Rising Household Diesel Consumption in the United States: a Cause for Concern? Evidence on Asymmetric Pricing.

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Abstract

Papers in the literature have thus far overlooked the projected increase in U.S. diesel car share when looking at asymmetries in petroleum pricing. This paper addresses this issue by comparing retail gasoline and diesel prices in order to see whether they rise faster than they fall given the price of their upstream input, crude oil. This phenomenon has been termed in the literature as "Rockets and Feathers." We apply the threshold vector error correction model (TVECM) of Hansen and Seo (2002) which has not yet been applied in the literature. We account for the 2008 structural break to crude oil and petroleum prices by splitting the sample using evidence from the recent structural break unit root test of Kim and Perron (2009). Both markets seem to price symmetrically before the 2008 break, but we find evidence of asymmetric pricing after 2008 in diesel prices, and not in gasoline prices. Given that the diesel market is small relative to the gasoline market and therefore more open to price exploitation, the ongoing cost increases associated with the policy of switching to Ultra Low Sulphur diesel (ULSD) from 2006 to 2010 could be at the heart of this asymmetry. With this in mind, the U.S. Federal Trade Commission should monitor diesel prices as the market share grows, in order to ensure that consumers are not adversely affected.

Keywords: Gasoline Pricing, Diesel Pricing, Asymmetric Adjustment, Threshold Cointegration, Cointegration, Structural Breaks.

JEL: C12, C22, Q40

1. Introduction and Literature Review

The market for petroleum in the United States has been hugely dominated by demand for ownership of gasoline powered cars. As of 2010 the share of diesel cars in new car sales was a mere 2.68%[1], far lower than in Europe where diesel cars take up over half of the market. As such, the literature on price adjustment in the U.S. petroleum markets has focussed almost exclusively on the retail price of gasoline. However, market experts expect that the share of diesel powered cars will treble[2] by the year 2015, a fact which has mostly been ignored by the literature to date and could have considerable welfare impacts if left without attention. This paper intends to address this issue, comparing the pricing of gasoline to that of diesel relative to the upstream crude oil price in the U.S.

The problem of asymmetric pricing has become of great interest to researchers due to the rising dependence of consumers on automobile transport, and therefore the welfare losses associated with price distortions and asymmetries in these markets. One of the earliest influential papers in this field was that of Bacon (1991) who followed a claim in the United Kingdom by the Monopolies and Mergers Competition that "the speed of adjustment of U.K. retail gasoline prices to cost changes is more rapid when costs rise than when they fall." In this context, asymmetric price adjustment has subsequently been termed the "Rockets and Feathers" problem.

This problem has indeed received some attention from the consumer protection agency in the U.S., the Federal Trade Commission (FTC), in the context of retail gasoline prices. This research is fairly mixed in its conclusions. For instance, Bachmeier and Griffin's (2003) work which started as an FTC paper, uses an asymmetric error correction model and finds no evidence for asymmetries in gasoline pricing implying a "very efficient market with few rigidities." On the other hand, Chesnes' (2010) FTC Working Paper finds, contrary to Bachmeier and Griffin (2003), that there *is* evidence for asymmetric pass-through. It is notable that even in the FTC, very little attention seems to have been paid to diesel price asymmetries.

Elsewhere, research into asymmetries in the North American retail gasoline market has also produced fairly mixed conclusions. For example, Godby et al (2000) used the bootstrap Threshold Autoregressive (TAR) procedure of Hansen (1996) and cannot find any significant impact of asymmetry in the Canadian gasoline market. More recently, Douglas (2010) used a univariate TAR model by Tsay (1989) to test the asymmetry hypothesis using the same dataset as the one used in this paper. He similarly finds that prices do not display as much asymmetry as believed and therefore that there is "minimal consumer welfare loss from asymmetry."

On the other hand, in an important work, Borenstein et al (1997) "test and confirm that retail gasoline prices respond more quickly to increases than to decreases in crude oil prices" when they look in depth at the production chain of gasoline in the United States. Chen at al (2005) use the threshold cointegration method of Enders and Siklos (2001), again using the same dataset as this paper. They find strong evidence of asymmetric adjustment to upstream crude oil prices, even when considering a variety of spot and future crude oil prices.

Whilst all of these papers address the problem of gasoline pricing, hardly any papers consider the case of diesel prices in the U.S. The only exception found is that of Johnson (2002) who looks at the gasoline and diesel markets in 15 major U.S. cities[3]. Using the Engle-Granger approach to

[1] Source: U.S. Department of Energy. See Section 2 and Footnote 4.
[2] Source: *Schmidt's Diesel Car Prospects to 2015.* See Section 2 and Footnote 5.
[3] I thank an anonymous referee for bringing this to my attention.

cointegration he confirms that pricing asymmetries exist in some of the cities for both gasoline and diesel by adding a dummy variable interaction into the error correction model. There has been more interest in diesel pricing in countries where diesel has a larger share of the car market. For example Liu et al (2010) analyse the comparison of gasoline and diesel prices in New Zealand using an asymmetric Mean Adjusted Lag (MAL) approach. They find that "oil companies adjust diesel prices upwards faster than they adjust them downwards" though this is not the case with gasoline.

This paper firstly analyses the developments in the retail gasoline and diesel markets in looking for reasons for asymmetric price adjustment. We also analyse the problem of structural breaks using a new test by Kim and Perron (2009) which endogenously determines the location of the breakpoint, leading to a scientific motivation for splitting the sample in this analysis. Then to test for asymmetry, we use the test of Hansen and Seo (2002), which has not been applied in the gasoline pricing literature, especially not to the U.S. This test is a multivariate threshold cointegration method estimated by maximum likelihood, unlike the less efficient two-step methods such as that of Enders and Siklos (2001) mentioned above. It should therefore provide a more robust insight into the true error correction behaviour of these series.

2. U.S. Petroleum Market – Scope for Pricing Asymmetries?

As mentioned before, the scope of this paper comes from the fact that previous studies have generally ignored the analysis of retail diesel prices in the U.S. market. This most probably has to do with the small volume of diesel light vehicles in the U.S. meaning that businesses rather than individuals would be affected by asymmetric pricing in the diesel market. Figure 1 shows a graph of the diesel share of new car sales in the U.S. taken from the *U.S. Department of Energy – Energy Efficiency and Renewable Energy*[4]. From this graph we can see that before 2009, the highest diesel share was a mere 0.82% in 2006. This figure has even risen from 0.08% in 1990. However, it can also be seen that, in general, this proportion has been rising over time, and in 2009 the share grew to a non-negligible 2.94%, before falling back to 2.68% in 2010. Even with this small rise, the omission of the analysis of retail diesel prices from previous papers is hardly surprising as this share is still small compared to the dominant sales of diesel cars in Europe which, according to *Schmidt's Diesel Car Prospects to 2015*[5] peaked at 53.3% in 2007.

[Figure 1 here]

However, the case for looking at retail diesel prices comes from the fact that industry experts seem to concur that the diesel share is going to grow in the next few years. The *Wall Street Journal* cites that "Frost & Sullivan predicts diesel unit sales will triple by 2015 from 2009 levels." [6] If accurate, this would bring the share of diesel cars to above 8% by 2015. Given this rising trend of ownership of diesel cars, it seems obvious that the analysis of price asymmetries in retail petroleum markets should also extend to diesel as a significant portion of U.S. citizens will be affected if market distortions were to be detected.

Turning to the petroleum market itself I will briefly analyse some reasons why market-specific factors may lead to symmetric or asymmetric pricing. Several authors justify asymmetric pricing through

[4] Available at URL: http://cta.ornl.gov/data/tedb30/Edition30_Chapter04.pdf [Accessed 20/06/2012.]
[5] Available at URL: http://www.eagleaid.com/dieselcarprospectsstudy.htm [Accessed 25/03/2011.]
[6] Available at URL: http://online.wsj.com/article/SB10001424052748703632304575451720343127244.html [Accessed 26/04/2011.]

market power arguments. Presence of oligopoly leading to asymmetric pricing has been used and empirically tested in papers such as Borenstein et al (1997). It may be reasonable to suggest that a low concentration of diesel cars means a shortage of retail diesel outlets, which are able to exploit monopoly power and price diesel asymmetrically. However, it is also key to note that diesel is still used fairly widely for fuelling heavy vehicles in the U.S.[7] While the *U.S. Energy Information Administration (EIA)* does not have data sources for the proportion of gas stations selling diesel, anecdotal evidence suggests that the figure is less than half. Given the nature of the diesel market, diesel pumps tend only to be found on interstate highways serving the trucking industry. This means there is more scope for localised monopoly powers than in the gasoline market, as gas stations in towns and cities off the interstate highways may struggle to provide diesel fuel profitably. This may allow diesel to be priced asymmetrically. For factual evidence, Figure 2 shows the share of motor gasoline to Ultra Low Sulphur diesel (ULSD) since 2007, using EIA data of U.S. prime supplier sales volume.[8] This confirms that the share of diesel sales is significantly lower than motor gasoline, only 27.6% in 2011, which again lends power to the argument that low diesel demand may force smaller sellers not to provide diesel and means there may be some enhanced reasons for suspecting asymmetric pricing than in the gasoline case.

[Figure 2 here]

On the other hand, another way people have tried to justify asymmetric pricing, such as the argument given by Douglas (2010) and Johnson (2002) is the search cost argument; that consumers tend to search for prices less intensely when prices decrease than when they increase. This means that suppliers can 'get away with' not posting lower prices when costs are falling, thus leading to the Rockets and Feathers phenomenon. In the context of the U.S. diesel market, since the vast majority of demand comes from truckers, as explained by Johnson (2002), who tend to buy in large quantities and therefore search more, we may expect that diesel prices are in fact *less* likely to be asymmetric given the underlying crude oil price. Which of these two opposite factors prevail is a matter for the empirical analysis to be conducted later.

One last consideration is that policy effects may have a role to play in the efficiency of price determination. After a proposal by the U.S. Environmental Protection Agency (EPA), since 2006 Ultra Low Sulphur Diesel (ULSD) fuel has been phased in as a requirement for highway vehicles. This is meant to reduce diesel engine sulphur emissions by up to 90%. As of late 2010, all highway vehicles had switched from the higher sulphur content diesel. As mentioned in an EIA report, some "diesel price premium can be attributed to costs associated with the transition to ULSD for highway freight fuel."[9] It may be the case that diesel pricing became more asymmetric after 2006 due to increases in the refinery costs to production, along with the market power argument of earlier, meaning increases in crude oil prices are passed on more quickly than decreases. This implies that there may be a structural break in the error correction response to the cointegrating vector some time after 2006, which must not be confused with structural breaks to the actual variables in the model, which we consider later. The next section discusses the data sources and looks as the stationarity properties of the data before moving onto an analysis of structural breaks using the new method proposed by Kim and Perron (2009).

[7] I thank an anonymous referee for raising this.
[8] Available at URL: http://www.eia.gov/dnav/pet/pet_cons_prim_dcu_nus_m.htm [Accessed: 21/06/2012.]
[9] Available at URL: http://www.eia.gov/oiaf/servicerpt/lightduty/chapter4.html [Accessed 19/06/2012.]

3. Data

3.1 Data Sources and Trends

This analysis follows the literature in looking for a simple bivariate long-run cointegrating relationship between retail petroleum prices and the upstream crude oil price similar to that of Chen et al (2005) amongst others. These can be described as:

$$p_t^G = \beta_0^G + \beta_1^G p_t^{CO} + \epsilon_t^G \tag{1}$$

$$p_t^D = \beta_0^D + \beta_1^D p_t^{CO} + \epsilon_t^D \tag{2}$$

for retail gasoline and diesel prices respectively. p_t^{CO}, p_t^D and p_t^G denote spot world crude oil price, and retail diesel and gasoline prices. We follow papers such as Borenstein et al (1997) in using the West Texas Intermediate (WTI) spot market price[10] as the proxy for world price as it is "the benchmark crude oil watched most closely in the U.S."[11] This data is at the weekly frequency taken from the EIA[12]. Retail gasoline and diesel prices are also taken from the EIA.[13] The retail prices are collected every Monday as the average price taken from around 900 gasoline outlets and 350 diesel outlets and are exactly the pump price paid by the consumer, which makes them appropriate for this analysis. The advantage of using this weekly dataset is that the data span combined with the data for crude oil is from January 1997 to June 2012, yielding a rich total of 806 observations. The same dataset, as mentioned above is also used by, amongst others, Chen at al (2005) and Douglas (2010).

Time series plots of the data are shown in Figure 3. This plot shows that while the gasoline and diesel prices clearly move in a similar fashion, there are also evident differences in the series. Most notably, even though the gasoline has in general been above the diesel price since 1997, there have also been significant spells where the diesel price has exceeded the gasoline price. It is these differences which make the two cases interesting to distinguish between in the analysis which follows.

[Figure 3 here]

Perhaps the most striking aspect of all three time series is the spike and crash in prices in 2008, caused by a number of factors including the bankruptcy of Lehmann Brothers in September of that year. With this in mind it is necessary to account for this break as it is likely that this will bias the results performed on the full sample. Rather than choosing an arbitrary location at which to split the sample, we will check for the structural break using a new unit root test proposed by Kim and Perron (2009) which allows the position of the break to be endogenously determined.

3.2 Unit Root Testing

In order to check the order of integration of the variables over the whole sample period we use the classic Augmented Dickey-Fuller test (1979) with the GLS-detrended extension of Elliot, Rothenberg

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[10] Though the Brent (Europe) crude oil price will be considered briefly later.

[11] An original version of this paper also used a proxy of the world crude oil price reported by the EIA, but which was discontinued in November 2011, so is not used here.

[12] Available at URL: http://www.eia.gov/dnav/pet/pet_pri_spt_s1_d.htm [Accessed 17/06/2012.]

[13] Available at URL: http://www.eia.gov/dnav/pet/pet_pri_gnd_dcus_nus_w.htm [Accessed 17/06/2012.]

and Stock (1996) and the test of Kwiatkowski, Phillips, Schmidt and Shin (1992) which reverses the test procedure to have stationarity under the null[14].

[Table 1 here]

The results in Table 1 show strong evidence that the WTI crude oil price and the retail diesel prices are nonstationary I(1) variables. The tests find mixed evidence of nonstationarity in the gasoline price with the ADF-type tests and the KPSS test finding opposite outcomes. However, given that all unit root tests resoundingly confirm that the data in differences are $I(0)$, we can conclude that there are no $I(2)$ processes in the data which is ideal for using a Johansen-type test for cointegration.

3.3 Structural Break Unit Root Testing

To formally test for the presence of the 2008 crash we observe in the variables, we use the new structural break unit root technique of Kim and Perron (2009). This procedure has several advantages over previous structural break unit root tests. The test allows for one structural break at a date which is determined by the model. Unlike tests such as that of Zivot and Andrews (1992), this test allows the break to occur under *both* the null and alternative hypotheses, as originally suggested by Perron (1989). This test is connected to that of Perron (1989) as they devise a procedure whereby the unknown breakpoint is first determined, and then added into the model as if it were known, meaning that the asymptotic distribution is the same as the original test of Perron (1989). To my knowledge this test has also not before been applied to this data.

I use the additive outlier (AO) model A3 of Kim and Perron (2009) as this allows for a break in both the constant and the trend, whereas models A1 and A2 allow for a break only in the constant *or* trend. Formally, for a given price series p_t, the model for the stationary data generating process (DGP) under the alternative is assumed to be:

$$p_t = \mu + \beta t + \mu_b DU_t + \beta_b B_t + u_t \tag{3}$$

Where $DU_t = B_t = 0$ if $T \leq T_1$ and $DU_t = 1, B_t = t - T_1$ for $t > T_1$. That is, both the mean and trend are allowed to break at the same location T_1. The error process is assumed to be the stationary ARMA process $A(L)\,u_t = B(L)\,\epsilon_t$ with $\epsilon_t \sim i.i.d.(0, \sigma_\epsilon^2)$. We do not consider the innovational outlier (IO) models also proposed in Kim and Perron (2009) as this makes the structural break evolve slowly over time according to the lag operator function $A^{-1}(L)\,B(L)$; something we are not interested in here. The test allows us to estimate the breakpoint fraction $\lambda \in (0,1)$ so that $T_1 = \lfloor \lambda T \rfloor$.[15] The unit root regression for model A3 is:

$$p_t = \alpha p_{t-1} + \sum_{j=0}^{k} w_j D(T_1)_{t-j} + \sum_{j=1}^{k} d_j \Delta p_{t-j} + u_t \tag{4}$$

Where p_t denotes the price series, detrended according to the deterministic components in the DGP displayed in Equation 3. k is chosen according to the Schwartz-Bayesian Information Criterion. $D(T_1)_t = 1$ if $t = T_1 + 1$ and 0 otherwise. Therefore Equation 4 incorporates a dummy at each point from k lags up until $T_1 + 1$. The test statistic is then the standard t-ratio on α, denoted t_α.

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[14] A constant and a trend were included for the test in levels and only a constant in the test in differences.
[15] The operator $\lfloor \cdot \rfloor$ denotes the integer part of the argument.

[Table 2 here]

The results of Model A3 are shown in Table 2. Using the critical values in Table IV.B of Perron (1989), we can see that again crude oil and diesel prices are seen to be I(1) and there is some evidence at 5% that gasoline prices may be I(0), though this does not seem natural. Most importantly, however, we see that the estimated breakpoints are incredibly similar for all three series and all of they occur seven weeks of one another between September and November 2008, during the crash. This clearly reflects economic and financial situations rather than policy changes in the petroleum markets, though the timing of the break is only two years after the implementation of the ULSD requirements.

Since large structural breaks tend to bias the estimates in cointegrating relations, not only of the error correction parameters but also of the short-run effects, it will be useful to split the sample at the location of the structural break. It is crucial to note that this reason for splitting the sample is distinct from the analysis of breaks to policy discussed in the previous section. However by splitting the sample in this way it will be possible to control for both, since the ULSD policy came into effect from 2006 to 2010, it will be possible to pick up the ongoing effects of this policy in the second subsample. While an ideal situation would be to split the sample in 2006 for the diesel model, the results of this regression would be biased by the rapid shifts in mean caused by the 2008 structural change, and it would be impossible to tell if changes in the results were driven by policy, or by statistical reasons regarding the structural break. However we will caveat the results with this notion.

3.4 Preliminary Linear Cointegration

Before proceeding to the TVECM analysis, we do a preliminary check that meaningful cointegrating relationships exist in the linear framework by looking at the trace test of Johansen's (1988) procedure. For the underlying cointegrating VECM, a constant is included in both the cointegrating vector and the short run equation since all of the variables seem to display some linear trend. Furthermore we include 3 lags of the differenced dependent variable[16] which corrects for serial correlation and allows for around a month's dynamic structure in the underlying VAR. The λ_{trace} statistics are reported in Table 3.

[Table 3 here]

The results of Table 3 indicate that the null hypothesis $r = 0$ is rejected in both cases and that $r \leq 1$ cannot be rejected, so we conclude that there exists a meaningful long-run cointegrating relationship between both gasoline and diesel prices with spot crude oil prices. With this in mind, in the next section we proceed to test for asymmetric cointegration using the method of Hansen and Seo (2002), which enables the analysis of asymmetric pricing in retail gasoline and diesel prices.

4. Threshold Cointegration

Threshold cointegration techniques have been introduced in order to relax the assumption of linear error correction to cointegrating vectors. The test of Enders and Siklos (2001), which is used by Chen et al (2005) in this literature, is a threshold autoregressive extension to the two-step univariate Engle-Granger (1987) cointegration approach. The test of Hansen and Seo (2002), on the other hand, is the threshold extension to the multivariate test of Johansen (1988). Using this model here can be considered a strong improvement over the likes of Chen et al (2005) as moving away from the two-step approach circumvents the problem of carrying over the errors from the first stage regression into

[16] Equivalent to 4 lags in the underlying VAR data generating process.

the second stage. This test has been applied elsewhere in the inflation literature by Esteve et al (2006). The linear cointegrating VECM (written with only one lag for simplicity) of Johansen (1988) for gasoline[17] can be written:

$$\begin{pmatrix} \Delta p_t^G \\ \Delta p_t^{CO} \end{pmatrix} = \begin{pmatrix} \mu_1 \\ \mu_2 \end{pmatrix} + \begin{pmatrix} \alpha_1 \\ \alpha_2 \end{pmatrix} w_{t-1}(\beta) + \Gamma \begin{pmatrix} \Delta p_{t-1}^G \\ \Delta p_{t-1}^{CO} \end{pmatrix} + u_t \tag{5}$$

Where $w_t(\beta) = p_t^G - \beta_1 p_t^{CO}$, with the coefficient on p_t^G normalised to 1. The threshold extension of Hansen and Seo (2002) on the other hand reformulates this model as:

$$\begin{pmatrix} \Delta p_t^G \\ \Delta p_t^{CO} \end{pmatrix} = \begin{cases} \begin{pmatrix} \mu_{11} \\ \mu_{12} \end{pmatrix} + \begin{pmatrix} \alpha_{11} \\ \alpha_{12} \end{pmatrix} w_{t-1}(\beta) + \Gamma_1 \begin{pmatrix} \Delta p_{t-1}^G \\ \Delta p_{t-1}^{CO} \end{pmatrix} + u_{1t} & if\ w_{t-1}(\beta) \leq \gamma \\ \begin{pmatrix} \mu_{21} \\ \mu_{22} \end{pmatrix} + \begin{pmatrix} \alpha_{21} \\ \alpha_{22} \end{pmatrix} w_{t-1}(\beta) + \Gamma_2 \begin{pmatrix} \Delta p_{t-1}^G \\ \Delta p_{t-1}^{CO} \end{pmatrix} + u_{2t} & if\ w_{t-1}(\beta) > \gamma \end{cases} \tag{6}$$

More generally for ease of exposition and to explain the following tests, Equation 6 can be written in compact form as in Section 2.2 of Hansen and Seo (2002) as:

$$\Delta x_t = \begin{cases} A_1' X_{t-1}(\beta) + u_t & if\ w_{t-1}(\beta) \leq \gamma \\ A_2' X_{t-1}(\beta) + u_t & if\ w_{t-1}(\beta) > \gamma \end{cases} \tag{7}$$

Where $x_t = [p_t^G, p_t^{CO}]'$ and $X_{t-1}(\beta) = [1, w_{t-1}, \Delta x_{t-1}, \ldots, \Delta x_{t-k}]'$, and the matrices A_1 and A_2 collect up the parameters μ, α and Γ from Equation 6.

In other words, there are two different matrices A_1' and A_2' which govern the dynamics of the system contingent on the value taken by the error correction term $w_t(\beta)$ relative to some threshold value γ. Note that in this specification with $w_t(\beta) = p_t^G - \beta_1 p_t^{CO}$, the condition $w_{t-1}(\beta) \leq \gamma$ reads $p_t^G \leq \gamma + \beta_1 p_t^{CO}$. So as an *a priori* prediction, for the hypothesis that petroleum prices increase faster than they decrease, we expect that the error-correction coefficient on p_t^G (and p_t^D in the diesel model) to be more negative in matrix A_1' than in A_2', that is $\alpha_{11} > \alpha_{21}$. The error correction term on p_t^G and p_t^D should not be positive or zero as these variables' coefficients are normalised to 1 in the cointegrating vector, so a non-negative error correction term would imply an explosive time path not consistent with equilibrium error correction.

To estimate the VECM based on the parameters β and γ, Hansen and Seo (2002) suggest the maximum likelihood estimation of this model using a grid search procedure in (β, γ) space since "conventional hill-climbing algorithms are not suitable" for maximising the non-smooth likelihood function they derive. In order to get sensible search regions for these they suggest using a region for these parameters based on the corresponding estimates from the linear model *a la* Johansen (1988) which can be proved to be consistent. As in Hansen and Seo (2002) we use a 300 by 300 grid to search for the estimators which maximise the log-likelihood.

The most important contribution of Hansen and Seo (2002) is that they provide a way to test the hypothesis:

[17] Clearly the equations for diesel will look exactly the same, but with the superscript 'G' replaced with 'D'.

$H_0: A_1 = A_2$ (Linear Cointegration)

$H_A: A_1 \neq A_2$ (Two-regime Threshold Cointegration)

In other words, since the set of linear VECMs is nested within the set of two-regime TVECMs, we can test H_0 directly. Note that this is *not* a test for symmetric vs. asymmetric pricing, as A_1 and A_2 include all of the coefficients including lagged dependent variables. Being able to test for the presence of a threshold, rather than simply assuming there is one, is a notable advance over the techniques used in papers such as Chen et al (2005). Hansen and Seo (2002) propose a heteroskedasticity-robust test for the presence of a threshold, and has the form of an LM test. They denote this statistic:

$$\mathrm{SupLM} = \sup_{\gamma_L \leq \gamma \leq \gamma_U} LM\ \beta, \gamma \tag{8}$$

where $[\gamma_L, \gamma_U]$ is the search region for the threshold γ and is set by using some π_0 percentile of the error correction term at both ends, where π_0 is the trimming parameter. We set the trimming parameter at 0.1 to ensure a non-trivial amount of observations in each regime. Hansen and Seo (2002) also provide a method to calculate the critical values and associated p-values to this SupLM statistic using fixed regressor and residual bootstrap methods.

If H_0 is rejected and we conclude that the TVECM model is the most appropriate, we can then test for asymmetric pricing. For this we use the Wald statistic for the equality of the error correction coefficient across the two regimes; the test that, for a given model the vector $[\alpha_{11}, \alpha_{12}]' = [\alpha_{21}, \alpha_{22}]'$. We also present the Wald test for the equality of the coefficients of the lagged dependent variables; the test that the matrix $\Gamma_1 = \Gamma_2$. These estimates are presented in the next section separately for gasoline and diesel and for the split sample sizes discussed in Section 3.3.

5. Results

The results of the test statistics described above are reported here in Tables 4(a) and 4(b). These results show that, when looking over the full sample period, the value of SupLM is not statistically different from zero at the 5% level for either the case of gasoline or diesel. On face value this seems to give evidence not to reject the null, and favour the linear cointegrating model, implying symmetric pricing. However, as mentioned before, having found statistical evidence for the structural break of 2008, subsample regressions were run to split the sample around the large boom and subsequent crash in crude oil prices. On performing these tests on the split sample, a different story emerges.

[Tables 4(a) and 4(b) here]

For the pre-break period, we can see from the middle column of Tables 4(a) and 4(b) that the SupLM test statistics are not statistically different from zero in either the gasoline or diesel case. This means that we can reject the presence of a threshold. Given that there are no significant structural breaks in this period it seems plausible to conclude in favour of H_0 in both cases and conclude symmetric pricing as dictated by a linear VECM model.

However for the post-break period we see the reverse conclusion that the SupLM statistics for both gasoline and diesel become statistically significant from zero at the 5% level or lower. This means that we reject the null hypothesis of linear cointegration in favour of the two regime threshold alternative. Given that the TVECM model is the most appropriate in both cases, it is now possible to look at evidence from the Wald tests. For the gasoline model, we see that the hypothesis of equality of the error correction coefficients cannot be rejected, so we would conclude that, like in the pre-break period, we have symmetric pricing post-break. The significance of SupLM in this case comes from inequality of the dynamic lagged dependent variables, indicating some nonlinear short-run behaviour in gasoline prices, which could be expected in times of financial crisis and fluctuation in economics aggregates. We therefore conclude similarly to Bachmeier and Griffin's (2003) work on gasoline pricing who state that this is a "very efficient market with few rigidities."

Turning to the diesel results however, we see that unlike gasoline we *do* reject the hypothesis of equality of the error correction coefficients across regimes for the post-break period. Therefore with this evidence we may conclude that asymmetric pricing is present in the retail diesel market. This is particularly interesting as the result changes when compared to the period before 2008. As mentioned before, it is impossible to allow simultaneously for the large structural break occurring in the data for the 2008 rise and crash of oil prices, and allow for the policy break to the error correction mechanism whilst still remaining within the threshold cointegration framework. Nevertheless, with the caveat that we would have preferred to have split the diesel sample in 2006, this analysis may lend some evidence that ongoing refining cost increases as a result of switching to Ultra Low Sulphur diesel from 2006 to 2010 were passed onto consumers in an asymmetric fashion by sellers exploiting some degree of market power. Johnson's (2002) study only looks at diesel prices over the period from 1996 to 1998 so is not really comparable with our study, but he finds more evidence for symmetry out of the 15 cities analysed, which is similar to our findings over that part of the sample.

For completeness, the estimation results of the TVECM models for both the gasoline and diesel systems in the post-break period are provided in the Tables 5(a) and 5(b). These are the only two models in which the null of linear cointegration is rejected. The most important part of these results is that they show the statistical significance of the individual error correction parameters, whereas the results in Tables 4(a) and 4(b) tested only joint hypotheses. Firstly we note that in both models Regime 2, where prices are above equilibrium, is the "extreme" regime (the term coined by Hansen and Seo, 2002) with only 12.71% of the observations in the diesel case and 20.65% in the gasoline case. In the "typical" regime diesel has the correctly signed error correction term ensuring the model reaches equilibrium. The model seems to show instability in Regime 2 as neither error correction term is significant, hence the conclusion overall that error is corrected asymmetrically. The gasoline model is in some sense 'better behaved' as the gasoline price carries a negative error correction parameter in both regimes, and the overall conclusion of symmetry. Note that in the "typical" regimes in both models, error correction comes only in the form of changes to the gasoline or diesel price, not the crude oil price, as we would expect *a priori*.

[Tables 5(a) and 5(b) here]

In addition to the analysis of structural breaks, some other robustness measures were considered. Since the EIA also publishes the Brent crude oil price in its section on world spot crude oil prices it was seen whether using this measure would have an impact on the results. This does not seem to be considered in other papers such as Borenstein et al (1997) primarily because in the U.S. WTI crude oil is still seen as the benchmark, whereas Brent crude oil is mainly seen as the benchmark for petroleum pricing in Europe. Furthermore it must be noted that *a priori* we would expect the results for diesel to change by more than the results for gasoline. After the Ultra Low Sulphur requirements for diesel from 2006 to 2010, production has demanded sweeter blends of crude oil. As WTI is indeed a sweeter type when compared to Brent crude, we may expect the WTI price to have a more relevant impact in the diesel model. Finally is should be noted that the Brent crude oil price has followed a very much divergent path to other world crude oil prices since 2010 for a variety of factors, most importantly due to North Sea oil supply contractions, which is independent of petroleum pricing in the U.S.

Nevertheless, the results for running the Hansen and Seo (2002) model for this type of crude oil price are displayed in the Appendix. The results are similar in that the SupLM test gives no evidence for a threshold in the pre-break period, but in the post-break period we do see a threshold present. In turning to the Wald statistics there seems to be no evidence for asymmetric pricing in diesel anymore. However on closer inspection we also see a dramatic difference in the Wald test on the dynamic coefficients, Γ, which become more statistically significant in both models. Therefore the greatest difference in using the Brent price relative to the WTI crude oil price comes from changes in the short run dynamics. This shows that the results of running this test can be sensitive to the underlying crude oil price used, though for testing the hypothesis of asymmetry we feel that the WTI results are the most reliable as the Brent price has diverged in ways which are in some sense independent of the US petroleum market. Perhaps an interesting area of future research would be to look at this kind of asymmetric pricing in European petroleum markets, where the Brent crude oil price is a more relevant measure[18].

6. Conclusion

This paper has added two elements to the literature on asymmetric pricing of petroleum in the U.S. Firstly, to analyse the question of pricing asymmetries relative to the price of crude oil, we used the multivariate threshold cointegration test of Hansen and Seo (2002), which we believe has not yet been used in the literature in any petroleum market analysis. Secondly this paper paid attention to the price of diesel which has been virtually ignored in the U.S. literature to date. Given that the market share of diesel cars is very low, but on the rise, findings of asymmetric pricing could have a significant impact on consumers in years to come.

In performing the tests, it was necessary to split the sample as the large structural break in the variables in 2008 due to the crude oil price spike and crash was likely to have biased results which used the full sample period. The recent structural break unit root test of Kim and Perron (2009) was used, which lets the position of the break be determined endogenously. This breakpoint was used to split the sample. Using Hansen and Seo's (2002) SupLM test for the presence of a threshold, along with associated fixed regressor and residual bootstrap p-values, we find evidence that pricing was symmetric for both gasoline and diesel in the calmer period before the 2008 crude oil price movements. However when turning to the post-break results, there is significant evidence for a threshold in both models, but in the case of gasoline there is no evidence of asymmetric pricing, only asymmetries in short-run dynamic behaviour.

[18] Note: As an additional robustness measure, for which we thank an anonymous referee's suggestion, we also checked for seasonal patterns in the data. Since demand for gasoline is stronger in the summer and demand for diesel is stronger in the winter for heating fuel, we may expect seasonality to play a role. However such seasonal patterns do not appear to be particularly strong in the price data we use. Furthermore as we use weekly data for which it makes little sense to apply seasonal adjustments, we dropped down to the monthly frequency. Not only did this lose a lot of data, but upon applying the X12-ARIMA seasonal adjustment method, if anything the data became *more* 'spiky'. As such, regressions were run but not reported as they contain little useful information, though the results for the seasonally adjusted data were similar to that of the non-seasonally adjusted data.

In the diesel model, on the other hand, a Wald test for equality of the error correction coefficients rejects the null meaning there is evidence that price rises in crude oil are passed on more quickly than price drops. A possible explanation for this is that, in the diesel market which is much smaller than that of gasoline, oligopolistic sellers asymmetrically passed on the increased refinery costs following the policy change of switching to Ultra Low Sulphur diesel (ULSD) which occurred between 2006 and 2010. This evidence seems to disappear when switching from using the WTI crude oil price to the Brent price, but we feel that the WTI results are more relevant, especially to diesel which requires a sweeter crude oil like WTI in the production process due to this switch to low sulphur fuel.

If asymmetry in diesel prices was caused by the ULSD policy, it could be expected that this will fade over time and we will see evidence for symmetric pricing in the near future. Nevertheless, this result could be of great interest to the U.S. Federal Trade Commission, in a similar way that Bacon's original (1991) paper was driven by the U.K.'s Monopolies and Mergers Competition. If the market share of diesel cars in the U.S. is set to rise to over 8% by 2015, then the question of asymmetric pricing in the diesel market should be carefully monitored so that consumers are not harmed by the pricing decisions of petroleum companies. The evidence from this paper would suggest that this increasing proportion of diesel consumers are currently experiencing asymmetric pricing strategies and should be wary that in current market conditions, there is evidence suggesting that diesel prices tend to rise more quickly than they fall.

7. Appendix

Appendix(a): Hansen and Seo (2002) Threshold Cointegration Results for Gasoline and Brent Crude Oil

	Gasoline		
	Full Sample	Jan 1997-Oct 2008	Oct 2008-Jun 2012
T	806	618	188
SupLM	28.6472	23.5116	36.59374
Fixed Regressor CV	33.1445	33.0545	30.3454
p-value	0.2030	0.5446	0.0036
Bootstrap CV	32.9836	31.4614	26.4934
p-value	0.1396	0.3892	0.0014
Wald(Γ)	30.9436	18.1454	93.3530
p-value	0.0020	0.1113	0.0000
Wald(α)	5.4153	6.9276	2.2991
p-value	0.0667	0.0313	0.3168

Appendix 4(b): Hansen and Seo (2002) Threshold Cointegration Results for Diesel and Brent Crude Oil

	Diesel		
	Full Sample	Jan 1997-Nov 2008	Dec 2008-Nov 2012
T	806	621	185
SupLM	36.8744	29.1141	53.01912
Fixed Regressor	31.7819	30.5103	30.28831

CV			
p-value	0.0082	0.0836	0.0000
Bootstrap CV	32.2641	31.6833	27.8041
p-value	0.0100	0.1082	0.0000
Wald(Γ)	34.1253	44.6939	77.0585
p-value	0.0006	0.0000	0.0000
Wald(α)	5.5227	5.0725	1.1857
p-value	0.0632	0.0792	0.5527

Appendix (a) and (b): Results for the gasoline and diesel model of the Hansen and Seo (2002) SupLM statistic to test the null of linear cointegration, with associated Fixed Regressor and Residual Bootstrap 5% critical values. Wald(Γ) and Wald(α) denote the Wald statistics for testing the equality across regimes of the lagged dependent variable coefficients and the error correction coefficients respectively.

8. Acknowledgement

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Figure 1: Diesel Share of New Retail Car Sales in the United States

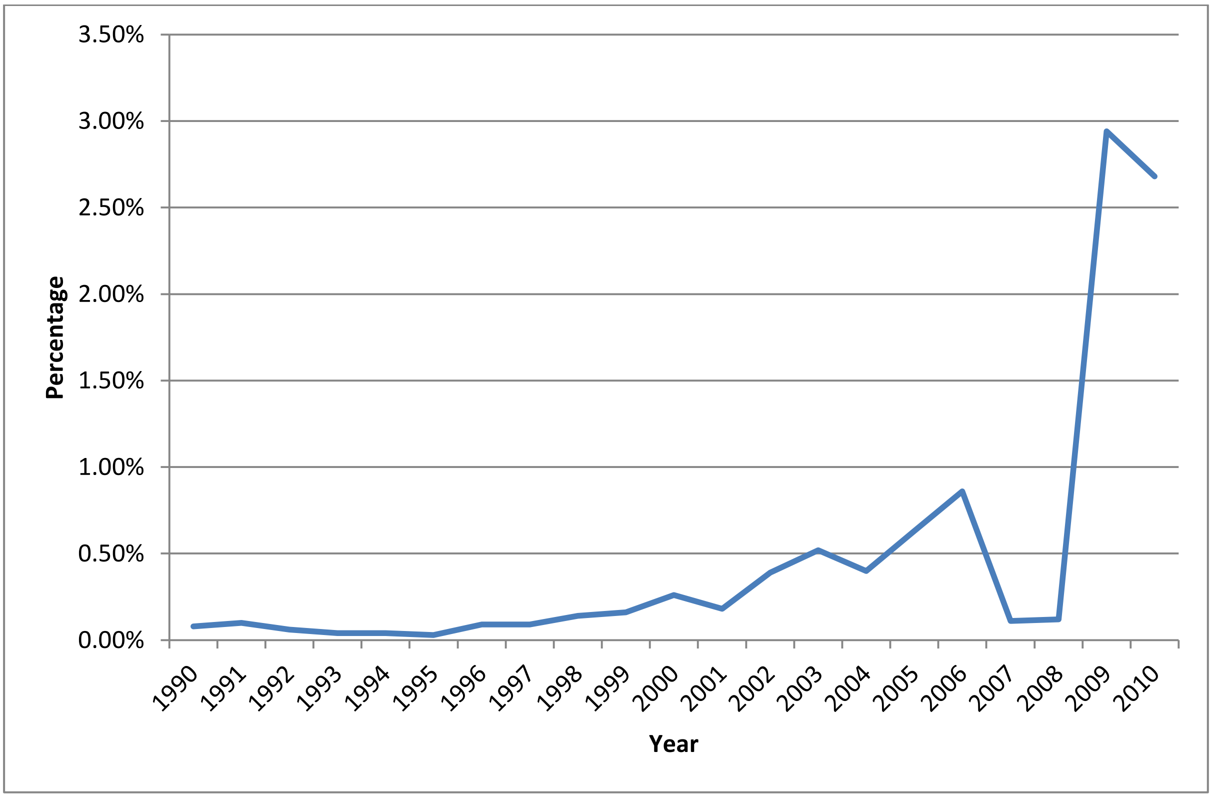

Figure 1: Time series showing the percentage share of new retail car sales going to diesel fuelled cars in the United States, from 1990-2010. Source: *U.S. Department of Energy – Energy Efficiency and Renewable Energy*.

Figure 2: Prime Supplier Sales of Motor Gasoline and ULSD from 2007-2012.

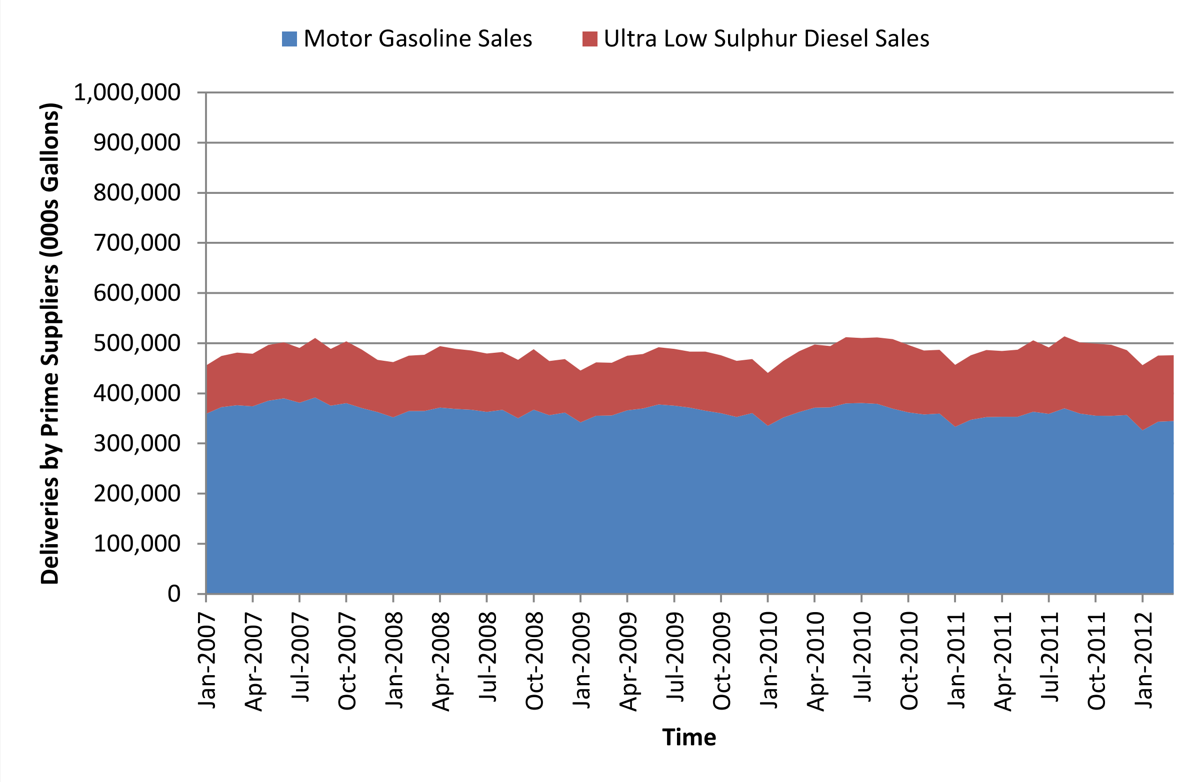

Figure 2: Prime Supplier Sales of Diesel and Gasoline. Source: U.S. Energy Information Administration.

Figure 3: Spot Crude Oil and Retail Gasoline and Diesel Prices in the United States

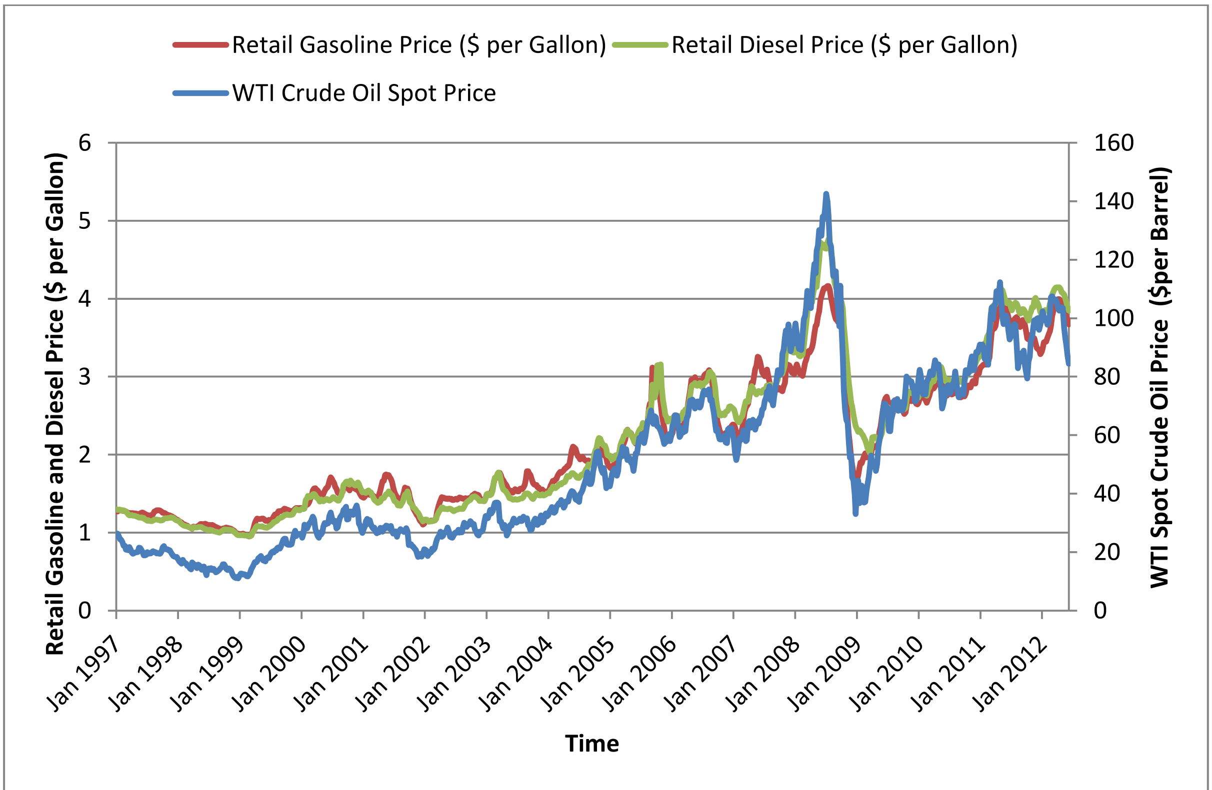

Figure 3: Time Series plots of spot WTI crude oil prices and retail gasoline and diesel prices. Source: *U.S. Energy Information Administration*.

Table 1: Unit Root Tests

Variable	ADF	Lag	ERS	Lag	KPSS	Lag
p_t^{CO}	-3.0874	1	-2.4284	1	0.1485 **	-
p_t^{D}	-3.3561 *	3	-2.6191 *	3	0.1810 **	-
p_t^{G}	-4.6780 ***	3	-3.9068 ***	3	0.1482 **	-
Δp_t^{CO}	-23.5157 ***	0	-22.9744 ***	0	0.0327	-
Δp_t^{D}	-10.6324 ***	2	-10.4685 ***	2	0.0299	-
Δp_t^{G}	-9.6956 ***	2	-9.5734 ***	2	0.0190	-

Table 1: Unit root tests on the levels and differences of the WTI spot crude oil price and the retail gasoline and diesel price. *, ** and *** denote rejection of the relevant null hypothesis at the 10%, 5% and 1% level respectively. Lag length, where relevant, was selected using the Schwartz-Bayesian Information Criterion.

Table 2: Structural Break Unit Root Test

Variable	λ	Break Date λT	t_α	k
p_t^{CO}	0.7655	26/09/2008	-3.3127	1
p_t^{D}	0.7705	21/11/2008	-3.2956	1
p_t^{G}	0.7667	31/10/2008	-4.3279 **	2

Table 2: Results of Model A3 of Kim and Perron (2009) on the levels of the WTI spot crude oil price and the retail gasoline and diesel price.

Table 3: Johansen's (1988) Trace Test

	Gasoline		Diesel	
H_0	λ_{trace}	C.V.	λ_{trace}	C.V.
$r = 0$	30.4717	15.49	35.3944	15.49
$r \leq 1$	2.3852	3.84	2.3767	.84

Table 2: Trace tests for number of cointegrating vectors in the gasoline and diesel systems. 5% Critical Values (C.V.) taken from Osterwald-Lenum (1992).

Table 4(a): Hansen and Seo (2002) Threshold Cointegration Results for Gasoline and WTI Crude Oil

	Gasoline		
	Full Sample	Jan 1997-Oct 2008	Oct 2008-Jun 2012
T	806	618	188
SupLM	31.7514	22.1964	29.71714
Fixed Regressor CV	31.9196	32.2956	30.2436
p-value	0.0532	0.6460	0.0600
Bootstrap CV	32.1867	30.8240	28.4802

p-value	0.0582	0.4262	0.0312
Wald(Γ)	32.5971	25.31246	63.5340
p-value	0.0011	0.0134	0.0000
Wald(α)	0.3167	9.8180	2.8886
p-value	0.8535	0.0074	0.2359

Table 4(b): Hansen and Seo (2002) Threshold Cointegration Results for Diesel and WTI Crude Oil

	Diesel		
	Full Sample	Jan 1997-Nov 2008	Dec 2008-Nov 2012
T	806	621	185
SupLM	26.3558	24.8406	32.3841
Fixed Regressor CV	31.5786	30.3143	30.42261
p-value	0.2376	0.2960	0.0228
Bootstrap CV	31.9665	31.5579	27.8208
p-value	0.2396	0.2870	0.0100
Wald(Γ)	66.6159	39.6384	138.9158
p-value	0.0000	0.0001	0.0000
Wald(α)	4.5334	0.8257	10.4729
p-value	0.1037	0.6618	0.0053

Tables 4(a) and 4(b): Results for the gasoline and diesel model of the Hansen and Seo (2002) SupLM statistic to test the null of linear cointegration, with associated Fixed Regressor and Residual Bootstrap 5% critical values. Wald(Γ) and Wald(α) denote the Wald statistics for testing the equality across regimes of the lagged dependent variable coefficients and the error correction coefficients respectively.

Table 5(a): TVECM Regression Results for the Diesel Post-Break Model

	Regime 1: (87.29% obs)					
	Dependent Variable: Δp_t^D			Dependent Variable: Δp_t^{CO}		
	Coefficient	S.E.	t-ratio	Coefficient	S.E.	t-ratio
$w_{t-1}\ \beta$	-0.0312	0.0134	-2.3341	0.0616	0.0476	1.2941
C	0.3576	0.1051	3.4025	-0.0218	0.4492	-0.0485
Δp_{t-1}^D	0.4286	0.0954	4.4937	-0.3624	0.3107	-1.1664
Δp_{t-2}^D	0.2691	0.0236	11.4246	0.2835	0.0780	3.6359
Δp_{t-3}^D	0.0576	0.0824	0.6984	0.6154	0.3019	2.0385
Δp_{t-1}^{CO}	-0.0074	0.0343	-0.2156	-0.0064	0.1357	-0.0475
Δp_{t-2}^{CO}	-0.1082	0.0492	-2.1996	-0.4006	0.2140	-1.8724
Δp_{t-3}^{CO}	-0.0071	0.0322	-0.2213	-0.1216	0.0967	-1.2584

	Regime 2: (12.71% obs)					
	Dependent Variable: Δp_t^D			Dependent Variable: Δp_t^{CO}		
	Coefficient	S.E.	t-ratio	Coefficient	S.E.	t-ratio
$w_{t-1}\ \beta$	0.0188	0.0307	0.6111	0.7557	0.2479	3.0483

C	-1.3007	0.7807	-1.6661	-19.5028	6.7920	-2.8714
Δp_{t-1}^{D}	0.3070	0.0876	3.5031	3.8885	0.8541	4.5531
Δp_{t-2}^{D}	0.0372	0.0231	1.6086	0.2726	0.1899	1.4352
Δp_{t-3}^{D}	0.0395	0.1416	0.2788	-1.6309	0.9900	-1.6474
Δp_{t-1}^{CO}	0.0751	0.0275	2.7302	-0.4222	0.1314	-3.2130
Δp_{t-2}^{CO}	-0.1422	0.1154	-1.2326	0.6171	0.6193	0.9964
Δp_{t-3}^{CO}	-0.0267	0.0206	-1.2943	-0.2129	0.2255	-0.9441

Table 5(b): TVECM Regression Results for the Gasoline Post-Break Model

	Regime 1: (79.35% obs)					
	Dependent Variable: Δp_t^{G}			Dependent Variable: Δp_t^{CO}		
	Coefficient	S.E.	t-ratio	Coefficient	S.E.	t-ratio
$w_{t-1}\ \beta$	-0.0304	0.0157	-1.9309	0.0318	0.0625	0.5095
C	0.4362	0.1348	3.2355	-0.4398	0.5383	-0.8170
Δp_{t-1}^{G}	0.4762	0.0698	6.8266	0.0642	0.2990	0.2146
Δp_{t-2}^{G}	0.2912	0.0255	11.4079	0.2628	0.0825	3.1852
Δp_{t-3}^{G}	0.2507	0.0809	3.0979	-0.0311	0.2655	-0.1171
Δp_{t-1}^{CO}	-0.0227	0.0331	-0.6853	-0.0978	0.1536	-0.6366
Δp_{t-2}^{CO}	-0.0710	0.0520	-1.3656	0.1599	0.1996	0.8011
Δp_{t-3}^{CO}	-0.1720	0.0306	-5.6255	0.1258	0.1183	1.0628

	Regime 2: (20.65% obs)					
	Dependent Variable: Δp_t^{G}			Dependent Variable: Δp_t^{CO}		
	Coefficient	S.E.	t-ratio	Coefficient	S.E.	t-ratio
$w_{t-1}\ \beta$	-0.0012	0.0342	-0.0345	-0.1648	0.1199	-1.3743
C	-0.2864	0.7636	-0.3751	4.1334	2.8038	1.4742
Δp_{t-1}^{G}	0.3643	0.1666	2.1860	0.9855	0.4306	2.2889
Δp_{t-2}^{G}	0.1252	0.0376	3.3335	0.0015	0.1459	0.0103
Δp_{t-3}^{G}	0.0911	0.1859	0.4899	0.0611	0.4156	0.1471
Δp_{t-1}^{CO}	0.0591	0.0545	1.0840	-0.1728	0.1357	-1.2727
Δp_{t-2}^{CO}	-0.3355	0.1357	-2.4722	-1.0109	0.4103	-2.4639
Δp_{t-3}^{CO}	0.0352	0.0660	0.5340	-0.3929	0.1746	-2.2506

Tables 5(a) and 5(b): Full regression results with standard errors and t-ratios from the post-break TVECM model of Hansen and Seo (2002) for diesel and gasoline.