the trivariate system results in Burke and Hunter (2012) and Kurita (2008), but over a shorter time frame that excludes the 2008 financial markets crisis as is the observation that the GC price is WE. However, as the MW price is dependent on $\beta_{.1}$ and $\beta_{.5}$, then it is appropriate to say that a linear relation between the LA and MW price is forced by the GC price and so in this case the MW and the LA are interdependent.

Hence, the exogenous variables appear to force the long-run equations, but in the case of the cointegrating exogenous variables the causality does not run the other way and for the weakly exogenous variable this is essentially a random walk. In the latter case the GC price is only impacted by shocks that impact this segment of the market and thus the history of demand and supply shocks that impact the price. Thus contrary to a competitive market, it is partitioned in the long-run.

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, but the reactivity of NE prices would suggest that this is not the case. To further investigate market structure it would be useful to study US company gasoline prices and search for WE price series with such data (Burke and Hunter, 2011). A difficulty associated with analysing company price series, is that they are volatile and that a similar historical data set does not seem to exist.[11]

[11] 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 as a compelling story still relates to the regional data.

6. Conclusion

For non-stationary variables, the Johansen methodology of cointegration and exogeneity testing appears an appropriate approach to investigate market performance. The empirical findings indicate that gasoline prices for 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 time series observations.

One problem with that approach is that the long-run restrictions are also binding on the short-run, this provides one reason why the test based on the ECM may be preferred. Furthermore, the ECM as part of an N dimensioned system with N error correction terms can be coherently defined (Boswijk, 1992). While Kremers, Ericsson and Dolado (1992) have shown that tests 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.

The single equation findings based on the ECM 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 the results 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 preferred model reveals that possibly three regional prices can be considered exogenous. It is derived here conditioned on the GC and this price does not react to other prices. If the long-run structure is further investigated it is suggested that the MW and the WC prices are CE for $\beta_{.1}$, $\beta_{.2}$ and $\beta_{.4}$ this implies that the prices are not responding to each other in the long-run.

The observed market behaviour in the long-run could be due to the geographical conditions or may be a reflection of the ownership of regional refinery capacity and their location across the US. 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. In this respect similar conclusions may also be relevant to countries such as the UK where significant concentration in refinery ownership has come under scrutiny especially following the fuel protests and related blockades of refineries in 2000.

The failure of the market for gasoline mirrors to some extent the conclusions of Forni (2004) and this implies anti-trust authorities resist further concentration in the industry via merger or acquisition. However, in this case the findings follow from the more subtle analysis related to the system of regional prices as compared with tests of stationarity that give rise to the conclusion the market may be competitive (Hunter and Tabaghdehi (2013)).

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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)[12]	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.

[12] q the lag order of each series has been selected by consideration of the maximum lag found via inspection of the correlogram of the individual logarithmic price series.

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

Hypothesis		Null ($r \leq 5$)	Statistics [p-value]
(WE)| r=5	P_{CA}	$\alpha_{1i}=0$, for i=1, …, 5	$\chi^2(5)$ = 18.872 [0.0020]**
	P_{EC}	$\alpha_{2i}=0$, for i=1, …, 5	$\chi^2(5)$ = 11.359 [0.0447]*
	P_{GC}	$\alpha_{3i}=0$, for i=1, …, 5	$\chi^2(5)$ = 5.1254 [0.4008]
	P_{LA}	$\alpha_{4i}=0$, for i=1, …, 5	$\chi^2(5)$ = 6.2379 [0.2838]
	P_{MW}	$\alpha_{5i}=0$, for i=1, …, 5	$\chi^2(5)$ =8.9639 [0.1105]
	P_{NE}	$\alpha_{6i}=0$, for i=1, …, 5	$\chi^2(5)$ =13.569 [0.0186]*
	P_{RM}	$\alpha_{7i}=0$, for i=1, …, 5	$\chi^2(5)$ = 32.671 [0.0000]**
	P_{WC}	$\alpha_{8i}=0$, for i=1, …, 5	$\chi^2(5)$ = 19.753 [0.0014]**
(LE)| r=4	P_{CA}	$\beta_{j1}=0$, for j=1, …, 5	$\chi^2(5)$ = 21.249 [0.0007]**
	P_{EC}	$\beta_{j2}=0$, for j=1, …, 5	$\chi^2(5)$ = 12.304 [0.0308]*
	P_{GC}	$\beta_{j3}=0$, for j=1, …, 5	$\chi^2(5)$ = 17.971 [0.0030]**
	P_{LA}	$\beta_{j4}=0$, for j=1, …, 5	$\chi^2(5)$ = 10.782 [0.0559]
	P_{MW}	$\beta_{j5}=0$, for j=1, …, 5	$\chi^2(5)$ = 26.335 [0.0001]**
	P_{NE}	$\beta_{j6}=0$, for j=1, …, 5	$\chi^2(5)$ = 1.0869 [0.9553]
	P_{RM}	$\beta_{j7}=0$, for j=1, …, 5	$\chi^2(5)$ = 40.178 [0.0000]**
	P_{WC}	$\beta_{j8}=0$, for j=1, …, 5	$\chi^2(5)$ = 29.493 [0.0000]**
Normalization (N) + (WE) P_{GC} | r=5		$\beta_{ii}=1$, for i=1, …, 5 $\beta_{ij}=0$, for $\begin{cases} i = 1, \dots, 5 \\ j = 1, \dots, 5 \\ i \neq j \end{cases}$ $\alpha_{3i}=0$, for i=1, …, 5	$\chi^2(5)$ = 5.1254 [0.4008]
SE = (LE) + (WE)| r=5	P_{CA}	$\alpha_{1i}=0$, for i=1, …, 5 $\beta_{j1}=0$, for j=1, …, 5	$\chi^2(10)$ = 35.633 [0.0001]**
	P_{EC}	$\alpha_{2i}=0$, for i=1, …, 5 $\beta_{j2}=0$, for j=1, …, 5	$\chi^2(10)$ = 20.717 [0.0232]*
	P_{GC}	$\alpha_{3i}=0$, for i=1, …, 5 $\beta_{j3}=0$, for j=1, …, 5	$\chi^2(10)$ =22.520 [0.0127]*
	P_{LA}	$\alpha_{4i}=0$, for i=1, …, 5 $\beta_{j4}=0$, for j=1, …, 5	$\chi^2(10)$ =30.611 [0.0063] **
	P_{MW}	$\alpha_{5i}=0$, for i=1, …, 5 $\beta_{j5}=0$, for j=1, …, 5	$\chi^2(10)$ =32.287 [0.0004]**
	P_{NE}	$\alpha_{6i}=0$, for i=1, …, 5 $\beta_{j6}=0$, for j=1, …, 5	$\chi^2(10)$ =19.658 [0.0327]*
	P_{RM}	$\alpha_{7i}=0$, for i=1, …, 5 $\beta_{j7}=0$, for j=1, …, 5	$\chi^2(10)$ =49.721 [0.0000]**
	P_{WC}	$\alpha_{8i}=0$, for i=1, …, 5 $\beta_{j8}=0$, for j=1, …, 5	$\chi^2(10)$ = 46.086 [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.