Petrol and Crude Oil Prices: Asymmetric Price Transmission

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Abstract:

This paper examines the relationship between crude oil prices, the dollar-pound exchange rate and petrol prices in the UK over the period 1982-2001. Quantitative methods were used to examine the existence of the long-run equilibrium and test for the presence of asymmetric patterns in the short-run responses to upstream price changes. Also the degree of asymmetry in the adjustment towards long-run equilibrium was analysed. Results confirm that short-run response is greater for increases in upstream prices and that the long-run equilibrium is reached faster after increase in upstream prices. Thus the opinion held by drivers in the UK is confirmed. Detailed analysis confirmed close relationship between the asymmetry and the size and change in the market margin.

Introduction

The study of the asymmetric price transmission is quite matured, but every year brings new developments. Unfortunately, they are focused on the empirical part of the problem. Samuel Peltzman (2000:466) stated even that:

> 'Output prices tend to respond faster to input increases than to decreases (....) It is found as frequently in producer goods as in consumer goods market (....) (it) suggest *a gap in the essential part of economic theory'* (emphasis added)

This paper is not aimed at filling this gap but only tries to establish whether petrol prices really respond faster to crude oil price increases than to decreases. This requires focusing mainly on the empirical work. The hope of the author is that further theoretical developments will bring an answer to the asymmetry puzzle..

The model created in this paper is designed to look for asymmetric patterns in the process of reverting to the long run equilibrium after shocks to the upstream variables. It can be use to assess asymmetries in interest rates and prices. It can also be used to detect the impact of the transaction costs (consistent versions of TAR and M-TAR models) in many markets. Last but not least, it may be able to detect the impact of foreign exchange rate bands. Its multiple use is a by-product of its original features[1].

The composition is as follows. First part presents up-to-date developments in the theory behind modelling asymmetric price transmission. It also reports results of previous studies. This review is aimed only at familiarizing the reader with the state-of-the art not at providing the comprehensive study of the modeling techniques. This is the task of the second part in which, the empirical methods developed for assessing asymmetry in price transmissions are described.

Next part describes the results of the application of time series techniques aimed at characterising the nature of the long-run relationship between crude oil prices and retail prices of 4 Star petrol. The last sections deal directly with the issue of asymmetry.

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[1] In the 1980s models similar to TAR and M-TAR were used for example to investigate the number of lynxes in their sunspots in Canada (sic!).

1 Theoretical Background

1.1 Review of the Literature

Perhaps the first study to analyse the response of pump prices to crude oil cost changes was by Sumner (1990), Using Engle and Granger's two-step procedure he examined the relationship between net retail prices of 4 Star petrol, the producer price of crude oil acquired by refineries and labour costs (index of weekly earnings in manufacturing) over the sample February 1981 - December 1989. The tests for cointegration were performed and existence of long-run relationship was established. For testing for the asymmetry he used the Wolfram-type segmentation of first differences of explanatory variables. The evidence of faster response of retail prices to increases as opposite to decreases in prices of crude oil was found. To check whether there was any offset through the ECM to the asymmetrical effects of cost changes the error correction term (lagged residuals from the level regression) was divided into Wolfram's manner. Results were not significantly changed, which can be seen as evidence that over this period asymmetry was persistent.

The problem of a different response to price increases and decreases was also considered in Bacon's (1991) article. His work on the UK petrol market was focused on the transmission of spot crude oil prices (the ex-Rotterdam price was used as proxy for producer prices) to retail petrol prices. Biweekly data for the period 1982-1989 was used, and since the prices of crude oil were in U.S. dollars the impact of the exchange rate was also taken into consideration. Evidence on asymmetry in transmission was found. Author concluded that increases in the crude oil prices are fully transmitted within 8 weeks while approximately one extra week is necessary for full transmission of the decreases in crude oil prices.

Manning's (1991) work was inspired by the debate over the pricing policies used by major oil companies (so-called majors) for the retail price of petrol in the UK. He considered whether the retail prices of 4 Star petrol exhibit a long-run relationship with the level of Brent oil prices and excise duties using the techniques based on simple ECM. After ascertaining that the long-run relationship does exist[2] he considered whether the short-run adjustment of petrol prices to changes in the oil price is symmetrical or whether petrol retailers change petrol prices less rapidly and by lower amount in times of falling oil prices. Monthly information

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[2] One reservation applies here: the t-statistic on error-correction term (lagged residual from level equation) has t-statistic equal to 2.45 only, which is obviously lower than critical values tabulated e.g. in Banerjee et al. (1993)), no other test for cointegration was reported.

from January 1973 to December 1989 was used to show that asymmetric pricing reactions by retailing companies was indeed present in the short-run although the difference between adjustment to increasing, as opposed to falling, oil prices was 'barely evident four months after the initial price change' (Manning 1991:1535). He also concluded that approximately 7% of any deviation from the long run equilibrium was corrected or eliminated each month. The equation 3 on page 1539 with small and easy-to-correct mistake ($\Delta \ln p_{t-1}$ instead of $\ln p_{t-1}$) gives slightly greater speed of adjustment – 7.29% of disequilibrium is eliminated every month. His findings that disequilibria were persisted (only 60% was supposed to be eliminated after 12 months) were in contrast with other studies on the UK market.

Reilly and Witt (1998) analysed the UK market and revisited the evidence of Bacon (1991) and Manning (1991) with monthly data for 1982-1995 stressing the importance of dollar-pound exchange rate and potential asymmetries associated with it, in addition to those of crude oil prices. An unrestricted ECM, which allows for short-run asymmetries in adjustment to changes in the exchange rate and crude oil prices, was used and the evidence of asymmetries was found. Authors concluded that the retail prices increase by more when the prices of crude oil increase, the same with the unfavourable changes in the exchange rate

The only study known to the author, concerned with the issue of interest in more than one country was by Galeotti, Lanza and Manera (2002). They used monthly data (from 1985 to 2000) on prices at two stages of transmission (first refinery stage and second distribution stage) for five European countries (Germany, France, UK, Italy and Spain) to test for asymmetric transmission. They also allowed for asymmetries caused by the exchange rate.

The enriched, asymmetric ECM was used to distinguish between asymmetries that arise from short-run deviations in input prices and from the speed at which the petrol price reverts to its long-run level. Statistical inference was based on bootstrapped F-tests of asymmetries in order to overcome the low-power problem of conventional testing procedures. The results indicated widespread differences among European Countries in both adjustment speeds and short run responses to crude oil prices rises and falls.

Transmission from crude oil prices to retail petrol prices in the UK was found to be asymmetric. Detailed analysis of the second stage of transmission (from spot wholesale price to retail prices) indicated that also on this stage asymmetries in the short-run responses and adjustment speed were present

1.2 Theory behind Testing for Asymmetric Price Transmission

1.2.1 Asymmetric ECM

The first model aimed at analysing the issue of asymmetric price transmission was developed by Houck in 1977. Many authors have used approaches for analysing price transmission, which differ in some technical regards but nonetheless can subsumed under the general title 'the Houck approach'. For an extended discussion on the drawback of this approach see Von Crammon-Taubadel and Meyer (2001) and von Cramon-Taubadel and Loy (1999) or von Cramon-Taubadel (1998). Since the early 1990s the tendency to use more sophisticated methods for analysis of prices transmission was clearly visible. For an example of such alternative techniques see Bacon (1991).

The commonly used way in the 1990s to analyse asymmetric price transmission was based on Error Correction Representation. It can be described as follows. If first-difference stationary time series are cointegrated (there is a long run relationship between them) then the ECM does exist and, in the most general form without the deterministic trend, can be depicted as:

$$\begin{aligned}\Delta \ln(p)_t &= \alpha_0 + \sum_{i=0}^{M_1} \beta_{1,i} * \Delta \ln(c)_{t-i} + \sum_{i=0}^{M_2} \beta_{2,i} * \Delta \ln(ex)_{t-i} + \\ &\sum_{i=1}^{M_3} \beta_{3,i} * \Delta \ln(p)_{t-i} + \pi_1 * \hat{u}_{t-1} + \varepsilon_t; \\ &\hat{u}_{t-1} = \ln(p)_{t-1} - \hat{\alpha} - \hat{\beta}_1 * \ln(c)_{t-1} - \hat{\beta}_2 * \ln(ex)_{t-1}\end{aligned} \tag{1}$$

or:

$$\begin{aligned}\Delta \ln(p)_t &= \alpha_0 + \sum_{i=0}^{M_1} \beta_{1,i} * \Delta \ln(c)_{t-i} + \sum_{i=0}^{M_2} \beta_{2,i} * \Delta \ln(ex)_{t-i} + \sum_{i=1}^{M_3} \beta_{3,i} * \Delta \ln(p)_{t-i} + \\ &\pi_1 * \ln(p)_{t-1} + \pi_2 * \ln(c)_{t-1} + \pi_3 * \ln(ex)_{t-1} + \varepsilon_t\end{aligned} \tag{2}$$

Where *M*s stand for the number of lags over which the shock is felt and ε represents the error term, which should be white noise. It is explicitly assumed that both explanatory variables are exogenous. In (1) and (2) the coefficients π_1 represents the estimate of the speed of adjustment of the downstream prices to shocks in the upstream prices. It is important to add that in equations (1) and (2) it is implicitly assumed that the speed of adjustment is the same in case of positive and negative changes in the upstream prices. Such reservation will show its importance later on.

It is possible to use ECM in testing for asymmetry in two ways. First way used by e.g. Borenstein et al. (1997) and Reilly and Witt (1998) is aimed at looking for the patterns in the short-run response like those depicted in Figure 1.

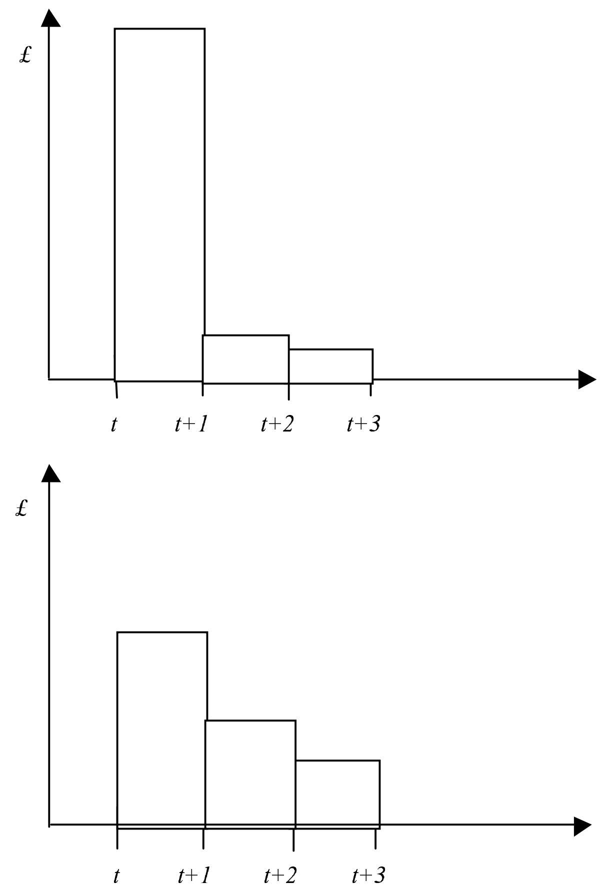

Figure 1 Example of asymmetry in the short-run responses.

Absolute values of changes in the petrol prices after increase and decrease of equal magnitude in crude oil prices at time t. Left panel shows responses to increases in crude oil prices.

It is based on separation of changes in the explanatory variables into positive and negative parts so that (2) becomes:

$$\begin{aligned}\Delta \ln(p)_t &= \alpha_0 + \alpha_1 * trend + \sum_{i=0}^{M_1} \beta_{1,i}^{+} * \Delta^{+} \ln(c)_{t-i} + \sum_{i=0}^{M_2} \beta_{1,i} * \Delta \ln(c)_{t-i} + \sum_{i=0}^{M_3} \beta_{2,i}^{+} * \Delta^{+} \ln(ex)_{t-i} + \\ &\sum_{i=0}^{M_4} \beta_{2,i} * \Delta \ln(ex)_{t-i} + \sum_{i=1}^{M_5} \Delta \ln(p)_{t-i} + \pi_1 * \ln(p)_{t-1} + \pi_2 * \ln(c)_{t-1} + \pi_3 * \ln(ex)_{t-1} + \varepsilon_t \end{aligned} \quad (3)$$

The coefficients on Δ^{+}s provide an estimate of the differential in the effect on net retail petrol prices of increases and decreases in crude oil prices or in the exchange rate. Those coefficients and their standards errors provide the basis for a statistical test of the symmetry hypothesis. The asymmetry is present when the coefficients on Δ^{+}s are statistically different from zero in the final model. The significance level used for testing the null that those coefficients are equal to zero is usually 5%.

The sign of those coefficients determines the direction of asymmetry in the response of net retail petrol prices to changes in the costs of producing petrol. For example, if the coefficient on contemporaneous positive first difference in crude oil price ($\beta^{+}_{1,0}$) is positive and statistically different from zero one can say that immediate response of the net retail petrol prices is greater for increases in the crude oil prices than for decreases. Analogous reasoning can be applied to the interpretation of β_2 and β^{+}_{2} - exchange rate coefficients.

Alternative specification of the ECM may be useful when it comes to estimation of the speed of adjustment to the long-run equilibrium after shocks to upstream prices. One technique was proposed by Granger and Lee (1989). They propose a modification to equation (1) that involves a Wolfram-type segmentation of the error-correction term (i.e. lagged residuals) into its positive and negative components:

$$\begin{aligned}
\Delta \ln(p)_t &= \alpha_0 + \alpha_1 * trend + \sum_{i=0}^{M_1} \beta_{1,i} * \Delta \ln(c)_{t-i} + \sum_{i=0}^{M_2} \beta_{2,1} * \Delta \ln(ex)_{t-1} + \\
&\quad \sum_{i=1}^{M_3} \beta_{3,1} * \Delta \ln(p)_{t-1} + \pi_1^{+} * \hat{u}_{t-1}^{+} + \pi_1^{-} * \hat{u}_{t-1}^{-} + \varepsilon_t; \\
\hat{u}_{t-1} &= \ln(p)_{t-1} - a_0 - \alpha_1 * trend - \beta_1 * \ln(c)_{t-1} - \beta_2 * \ln(ex)_{t-1}; \\
\hat{u}_{t-1}^{+} &= \max(0, \hat{u}_{t-1}); \\
\hat{u}_{t-1}^{-} &= \min(0, \hat{u}_{t-1});
\end{aligned} \tag{4}$$

It is usually recommended to estimate (4) after testing for cointegration in a 'traditional' way without splitting residuals into positive and negative parts. For testing whether the transmission is symmetric and downstream prices adjust equally quickly to increases and decreases in upstream prices von Cramon-Taubadel and Meyer (2001) proposed simple F-test of the null hypothesis of equal coefficients on positive and negative residuals ($H_0{:}\pi^{+} = \pi^{-}$). If the null is rejected and $|\pi^{+}| < |\pi^{-}|$ the downstream prices adjust more rapidly to the increases in the upstream prices than to decreases so that the long-run equilibrium price is reached much faster.

1.2.2 Asymmetric Time-Series Models

While the asymmetric ECM approach, as described by (3) and (4), may be superior to earlier approaches when the price series in question are cointegrated, the Monte Carlo experiments reported by Pippenger and Goering (1993), Balke and Fomby (1997) and Enders and Granger (1998) demonstrated that tests for unit roots and cointegration have some serious drawbacks in the presence of asymmetric adjustment.

They suggest that when adjustment towards new long run equilibrium is asymmetric and the degree of asymmetry is significant, the threshold-autoregressive (TAR) and momentum-TAR models should be used (see Enders and Siklos (1999: 23) for power comparisons.

Those techniques are especially useful when researchers are interested in analysing the process of adjustment of downstream prices after the shocks to upstream prices like those depicted in Figure 2

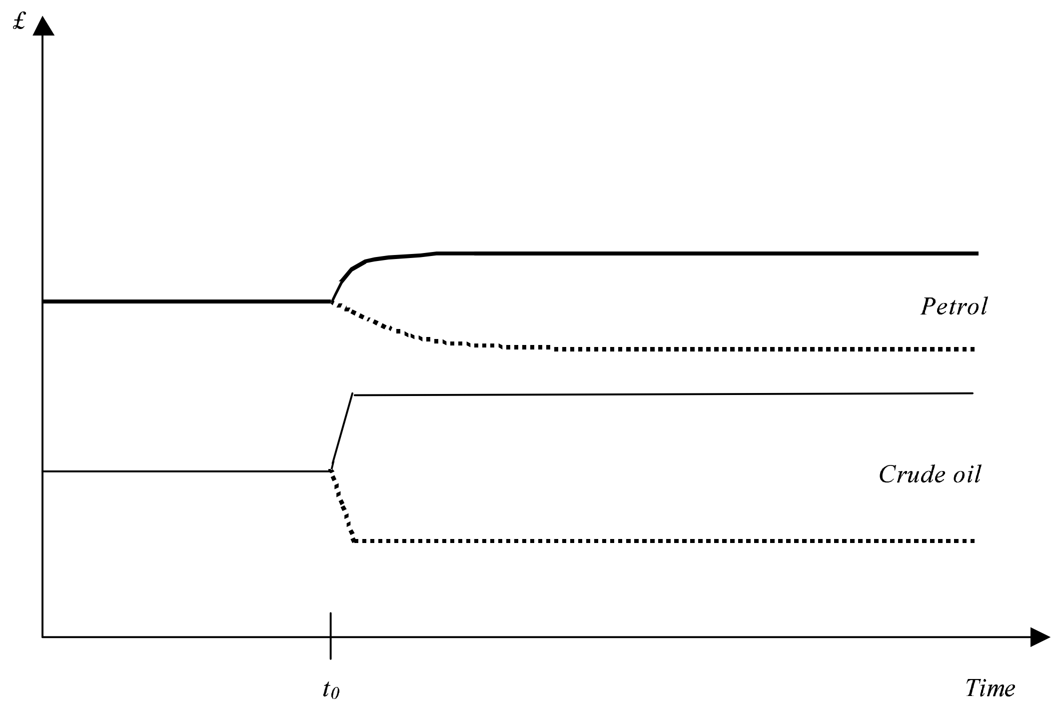

Figure 2 Example of asymmetry in the long-run speed of adjustment.

Process of adjustment of petrol prices after the increase and decrease in crude oil prices at time t_0. Dashed line represents adjustment after a decrease in crude oil prices.

They are aimed at capturing the difference between the speeds of adjustment of downstream prices (petrol prices) to negative (caused by increases in crude oil prices) and positive (caused by increases in crude oil prices) disequilibria. What is more, those techniques allow for taking into account the size of the disequilibrium.

Mentioned alternatives to conventional techniques can be seen as an extension to the standard Engle-Granger (EG) testing strategy. The EG method assumes linearity and symmetric adjustment. In the simplest case, this two-step methodology entails:

1) Using OLS to estimate the long-run relationship:

$$\ln(p)_t = \alpha + \beta * (\ln(c)_t + \ln(ex)_t) + u_t \tag{5}$$

Where all variables are the individual I(1) components, α and βs are the estimated parameters and u_t is the disturbance term that can be correlated.

2) OLS estimation of the π in the regression equation:

$$\Delta\hat{u}_t = \pi * \hat{u}_{t-1} + \varepsilon_t \tag{6}$$

Where ε_t is a white noise disturbance and the residuals from (5) are used to estimate (6). Rejecting the null hypothesis of no cointegration (i.e. accepting the alternative hypothesis: $\pi\in]-2;0[$), implies that the residuals in (5) are stationary with mean zero. As such (5) represents long-run equilibrium - attractor such that its pull is strictly proportional to the absolute value of u_t.

Enders and Siklos (1999) and Enders and Granger (1996) proposed that the following equation (TAR) should be estimated if the degree of asymmetry is significant.

$$\Delta\hat{u}_t = \pi_1 * I(\hat{u}_{t-1}) * \hat{u}_{t-1} + \pi_2 * (1 - I(\hat{u}_{t-1})) * \hat{u}_{t-1} + \varepsilon_t \tag{7}$$

Where I(u^$_{t-1}$) is the indicator function (indicator in short) such that:

$$I(\hat{u}_{t-1}) = \begin{cases} 1 \Leftrightarrow \hat{u}_{t-1} > \tau \\ 0 \Leftrightarrow \hat{u}_{t-1} < \tau \end{cases} \tag{8}$$

The value of the threshold is represented by τ. Setting the threshold at the level different from zero is the only difference to equation (6).

The necessary and sufficient condition for the stationarity of $\{u_t\}$ in such case is: $(\pi_1<0)\wedge(\pi_2<0)\wedge[(1+\pi_1)*(1+\pi_2)<1]$ for given values of τ von Cramon-Taubadel and Meyer (2001:6). If this condition is met, $u_t=0$ can be considered the long-run equilibrium value of the system. The adjustment of the system in such case is $\pi_1 * u_{t-1}$ when $u_{t-1}>\tau$ (e.g. after a sudden and large decrease in the upstream prices) and $\pi_2 * u_{t-1}$ in the opposite case.

Enders and Siklos (1999) tabulated critical values of an F-test (so-called φ* test, or just φ when threshold equals zero) that can be used to test join hypothesis that $\pi_1 = \pi_2 = 0$. If this hypothesis is rejected one can conclude that the cointegration between variables of interest is present. According to Balke and Fomby (1997:4), this test can be seen as additional tool for testing for presence of the long-run relationship.

Enders and Granger (1996) also tabulated critical values for so-called t-max* test (or just t-max when threshold is assumed to be equal zero). It is based on maximum t-statistic of estimate of π_1 or π_2. If the less negative t-statistic exceeds the critical values one can conclude that long run relationship between variables exists. This also can be seen as a sign of cointegration.

If the existence of long-run relationship is established, following Tong (1990), who showed that the OLS estimates of π_1 and π_2 have an asymptotic multivariate normal distribution, it is possible to use standard F-test of the null hypothesis that $\pi_1=\pi_2$ in order to test whether adjustment is symmetric. When null is rejected and an alternative hypothesis that $|\pi_1| < |\pi_2|$ is accepted, one can conclude that downstream prices adjust to the new long-run equilibrium faster when crude oil prices increase than when crude oil prices decrease.

In general, the true value of τ is unknown and needs to be estimated. When possibility of threshold other than zero is real, Enders and Siklos (1999) suggest utilising so-called 'Chan's approach'. Chan (1993) showed that, under some regularity conditions, the least squares (LS) estimator of a stationary ergodic threshold autoregressive model is strongly consistent. After deriving the limiting distribution of the OLS estimator of the threshold parameter he showed that it is N consistent (Chan (1993: 520) so searching over the potential threshold values so as to minimise the sum of squared errors from the fitted model yields a super-consistent estimate of the threshold.

Usual procedure is based on sorting estimated residual series from (5) in ascending order, discarding the largest and the smallest 15% and considering each of the remaining 70% as possible thresholds. The threshold yielding the lowest residual sum of squares in (7) is deemed to be the appropriate estimate. Harris and Silverstone (1999) suggested using the percentiles of the residuals, which is less computation-burdensome.

In (7) the indicator depends on the level of lagged residuals from the long-run equation. Caner and Hansen (1998) and Enders and Granger (1998) suggested an alternative such that threshold depends on the previous period's *change* in residuals so (8) becomes:

$$I(\hat{u}_{t-1}) = \begin{cases} 1 \Leftrightarrow \Delta\hat{u}_{t-1} > \tau \\ 0 \Leftrightarrow \Delta\hat{u}_{t-1} < \tau \end{cases} \tag{9}$$

Such models are called momentum-threshold autoregressive (M-TAR) models in that the $\{u_t\}$ series exhibits more 'momentum' in one direction than the other. Of course setting $\tau=0$ in many cases is natural and is recommended. Including lagged changes in the $\{u_t\}$ sequence, just for TAR models is allowed. What is more, utilizing Chan's procedure for finding estimate of the threshold is possible.

Testing for cointegration involves the F-test (this time named φ*(M) test or φ (M) if threshold equals zero) of the join hypothesis that $\pi_1= \pi_2=0$ and comparing it with the critical values tabulated again in Enders and Siklos (1999). Also t-max tests are allowed (they were called t*-max(M)). Again in is important to add that for reasonable degree of asymmetry conventional tests work reasonably well.

M-TAR model were initially advised in analysis of systems in which adjustment was assumed to come from the outside. For example, consider analysis of the exchange rate in the presence of managed float. In such case the exchange rate authority may want to mitigate large changes in the exchange rate without attempting to influence the long run level of the rate. Enders and Siklos (1999) considered the term structure of the interest rates and concluded that M-TAR model might be used to test for asymmetry in such case because the Federal Reserve might take strong measures to offset shocks to the term structure relationship when such shocks are deemed to indicate increases in inflationary expectations. Shocks that indicate decreases may not be offset so readily.

Caner and Hansen (1998) argued that M-TAR adjustment might be superior to a simple TAR-type adjustment on purely statistical basis. According to their work, if $\{u_t\}$ is a near-unit-root process, setting the indicator using lagged first difference of the residuals (just as in (9)) can perform better than the specification described by (7) with the indicator set according to (8).

The M-TAR model is particularly useful in detecting asymmetrically ‘steep’ movements in a series. The TAR model may be preferred when one is interested in capturing asymmetrically ‘deep’ movements. For the discussion of those properties of the series see Sichel (1993). For our purposes it is enough to say that negative ‘deepness’ (i.e. $|\pi_1| < |\pi_2|$ in (7)) with the indicator set according to (8)) of a series implies negative skewness relative to the mean or trend. The deviation of observation below this mean or trend exceeds in such case the average deviation of observation above. Of course, positive deepness suggests the opposite.

‘Steepness’ of the time series implies that its first differences exhibit negative skewness. In such case the sharp decreases in the series are larger but less frequent than more moderate increases. The figure below shows clearly what mentioned definitions mean:

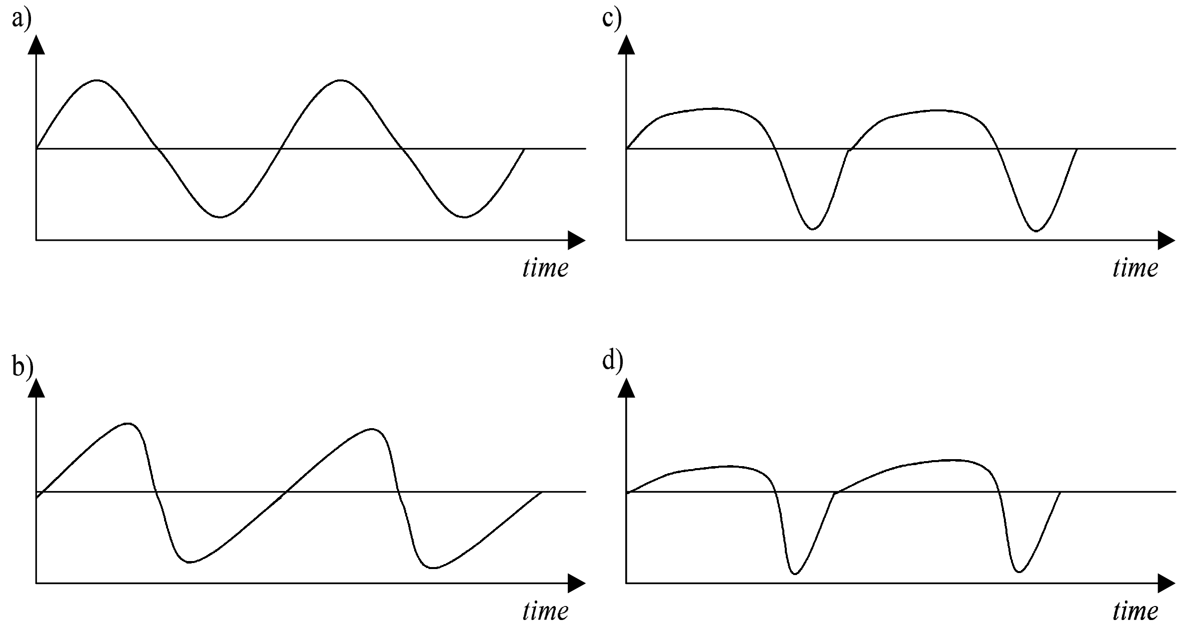

Figure 3 Characteristics of time series:

a) symmetric; b) steep; c) deep; d) steep and deep.

It is important to stress that the character of the M-TAR type asymmetry is fundamentally different from the asymmetry that is tested for by the Houck, ECM and TAR approaches. According to the M-TAR approach, a correction to the margin between prices at different levels of the transmission chain does not depend on the size of this margin at a given point in time but rather on the magnitude and direction of its *change* in the previous period. It is in this sense that M-TAR-type asymmetry is said to exhibit 'momentum'.

2 Cointegration, long-run relationship and its stability

2.1 Data Used, Properties of the Series, Tests for Stationarity

The chain of transformation of crude oil into 4 Star petrol is modelled as a one-stage process. Such approach is consistent with that of Sumner (1990), Manning (1991), Reilly and Witt (1998).

Following Manning (1991), crude oil was considered to be the only significant cost to affect the price paid by drivers. Data on costs of transportation, storing crude oil and petrol was not available on any reliable and constant basis over the sample period. Sample covers last 20 years (January 1982 to December 2001) and it seems that it is the longest time period ever examined for the UK petrol market.

The variables are the monthly series for retail UK petrol price (*p*) net of VAT tax and the excise duty, Brent oil price (*c*) and the dollar/sterling exchange rate (*ex*). Data on typical retail prices of leaded gasoline classified according to the British Standard BS4040 as 4 Star (i.e. 97 octanes) was taken from the Digest of UK Energy Statistics. The data is representative of prices paid on or about the 15th of the month and is based on information provided by oil companies. Prices of crude oil are obtained from DataStream. They represent average FOB spot prices in USD paid at the beginning of the month for one barrel of Brent API 37.4° crude oil on so-called 'Rotterdam market'. This generic term is given to trade in oil product in Northwest Europe and takes its name from the large refining and storage complex in the Antwerp, Rotterdam and Amsterdam area..

The exchange rate is introduced because the prices of crude are expressed in US dollars. As pointed out by Beacon (1991), Reilly and Witt (1998), Galeotti et al. (2002) the exchange rate may be relevant source of asymmetry in non-US countries. Hence the importance of allowing separately for positive and negative changes in exchange rates. Average monthly exchange rate between US dollar and British pound was taken from the DataStream.

The sample size was limited to the last-but-one year due to the changes in the benchmark classification applied by the major players (e.g. Platt's) because of the discussion about the accuracy of WTI and Brent benchmarks in the lights of questions about competition between different crudes.[3]

Simple logarithmic transformation was applied to all the series in order to prevent problems caused by quadratic trends (Banerjee et al. (1993)). Borenstein et al. (1997) pointed out that using logarithms implies that crude-retail margin increases with the price of crude oil. As an alternative he proposed using the raw data. Unfortunately, that implies constant nominal margin in the level equation, which also can be problematic and is not consistent with the findings of the OFT (1998). What is more, some authors (Borenstein et al. (1997: 312)) did admit that with longer samples, which include observations from periods of high and variable inflation, using level equations to test for asymmetries is not justified.

Over the whole sample period the properties of the series were investigated using the typical array of tests. To avoid bias caused by exclusion of the relevant trend in the data generating process the non-standard F-test for the null hypothesis of root in the unit circle and no deterministic trend was applied for all series. In all cases the statistics were smaller than the critical values for 250 observations and 5% significance.

[3] For more details see various papers about the LINK model, created at the Boston University by Robert Kaufmann.

Augmented Dickey-Fuller tests with lag order of twelve in all cases lead to conclusion that unit root may be present in all series. For all series rejecting the null of a unit root at 1% significance level was impossible. Plots of autocorrelation functions for all level variables pointed out slow decay (20 months) characteristic of first-difference stationary series. Alternative tests - CDRW for residuals from regressing on a constant were also performed, they supported the hypotheses of all series being I(1).

With usual care, series could be classified as first-difference stationary. Although in some cases it required using critical values harsher than 5%, classifying them in such way, may have prevented problems with estimation the level equations (like those encountered by Balke et al. (1998)).

To ensure that the prices are indeed I(1), the same array of diagnostic tests was applied to first differences. In all cases null hypothesis of a unit root was rejected at typical significance levels.

2.2 Long-run Relationship, Testing for Cointegration

Following Sumner (1990), Manning (1991), Shin (1994), Reilly and Witt (1996), Borenstein et al. (1997) crude oil prices and the exchange rate were assumed to be exogenous. Application of OLS to the level equation with I(1) series may lead to 'nonsense' or 'spurious' regressions unless variables are cointegrated. This issue was considered so many times that any details might be omitted, only the basic results from the testing procedures are described below. The linear trend was used to capture the effects of other costs on the retail petrol prices (for example real wages were increasing over the sample period). Estimation of the level equation lead to following results:

$$\ln(p)_t = \beta_1 + trend + \beta_2 * \ln(c)_t + \beta_3 * \ln(ex)_t + \varepsilon_t$$
$$\ln(p)_t = 1.3894 + .003729 * \text{trend} + .49011 * \ln(c)_t + .37999 * \ln(ex)_t \qquad (10)$$

						Lagrange Multiplier χ^2 tests				
		R^2	DW	AIC	SBC	Auto - correlation	RESET	ARCH	Hetero-scedascity	JB Normality
Model with linear trend	Value	.70	.40	217.33	210.37	157.2	.58	122.3	3.20	.618
	P-value					.000	.446	.000	.300	.003

As a first test for cointegration the ADF test was applied to residuals obtained from estimation as in (6). The optimal lag length was chosen in accordance with information criteria. SBC, HQC and AIC criteria indicate that typical DF statistic without lags is appropriate. Its value (-5.1729) exceeds the critical value at 5% significance level (-4.1725) allowing us for rejecting the null of residuals being I(1). Also CRDW statistic exceeds critical value for 4 regressors i.e. 0.3. indicating cointegration.

The critical values based on response surfaces also support the view that the long run relationship between prices exist. Regressing changes in residuals on residuals lagged one period as in (6) gave t-statistic equal to -5.1514. The critical values at 1% and 5% significance level are -4.74551, -4.16963, accordingly. Restricted and unrestricted ECMs were also used to test for cointegration. Using lagged residuals from (10), equation (1) was estimated:

$$\Delta\ln(p) = \alpha + \beta_1 * \Delta\ln(c)_t + \beta_2 * \Delta\ln(ex)_t + \pi * \hat{u}_{t-1} + \varepsilon_t$$

$$\Delta\ln(p) = .033287 + .27464 * \Delta\ln(c)_t + .20636 * \Delta\ln(ex)_t - .18176 * \hat{u}_{t-1} \qquad (11)$$

[.094463][8.1126] [1.5272] [4.9983]

						Lagrange Multiplier χ^2 tests				
		R^2	DW	AIC	SBC	Auto - correlation	RESET	ARCH	Hetero-scedascity	JB Normality
(11)	Value	.268	1.89	357.73	350.78	16.36	.125	7.12	2.17	81.88
	P-value					.175	.724	.856	.145	.0

The model above passes the array of diagnostic tests. The fact that residuals do not pass the JB test does not affect the efficiency of the OLS estimation. The t-statistic on the lagged residuals suggests that they are significantly different from zero, taking into consideration the critical values at 5% and 1% significance level as given by McKinnon (1991). It also supports the hypothesis that all series are cointegrated.

As the next step in testing for linear cointegration unrestricted ECM in the form given by (2) was employed.

$$\Delta\ln(p)_t = \alpha + \beta_1 * \Delta\ln(c)_t + \beta_2 * \Delta\ln(ex)_t + \pi_1 * \ln(p)_{t-1} + \pi_2 * \ln(c)_{t-1} + \pi_3 * \ln(ex)_{t-1} + \varepsilon_t$$

$$\Delta\ln(p)_t = .24115 + .27560 * \Delta\ln(c)_t + .20638 * \Delta\ln(ex)_t \qquad (12)$$

[.3.4946] [7.9263] [1.4856]

$$- .16974 * \ln(p)_{t-1} + .084461 * \ln(c)_{t-1} + .067155 * \ln(ex)_{t-1}$$

[N/A] [N/A] [N/A]

						Lagrange Multiplier χ^2 tests				
		R^2	DW	AIC	SBC	Auto - correlation	RESET	ARCH	Hetero-scedascity	JB Normality
(12)	Value	.263	1.90	354.93	344.5	17.54	.375	7.63	1.66	84.69
	P-value					.130	.540	.813	.197	.000

Non-standard Pesaran F-tests for join insignificance of all lagged values equals to 7.6811 and exceed the critical value for 3 variables, intercept and a trend for 5% and 1% significance level (i.e. 5.972 and 7.584, respectively), indicating cointegration.

All tests above indicate that long-run relationship between variables under scrutiny really exists. This rules out possibility of 'spurious regression'. Since it was proven that with cointegration OLS estimators are not only consistent but also N consistent, we can treat βs from estimations above as reliable estimates of real parameters.

The point is that since those estimates are superconsistent OLS estimation gives residuals that represent unpredictable shocks to the system. This gives a strong base for several tests for asymmetric adjustment of the system that will be used in the next part.

One remaining problem is connected with the stability of the parameters. The CUSUM test was applied to the whole sample. As the picture below shows the null hypothesis of the parameter stability over the sample period must be rejected at 5% significance level.

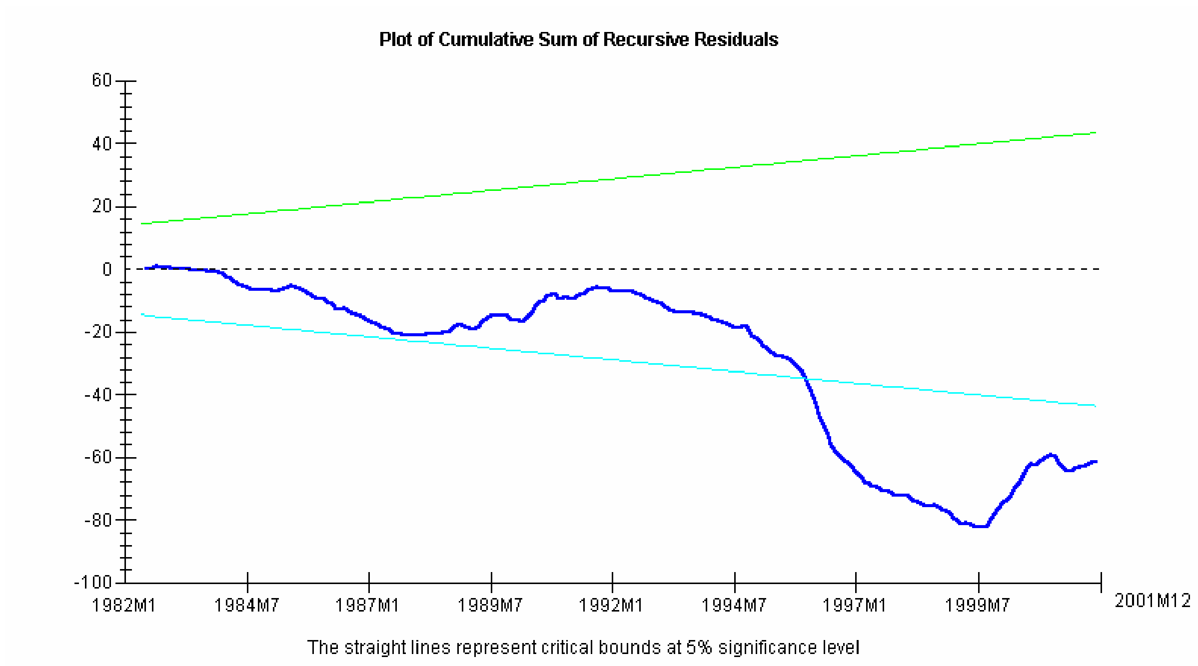

Figure 4 CUSUM test

Straight lines represent the critical values at 5%

As the next step to ensure if the changes were caused by one-time deviation or change in the parameters the CUSUMSQ test was applied. Result given below suggests instability in the long-run relationship around 1996.

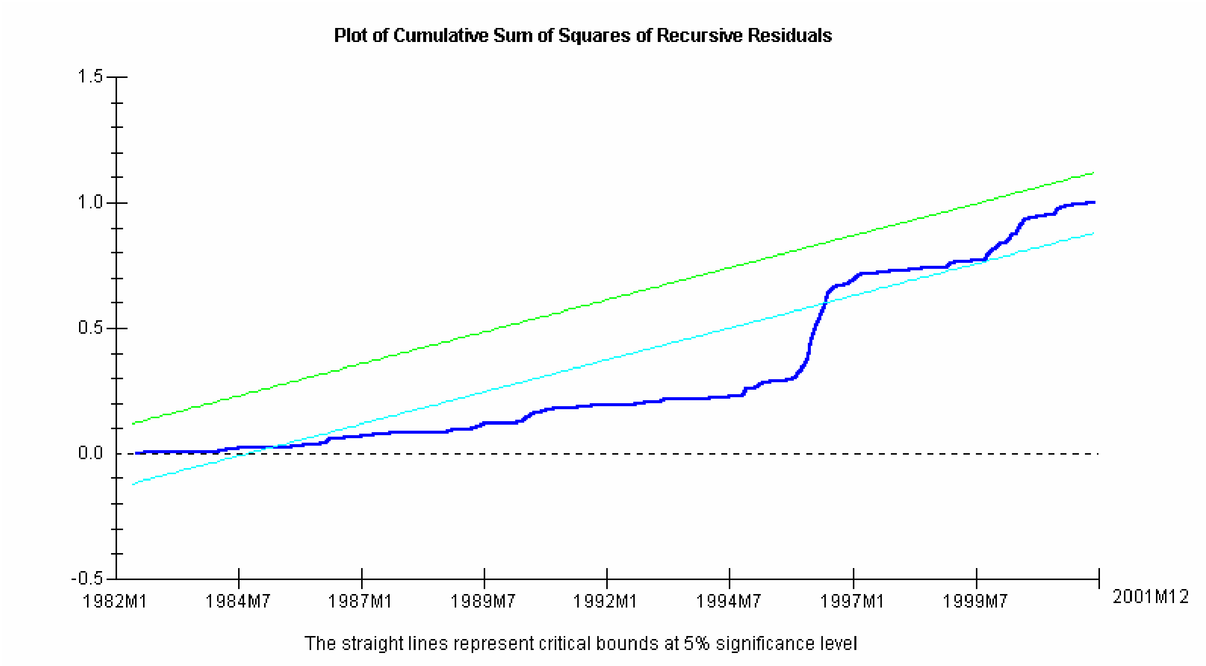

Figure 5 CUSUMSQ test

Straight lines represent the critical values at 5%

Also the behaviour of the standard errors of recursive regression was examined. As Figure 6 shows, around 1996 there was a sudden increase in the value of standard error.

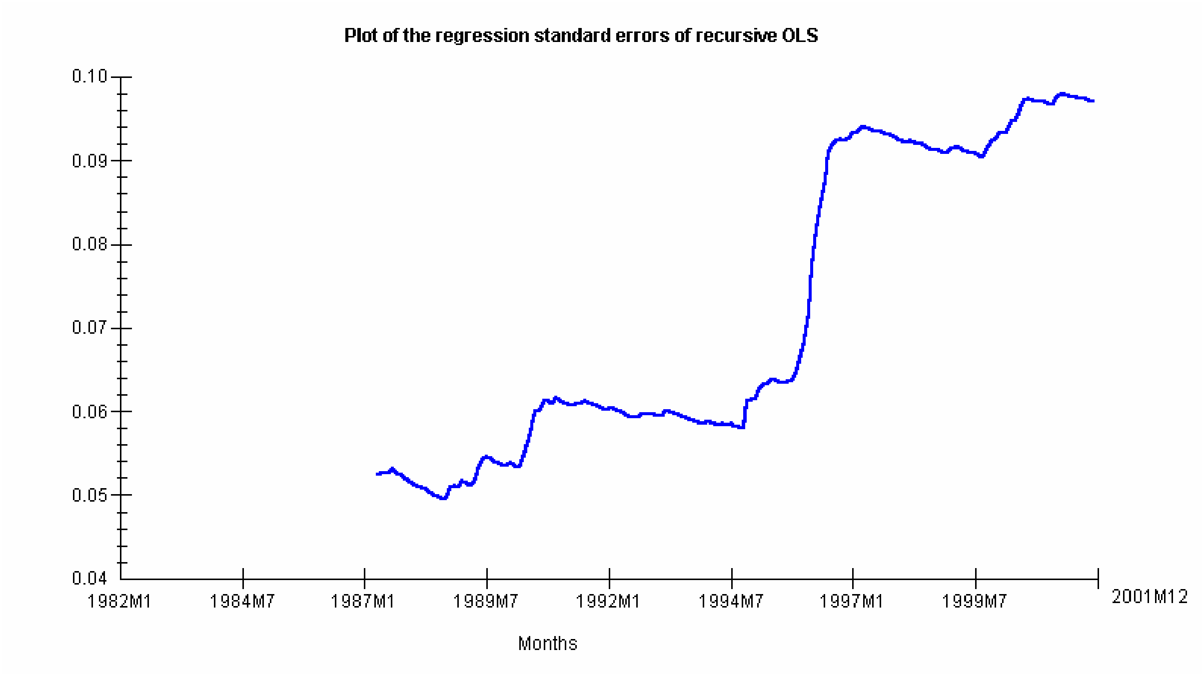

Figure 6 Standard errors of recursive regression.

To check thoroughly the properties of (10) one-step-ahead Chow test was used. The model was calibrated in the first 10 years (the standard error for that period was equal to .059453). The test statistic scaled by the 1% critical value for the χ^2 distribution in Figure 7 is represented by the blue line. At the end of the sample instability is most likely to be caused by the introduction of two different taxes.

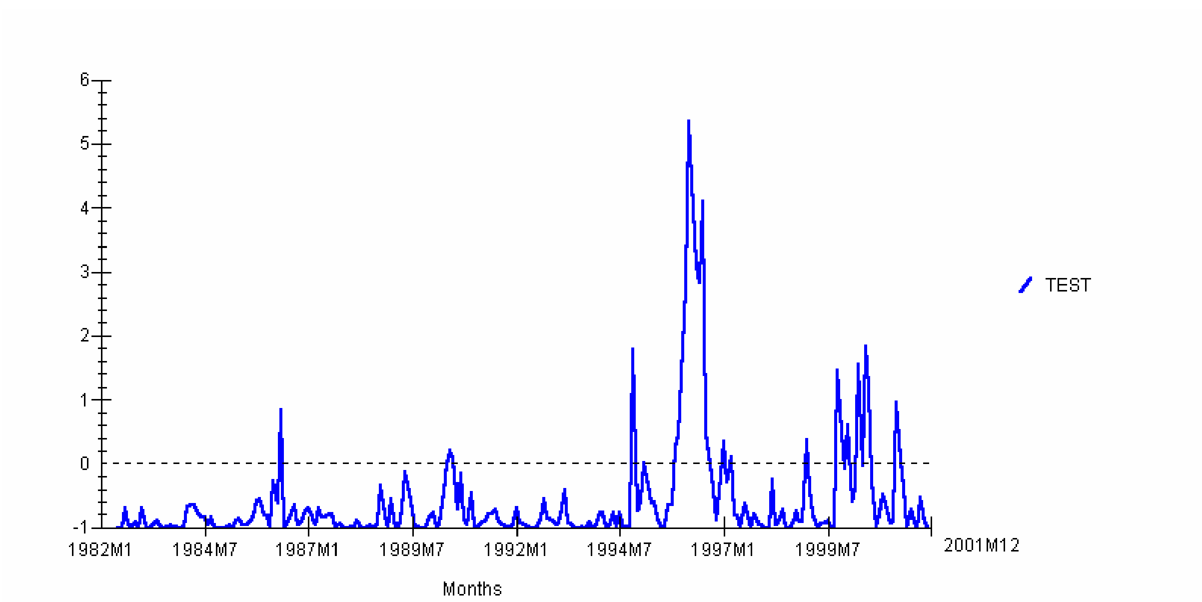

Figure 7 One-step-ahead Chow test

The most probable explanations of this instability are connected with the Pricewatch programme aimed at competing with supermarkets which started to sell petrol in the mid

1990s. Pricewatch programme consisted of granting by the big refiners a condition of payment to the retail station in order to allow them to lower the prices and compete with supermarkets. For the discussion about this programme and its impact on the UK market see the OFT report (1998) or Williamson and Taylor (2000).

3 Testing for Asymmetry

3.1 Asymmetry in the Short-run Adjustment

To test whether the short-run response to changes in the crude oil prices is symmetric the ECM model as described by equation (3) was estimated over the whole sample. Following Banerjee et al. (1993), the Hendry's 'general-to-specific' procedure was used to determine the appropriate lag structure. The significance level at each step of the procedure was equal to 0.5%.

For testing the null hypothesis of symmetry in short-run response conventional critical values were used i.e. the t-statistics on the coefficient on positive first differences *($\beta^+ s$)* were compared to conventional critical values at 5% significance level. After the testing down the following model was obtained:

$$\begin{aligned}\Delta \ln(p)_t &= \alpha_1 + \alpha_2 * trend + \beta^+{}_{1,0} * \Delta^+ \ln(c)_t + \beta_{1,0} * \Delta \ln(c)_t + \beta_{1,1} * \Delta \ln(c)_{t-1} + \\ &\beta^+{}_{2,0} * \Delta^+ \ln(ex)_t + \beta_{2,1} * \Delta \ln(ex)_{t-1} + \pi_1 * \ln(p)_{t-1} + \pi_2 * \ln(ex)_{t-1} + \pi_3 * \ln(c)_{t-1} + \varepsilon_t \end{aligned} \tag{13}$$

After further analysis several restriction on the coefficients were imposed, they were all supported by appropriate test statistics. The final model was:

$$\begin{aligned}\Delta \ln (p)_t = {} & \underset{[3.9333]}{.23209} + \underset{[.54526]}{.0002869} * trend + \underset{[2.2934]}{.16086} * \Delta^+ \ln(c)_t + \underset{[5.7391]}{.16357} * (\Delta \ln (c)_t + \Delta \ln(c)_{t-1}) + \\ & \underset{[1.9265]}{.42302} * \Delta^+ \ln(ex)_{t-1} + \underset{[2.5067]}{.34391} * \Delta \ln(ex)_{t-1} - \underset{[N/A]}{.15244} * \ln(p)_{t-1} + \underset{[N/A]}{.067310} * (\ln(ex)_{t-1} + \ln(c)_{t-1})\end{aligned} \tag{14}$$

						Lagrange Multiplier χ^2 tests				
		R^2	DW	AIC	SBC	Auto - correlation	RESET	ARCH	Hetero-scedascity	JB Normality
(14)	Value	.341	1.94	364.87	350.98	14.809	.0665	2.3686	.135	174.05
	P-value					.252	.796	.999	.712	.000

According to the results from (14) 1% increase in the prices of crude oil at *t* causes the prices of petrol to increase by .323% (.323 = .160+.163) at *t* and by further .163% in the next period. In contrast to that, 1% decrease in the prices of crude oil causes fall by only .16% at *t* and by similar amount in the next period (ceteris paribus).The first estimate is consistent with the estimate of .224% by Reilly and Witt (1998) and that by Galeotti et al. (2002) (.198.%). Sumner's (1990) estimate of the immediate response to 1% increase in crude oil prices equal to .38% is quite close to the results obtained.

It is visible that the movements in the exchange rate also seem to cause asymmetric short-run response. The fact that the coefficient on $\Delta log(ex)_t$ was not significantly different from zero at 0.5% suggests that favourable change in the exchange rate is not passed to prices of petrol immediately. In the next period after 1% favourable change prices of petrol fall by approximately 0.34% (ceteris paribus). The situation with unfavourable changes is quite different – at t there is a transmission of 1% change equal to 0.42% and it is also present in the next period.

Galeotti et al (2002) also found that short-run response to a favourable change in the exchange rate is insignificantly different from zero. Their estimate of response to unfavourable exchange rate is equal to 0.732 and is close to that obtained by Reilly and Witt (1998). Estimate reported above is lower, but also significantly different from zero.

The estimates on the levels' variables in (14) can also be used to numerically solve for the long-run relationship. In a long-run steady-state, the short-run effects go to zero, ($\forall i,j \in N$ $ln(p)_{t-i}= ln(p)_{t-j}$ and $ln(ex)_{t-j} =ln(ex)_{t-i}$) and it is possible to re-express (14) and solve for $ln(p)_t$ to obtain:

$$\ln(p)_t = 1.5225 + .001882052*trend + 0.441551*(\ln(ex)_t + \ln(c)_t) \tag{15}$$

The long-run estimate reflecting the effect of crude oil prices expressed in pounds is close to the estimate of 0.58 reported by Reilly and Witt (1998). The long-run coefficient reported by Manning (1991) equal to 0.3055 is much lower than result above. The long-run estimate of unity reported in Bacon (1991) is entirely not in agreement with findings. The Wald test of such null hypothesis supported this view with all force – null of full passthrough was definitely rejected.

Figure 8 Plot of recursive coefficients for asymmetry caused by the crude oil prices

The interesting point is that results of recursive regression reported by Reilly and Wit t (1998) show the same pattern with sudden increase at the beginning of the Gulf war and fairly stable values till 1995.

3.2 Testing for Asymmetry in the Long-run Speed of Adjustment

3.2.1 ECM Representation

To assess the impact of the Pricewatch programme on asymmetry in the adjustment process the rolling OLS with the window size equal to 5 years was applied to the model given by (4) . Steps were as follows:

1) Ordinary rolling regression in the following form:

$ln(p)_t = \alpha_0 + \alpha_1 * trend + \beta_1 * ln(c)_t + \beta_2 * ln(ex) + \varepsilon_t;$

2) Saved rolling coefficients were used to compute rolling residuals, which can be seen as crude proxies for the shocks to the system;
3) Another rolling regression of changes in the residuals on lagged residuals divided into positive and negative parts. The figures below show the values of the rolling coefficients obtained in this way.

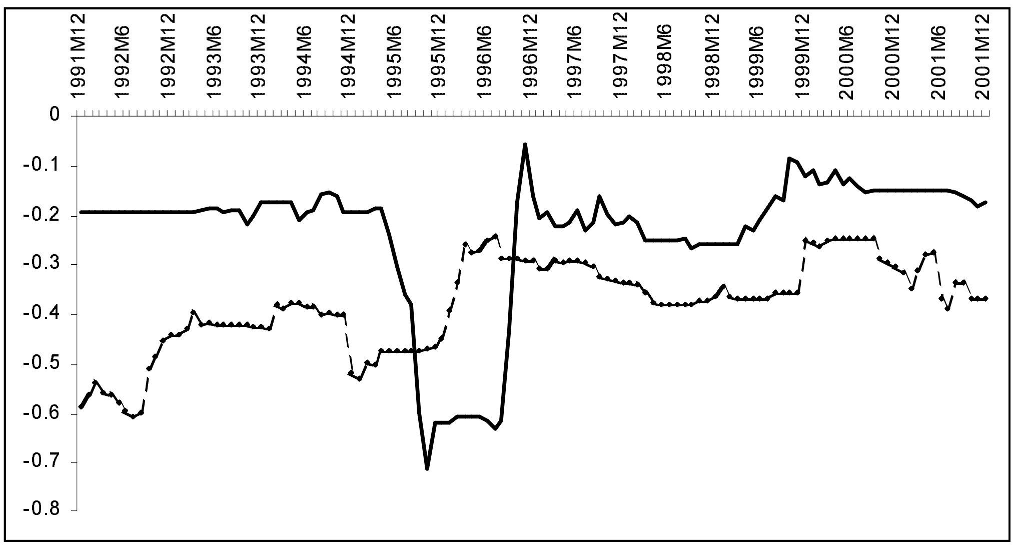

Figure 9 Recursive estimates of speeds of adjustment

It is clear that increased competition caused by entry of the new station owned by supermarkets and the Pricewatch campaign aimed at competing for the market share with them increased the speed in which the agents responded to shocks caused by falling prices of crude oil and favourable exchange rate movements. It is also visible that change was only temporary. To asses for it 5 impulse dummies were used. The estimation of the long-run equation with dummies gave:

$$\ln(p)_t = \alpha_1 + \alpha_2 * trend + \beta_1 * \ln(c)_t + \beta_2 * \ln(ex)_t + \sum_{i=1}^{5} \vartheta_i * D_i + \varepsilon_t \qquad (16)$$
$$\ln(p)_t = 1.4008 + 0.0004374 * trend + .48936 * \ln(c)_t + .40411 * \ln(ex)_t$$

						Lagrange Multiplier χ^2 tests				
		R^2	DW	AIC	SBC	Auto - correlation	RESET	ARCH	Hetero- scedascity	JB Normality
(16)	Value	.76	.54	235.5	219.84	135.38	.0027	4.1995	1.980	1.8984
	P-value					**.000**	.958	.000	.159	.387

Introducing the dummies into (10) did not change the values of the estimates but resulted in the normality of residuals as indicated by JB test statistics. All tests for cointegration were repeated and results were unchanged. Excluding only 5 observations allowed to test for asymmetric patterns in the transmission with greater confidence.

Next the techniques proposed by Granger and Lee (1989) were used. Residuals from (16) were used to test for asymmetry in the adjustment to the long-run equilibrium just as proposed by equation (4). The results are given below

$$\Delta\ln(p)_t = \alpha + \beta_1 * \Delta\ln(c)_t + \beta_3 * \Delta\ln(ex)_t + \pi^+ * \hat{u}^+_{t-1} + \pi^- * \hat{u}^-_{t-1} + \varepsilon_t$$
$$\Delta\ln(p)_t = \text{-}.0064509 + .27928 * \Delta\ln(c)_t + .21439 * \Delta\ln(ex)_{t-1}$$
$$[1.1578] \qquad [8.3636] \qquad [3.9951] \qquad (17)$$
$$\text{-}.12740 * \hat{u}^+_{t-1} - .\text{-}.32393 * \hat{u}^-_{t-1} + \varepsilon_t$$
$$[1.6968] \qquad [4.3765]$$

						Lagrange Multiplier χ^2 tests				
		R^2	DW	AIC	SBC	Auto - correlation	RESET	ARCH	Hetero- scedascity	JB Normality
(17)	Value	.297	1.85	361.61	352.92	17.349	1.937	5.972	1.010	100.424
	P-value					.137	.164	.917	.315	.000

Testing for the symmetry in the adjustment speed was based on simple F test of the null hypothesis of equality of $\pi^+$and π^-. Its value (2.406) allows rejecting null of equal speed of adjustment only at 15% significance level. What is more results of (17) are not consistent with the requirements set by Cook et al. (1998) since the coefficient on positive residuals is significantly different from zero only at 10% not at standard 5% significance level.

In order to get more information about the dynamics of the system the restricted ECM was enriched by lagged explanatory variables. For determining the appropriate lag length the usual Hendry's approach was used.

The most generous model estimated included lags up to order of 6 in case of exchange rate and crude oil prices and 5 lagged differences in retail prices. Therefore to get the overall significance level equal to 5% at each stage of testing-down procedure the significance level for each lagged variable was set at 0.003.

After the 'testing down' procedure the following model was obtained.

$$
\begin{aligned}
&\Delta \ln(p)_t = \alpha + \beta_1 * \Delta \ln(c)_t + \beta_2 * \Delta \ln(c)_{t-1} + \beta_3 * \Delta \ln(ex)_{t-1} + \pi^+ * \hat{u}^+_{t-1} + \pi^- * \hat{u}^-_{t-1} + \varepsilon_t \\
&\Delta \ln(p)_t = \text{-}.0050829 + .27892 * \Delta \ln(c)_t + .12725 * \Delta \ln(c)_{t-1} + .40727 * \Delta \ln(ex)_{t-1} \\
&\qquad [.95035] \qquad [8.6050] \qquad [3.9465] \qquad [3.1543] \\
&\qquad \text{-}.11119 * \hat{u}^+_{t-1} - .\text{-}.27891 * \hat{u}^-_{t-1} + \varepsilon_t \\
&\qquad [1.5443] \qquad [3.8907]
\end{aligned} \tag{18}
$$

						Lagrange Multiplier χ^2 tests				
		R^2	DW	AIC	SBC	Auto - correlation	RESET	ARCH	Hetero-scedascity	JB Normality
(18)	Value	.358	1.92	369.98	359.56	13.84	1.559	1.927	83025	160.2
	P-value					.311	.212	1.000	.362	.000

Results from (18) also suggest that the speed of adjustment differs according to the nature of the shock to the system i.e. whether upstream prices increased or decreased. However the empirical test allowed for rejecting the null of the same speed of adjustment at 10%. The most likely explanation was that the model used was relatively simple and separating the residuals according to their sign only was naive. More sophisticated methods begged to be used.

3.2.2 TAR and M-TAR models

Techniques described by Enders and Siklos (1999) were not developed to deal with three and more variables, therefore impact of exchange rate could not be evaluated separately.

It was possible to overcome this problem by analysing the price transmission between prices of crude oil expressed in British pounds and net retail prices of 4 Star petrol. Since the price of Brent crude expressed in pounds equals product of exchange rate and price expressed in US dollars, the logarithmic transformation of crude prices in pounds is simply sum of logs of price in dollars and log of exchange rate.

Using prices in domestic currency will be allowed only if in the long run equation coefficients on those two variables are equal to each other. The Wald test for testing the null of equality of β_3 and β_4 was used. The value of the test statistic was 2.3304 with associated p-value equal to .127. It is lower than the critical values and it does not allow for rejecting the null of equal impact at conventional significance levels.[4]

Therefore, it is possible to assume that in the level equation the values of the coefficient on crude oil prices and exchange rate is the same so they can be replaced by the new variable, sum of those two.

The same array of tests both diagnostic and for cointegration was applied to the new model. The results are given below.

$$\ln(p)_t = \alpha_0 + \alpha_1 * trend + \beta_2 * (\ln(c)_t + \ln(ex)_t) + \varepsilon_t$$

$$\ln(p)_t = 1.4864 \quad + \quad .003600 * \text{trend} \quad + 47251 * \ (\ln(c)_t + \ln(ex)_t) \tag{19}$$

						Lagrange Multiplier χ^2 tests				
		R^2	DW	AIC	SBC	Auto - correlation	RESET	ARCH	Hetero-scedascity	JB Normality
(19)	Value	.706	.38	217.15	211.94	159.47	.1137	16.961	2.62	.70
	P-Value					.000	.736	.000	.347	.004

The properties of the new series were thoroughly examined. The ADF test with lag length indicated by information criteria gave the test statistic equal to -2.5364 for levels and -14.6288 for the first differences. Therefore, it is possible to conclude that new series is also I(1). Typical array of tests for cointegration was also used. What is expected, they show clear signs of cointegration.

The simplest TAR model in a form given below was estimated over the sample period. The residuals were divided according to the basic indicator i.e.

$$I(u_{t-1}) = \begin{cases} 1 \Leftrightarrow u_{t-1} > 0 \\ 0 \Leftrightarrow u_{t-1} < 0 \end{cases}$$

[4] Inefficiency of estimation caused by the autocorrelation present was overcome by the Wald tests based on the Newey-West adjusted variance-covariance matrix with different weights. Using the version with the equal weights and truncation point set at 25% of the sample the test statistic equals .74757 which is less than critical value for χ^2 with one d.o.f. (3.84). The Wald test with similar truncation point but different weights i.e. Barlett, Tukey, Parzen gave similar results i.e. .58105, .55708, and .49758, respectively. Changing the truncation point did not affect the result.

Threshold set at 0 reflects a implicit assumption that the system under scrutiny behaves differently when upstream prices increase and decrease. The issue of the magnitude of the shocks and its impact on the speed of adjustment will be discussed later

$$\Delta\hat{u}_t = \beta^+ * I(u_{t-1}) * \hat{u}^+_{t-1} + \beta^- * (1 - I(u_{t-1})) * \hat{u}^-_{t-1} + \varepsilon_t$$
$$\Delta\hat{u}_t = \beta^+ * \hat{u}^+_{t-1} + \beta^- * \hat{u}^-_{t-1} + \varepsilon \qquad (20)$$
$$\Delta\hat{u}_t = -.21648 * \hat{u}^+_{t-1} - .32466 * \hat{u}^-_{t-1}$$
$$[3.4599] \qquad [5.1982]$$

						Lagrange Multiplier χ^2 tests				
		R^2	DW	AIC	SBC	Auto - correlation	RESET	ARCH	Hetero- scedascity	JB Normality
(20)	Value	.141	2.12	332.18	328.71	17.33	1.2948	6.3306	.42671	235.7718
	P-value					.138	.255	.899	.514	.000

According to the critical values tabulated by Enders and Granger (1998) the value of the maximum t-statistic (i.e. -3.4599) gives a basis for rejecting the null of no threshold cointegration between variables at 5% and 1% significance level (appropriate critical values are for 250 observations –2.12 and –2.53 respectively. F-test for the join insignificance of both lagged variables (φ test) equals 38.99 and clearly exceeds critical values -5.87 and -8.04, at 5% and 1% significance level. The hypothesis of equal speed of adjustment could not be rejected by the standard F-test. The test statistic equal to 1.4976 is lower than critical values at 5%.

A clear drawback of modelling asymmetry by using specification described by (20) is that it imposes that characteristics of the adjustment differ only depending on the signs of the disequilibria. The more realistic assumption should be that speed of adjustment might differ if the disequilibria are significant. There are many reasons to predict that this may be the case.

For example, if upstream prices increase by a lot - so that margins at all stages of modelling approach zero, agents are more likely to adjust prices toward new long-run equilibrium very fast. But when upstream prices decrease significantly agents may postpone lowering prices and enjoy bigger margins for longer.

Therefore, consistent TAR model as described by equation (8) was employed to find the value of the threshold is based on the one used by modification of method used by Harris and Silverstone (1999). The analysis suggested that the value of the threshold should be set at –0.05969, in such case RSS equals .83589 which is lower than that for model described by(20) (.85388).

The model with such threshold is described below:

$$\Delta\hat{u}_t = \beta_1 * I(u_{t-1}) * \hat{u}_{t-1} + \beta_2 * (1 - I(u_{t-1})) * \hat{u}_{t-1} + \varepsilon_t$$

$$\Delta\hat{u}_t = -.17884 * I(u_{t-1}) * \hat{u}_{t-1} - .38302 * (1 - I(u_{t-1})) * \hat{u}_{t-1} \qquad (21)$$

$$\qquad [3.0256] \qquad\qquad [5.8579]$$

						Lagrange Multiplier χ^2 tests				
		R^2	DW	AIC	SBC	Auto - correlation	RESET	ARCH	Hetero-scedascity	JB Normality
(21)	Value	.155	2.10	334.111	330.6346	17.5943	.11303	5.971	.32377	244.6854
	P-value					.129	.737	.918	.599	.000

Consistent version is clearly superior to the 'ordinary' TAR. It is indicated by the values of SBC and AIC and by the test statistics. The t-max* statistic in (21) equals –3.0256 and exceeds the critical values at 5% and 1% (-1.84 and –2.31, respectively) again indicating cointegration. The F-test for the equality of coefficients has a value equal to 5.366 with p-value equal to 0.021 allowing us to rejecting the null of equal speed of adjustment to positive and negative shocks.

The conclusion from all TAR models above is that the system exhibits slower adjustment to the positive disequilibria that exceed the threshold values. In other words system seems to adjust to the equilibrium in the same manned when small increases and decreases in upstream prices occur but if increases exceed the critical value the adjustment to the equilibrium is much faster.

It is entirely consistent with the results of survey among retailers (mainly independent retailers) in OFT Report (1998: 51). Majority of them claimed that they are reluctant to match their competitors if they operate at very low margin and/or that they feel that their margins are too small. It may suggest that they will use the opportunity to raise margins, by postponing lowering the prices when upstream pries fall but will increase prices very rapidly whenever upstream prices increase. That is why in the latter case new long-run equilibrium is achieved much faster. Their responses clearly indicate that the size of the margin matters – if it is too low they are reluctant to lower the prices to the new long-run equilibrium level. Figure 10 shows simulated adjustment over 24 months after positive and negative shocks which caused 15% discrepancy between current prices and their long-run equilibrium levels.

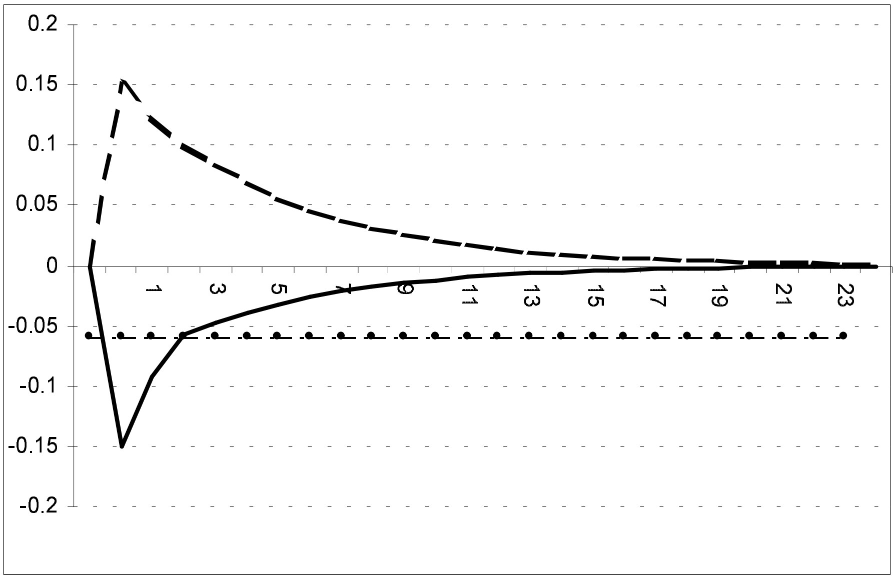

Figure 10 Adjustment to the long-run equilibrium

Dashed line corresponds to the positive disequilibrium caused for example by decrease in the crude oil prices.

This finding can be a basis for the next model in which behaviour of the system depends not only on the direction of the disequilibria but also changes in them affect the adjustment to the long-run equilibrium. Estimation of M-TAR model gave following results.

$$\Delta\hat{u}_t = \beta_1 * I(\Delta\hat{u}_{t-1}) * \hat{u}_{t-1} + \beta_2 * (1 - I(\Delta\hat{u}_{t-1})) * \hat{u}_{t-1} + \varepsilon_t$$

$$\Delta\hat{u}_t = -.15352 \underset{[2.4128]}{*} I(\Delta\hat{u}_{t-1}) * \hat{u}_{t-1} - .38019 \underset{[6.2991]}{*} (1 - I(\Delta\hat{u}_{t-1}) * \hat{u}_{t-1} \qquad (22)$$

						Lagrange Multiplier χ^2 tests				
		R^2	DW	AIC	SBC	Auto - correlation	RESET	ARCH	Hetero- scedascity	JB Normality
(22)	Value	.167	2.12	333.33	329.86	17.895	.38428	6.7575	.08316	247.6153
	P-value					.119	.535	.873	.773	.000

The t-max(M) statistic exceeds the critical values at 5% -1.99 and is very close to the critical value at 1% (i.e. –2.45). The conventional F-test rejects the null of equal speed of adjustment is rejected at 5% and 1% with the statistic equal to 6.68. As above shows the system adjusts to approximately 15% of the disequilibrium if it follows sudden decrease in the prices of crude oil but 38% of disequilibrium caused by increase in the prices of crude oil of similar magnitude is eliminated over one year.

In light of properties of the M-TAR model described in chapter 1.2, it indicates that sudden decrease in the margin prompts sharp answer but when margins are unchanged or increased adjustment to the new long run equilibrium takes a little bit longer. One can conclude that in the transmission chain agents delay lowering prices when their margins are stable or increasing but changes in the upstream prices that reduce the margin are passed fully onto the consumer.

The findings seem to be consistent with the results of the OFT survey (1998:52) Almost half of the retailers responded that they would not match the competition in lowering prices if it means reducing their margin. Their answers also suggest the *change* in the margin is as important as its size.

Also the consistent version of the M-TAR model was estimated. The aim of this model was to account for the possibility that behaviour of the system may depend not only on the sign of the change in the disequilibrium but also on its magnitude. Therefore I estimated the consisted version of the M-TAR model using the first differences of the residuals.

Indicator was set according to (9). The method of finding the threshold was analogous to that used before in case of consistent TAR model. The estimated model that gave the lowest residual sum of squares (0.82965 compare to 0.83233 in case of ordinary M-TAR) had a threshold equal to -.0034.

$$\Delta\hat{u}_t = I(\Delta\hat{u}_{t-1}) * \beta_1 * \hat{u}_{t-1} + (1 - I(\Delta\hat{u}_{t-1})) * \beta_2 * \hat{u}_{t-1} + \varepsilon_t$$

$$\Delta\hat{u}_t = -.15026 * I(\Delta\hat{u}_{t-1}) * \hat{u}_{t-1} - .38923 * (1 - I(\Delta\hat{u}_{t-1}) * \hat{u}_{t-1} \qquad (23)$$

$$\qquad [2.3987] \qquad\qquad [6.3762]$$

						Lagrange Multiplier χ^2 tests				
		R^2	DW	AIC	SBC	Auto - correlation	RESET	ARCH	Hetero- scedascity	JB Normality
(23)	Value	.164	2.12	333.71	330.24	17.95	.35919	6.6967	.05980	251.0294
	P-value					.117	.549	.877	.807	.000

The results of this model are consistent with those from normal M-TAR model. However, as indicated by AIC and SBC it can be seen as superior. The asymmetric response is more visible and the Wald statistic allows for rejecting the null of equal coefficient at 1% significance level (7.4643 with p-value equal to .006). The fact that in this case threshold is different from zero and its negative may suggest that change in the margins must exceed specified value in order to prompt fast response of the agents. If disequilibria changed by less than –.0034, adjustment is weak, but if the change exceeded this threshold (i.e. there was a sudden drop in the value of the disequilibrium – corresponding for example to sudden increase in the upstream price) response of agents is much quicker and the disequilibrium is eliminated much faster.

Conclusions

Econometric analysis confirmed that asymmetric response of petrol prices to changes in the upstream prices exists. Thus the belief held by the drivers is confirmed. Detailed analysis based on the latest developments in the TAR and M-TAR models confirmed that the size and changes in the market margin determine the size of the asymmetry. This may support the view that the behaviour of the agents is responsible for the characteristics of the process of price transmission. This finding, although intuitive, was not confirmed so far by any formal study, perhaps because the techniques were not developed sufficiently.

The analysis described above allows for a closer look at the nature of asymmetry, and could be sign as a sign that asymmetry is a inherent feature of the real petroleum markets.

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