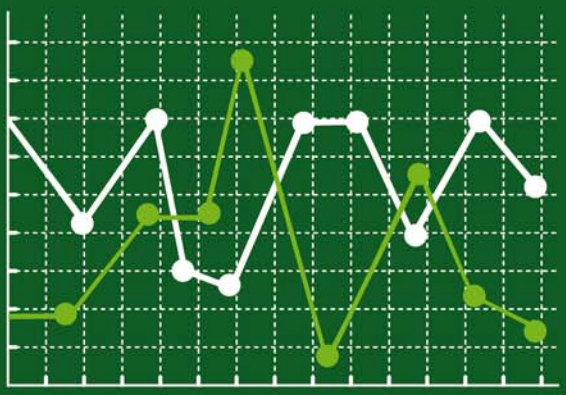
a/ working paper

WP 2014/07 (December)

Long- and short-run price asymmetries in the Italian energy market: the case of gasoline and heating gasoil

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LONG- AND SHORT-RUN PRICE ASYMMETRIES IN THE ITALIAN ENERGY MARKET: THE CASE OF GASOLINE AND HEATING GASOIL

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Abstract: Using monthly data from 1994 to 2012 we study the long-run relation between the pre-tax retail prices of petrol and heating gasoil with crude price and the nominal exchange rate. We find a strongly significant long-run relation. We then use the nonlinear ARDL (NARDL) model to assess the asymmetries on both the short- and long-run elasticities. The estimation results confirm the presence of a strong asymmetry in the long-run elasticities.

J.E.L. Classification: C53, F32, H62.

Keywords: energy prices, asymmetric cointegration.

[¶] Corresponding author. Paper presented at the 16th INFER Annual Conference, held at the Faculty of Economics, University Gabriele D'Annunzio, Pescara (Italy), May 29th-31st 2014.

LONG- AND SHORT-RUN PRICE ASYMMETRIES IN THE ITALIAN ENERGY MARKET: THE CASE OF GASOLINE AND HEATING GASOIL

1. Introduction

The asymmetry between gasoline and crude oil prices has long been studied in the theoretical and applied literature, starting from Bacon (1991) "rockets and feathers" paper. The recent meta-analysis by Perdiguero-García (2013) indicates that price asymmetry is pervasive and depends mainly on the non-competitive nature of retail markets: the last segment of the market (i.e., the one where final consumers are involved)[1] is more likely to show price asymmetries, relative to the upper segments, which face higher levels of competition in international markets. These results are consistent with those of Frey and Manera (2007), who carry out a meta-analysis on the econometric models of asymmetric price transmission in various markets and find that asymmetry is robust across different settings: only 13% of the estimated models considered in the survey show no evidence of any kind of asymmetry. Moreover, they find that the asymmetry is not significantly affected by the kind of commodity considered (gasoline, agriculture, alimentary), which implies that the asymmetric price mechanism found in the gasoline market is a consequence of a generic asymmetry of price transmission.[2]

Despite being a relatively general phenomenon (see Peltzman, 2000), price asymmetries have received a special attention in the market of crude-derived fuels, both because of the relevance of these products for the public at large, and because of the large swings in input prices.[3] In the last years, the relevance of gasoline pricing is stronger than ever in the Southern countries of the Eurozone, owing to

[1] The last segment regarding the "relationship between the spot price and the retail price paid by consumers, the relationship between the price of crude oil and the price paid by consumers or the wholesale price and the retail price paid by consumers" (Perdiguero-García, 2013, pag. 392).

[2] While in the general meta-regression results, the probability of rejection of symmetric price behaviour is not influenced by the kind of commodity market considered, it is reduced by the dummy variable indicating agriculture only in the subset of error correction models.

[3] Nearly 49% of works included in Frey and Manera's work cover the gasoline market.

the ongoing debate on the possible consequences of an euro breakup. A recurrent argument against a nominal realignment of the exchange rates within the Eurozone is that the expected nominal devaluation would lead to an immediate increase in gasoline prices in Southern countries, with devastating effects on their economies.

The purpose of this study is to examine the relation between the pre-tax retail prices of gasoline and heating gasoil, with crude price and the nominal exchange rate, focusing on the Italian market, in order to assess the existence of any possible price asymmetry. The Italian gasoline market has been the object of previous analyses, while to the extent of our knowledge there are no published studies on the relation between crude price and heating gasoil. However, in some cases the studies on gasoline price do not examine separately the impact of exchange rate variation on the retail price, and in general they rely on "asymmetric ECM" (A-ECM) approaches where asymmetry is allowed only in the adjustment parameters (short-run elasticity and error correction parameter), not in the long-run elasticities. In other words, these models rule out the existence of two different long-run equilibria, one with rising, and another one with falling prices. Honarvar (2009) finds that the fact of considering asymmetric adjustments around a common cointegrating relation is a shortcoming of the previous studies.[4] In order to improve over the previous empirical evidence, in this study we adopt the nonlinear autoregressive distributed lag (NARDL) approach proposed by Shin *et al.* (2013), that allows for asymmetries in both the short- and long-run parameters.[5]

The remain of the paper falls in four sections. Section 2 provides a survey of previous studies on price asymmetries in the Italian market. Section 3 describes the NARDL approach used in this study. Section 4 presents the results. Section 5 concludes.

2. Asymmetries in gasoline pricing: a survey of the Italian evidence

Galeotti *et al.* (2003) study the price of gasoline in France, Germany, Italy, Spain and U.K., using monthly data from 1985 to 2000. Asymmetries are modelled

[4] Another shortcoming of the previous analyses is that the generally fail to report the long-run elasticities, i.e., the parameters of the "common" long-run equilibrium relation.

[5] A previous application of this approach is Atil, A. *et al.* (2014). However, these authors do not consider the exchange rate pass-through.

via an A-ECM (asymmetric error correction model), which allows for different responses to positive and negative variations in the speed of adjustment (long-run, or persistent, asymmetry) and in the lagged explanatory variables (short-run, or transitory, asymmetry). The effect of crude oil (C) and exchange rate between the USD and local currencies (ER) on pre-tax retail gasoline price (R) are estimated either in a two-stage (production and distribution levels) and in a single-stage time setting: in the former case, in production C and ER enter directly into the gasoline spot price (S) and this enters, in distribution, as the only explanatory variable in the equation of R, while in the latter case, R is modelled as a function of C and ER. In general, even with several differences across countries, results confirm the presence of asymmetric price adjustments, with gasoline price adjusting more rapidly to positive relative to negative crude oil price variations and are stronger in the second stage, which is attributed to a more competitive environment in the refining sector respect to the distribution sector. As far as the exchange rate is concerned, its effects are more diversified: while in the first stage asymmetries emerge significantly, evidence is more scant in the single-stage. Results for Italy show that: 1) in the first stage there is short run exchange rate asymmetry;[6] 2) in the second stage there is asymmetric adjustment to equilibrium and short run price asymmetry; 3) adjustment to equilibrium in the first stage is significant only for upward deviations; some kind of asymmetry which involves only negative variations of the exchange rate emerges in the single-stage.[7]

An extension of the work by Galeotti *et al.* (2003) has been done in Grasso and Manera (2007). The authors consider the same countries, with monthly data spanning from 1985 to 2003, and also base the analysis on two-stage and single-stage models. In addition to the A-ECM framework (though with richer dynamics), they also consider threshold ECM (TAR ECM) and ECM with threshold cointegration (M-TAR ECM). Results are heterogeneous, but some conclusions hold in general: 1) when A-ECMs display price asymmetry this is mainly at the distribution level and in this stage adjustment occurs more gradually than at in the first stage, results which suggest the presence of an oligopolistic retail market; 2)

[6] F-statistics for testing the symmetry hypothesis for exchange rate dynamics are not reported.
[7] As stated in the previous note, no F-statistics are reported.

only unfavourable movements in the exchange rate enter significantly in A-ECMs in the production level, while this evidence disappear in the second stage; 3) threshold ECMs increase the evidence for asymmetric pricing behaviour in the oil market with respect to A-ECMs (some sort of asymmetry occurs in the former case in 34% and 16% of cases, respectively); 4) ECMs with threshold cointegration are better at capturing long run asymmetries with respect to A-ECMs.[8] Specifically for Italy A-ECMs show that: 1) results are quite different from those reported in Galeotti *et al*. (2003) as asymmetry emerges only in the second stage and it is relatively less pronounced; 2) no asymmetry emerges in the first stage; 3) asymmetries with significance only on the positive or negative coefficient is present in the speed of adjustment in the second- and single-stage and in exchange rate variations in the single-stage. Estimates of the threshold ECMs for Italy display on the one hand a weak asymmetry from price dynamics and stronger asymmetry from the lagged exchange rate, and, on the other hand, they display asymmetry in the single-stage in the speed of adjustment and from (contemporaneous) exchange rate variations. Long term asymmetric cointegration in the Italian market exist with the MC-TAR model in either the two-stages setting (both in the distribution and retail level) and in the single-stage model; moreover, long run asymmetries are found to be statistically significant in the first stage for errors above and below the threshold, while in the second and single-stage they are significant only when errors are below the threshold.

Romano and Scandurra (2009) study two different stages of price formation, namely crude oil to wholesale prices and wholesale prices to pump prices. They use weekly data (from January 2000 to November 2008) and the framework is and A-ECM augmented by lags of the dependent variable. In the models they consider, exchange rate effects are not present, as the crude price is converted in local currency. This corresponds to imposing the (unverified) constraint that the two elasticities of gasoline price to crude oil and the exchange rate are equal. The outcomes of the estimates show that both in the wholesale and retail market asymmetric effects exist only in the autoregressive behaviour of price, while the speed of adjustment and price dynamics are symmetric. The same authors update

[8] In all countries, estimates of threshold cointegration are found with consistent M-TAR models.

their estimates[9] and extend the augmented A-ECM by including a volatility variable of the considered market proxied by GARCH estimates of the standard deviations of oil quotation and industrial price (Romano and Scandurra, 2012). Moreover, besides estimation in the full sample, they estimate the models in two different subsamples based on structural change test on the volatility variable: the break occurs in 22 September 2008 and separates the period of low volatility (pre-break) from a period of high volatility (post-break). They find no evidence of asymmetry in the wholesale market and volatility is not a significant explanatory variable of price formation, while mixed findings are found in the retail market: with the full sample, short term asymmetry arises in the response of gasoline prices to wholesale prices, both contemporaneous and lagged; in the first subsample, asymmetry is present in the autoregressive part of price formation and lagged wholesale prices; in the second subsample, no asymmetries emerge; in the whole sample and in the two subsamples, volatility is an important determinant of gasoline prices.

Table 1 summarizes the results of the previous studies that adopt an approach similar to ours, i.e., where a single stage of transmission from crude to the final product price is considered. The Table does not display the long-run elasticities, since these are usually not reported in the previous studies. The results confirm the existence of asymmetries, but the evidence on their directions is rather mixed. In general, the speed of adjustment (resumed by the size of the error correction coefficient) is greater in case of positive shocks, but as far as the size of the impact is concerned, the reaction is in some cases greater to positive than to negative shocks, and in other case the reverse occurs. The order of magnitude of exchange rate pass-through is especially controversial, going in case of devaluation from about 10% (0.09) to more than 100% (1.10), the latter estimate showing a possible "overshooting".

3. Data sources and time series properties

The gasoline pre-tax retail prices (R) come from the *Oil bulletin* of the European Commission Energy markets observatory and statistics,

[9] Data span from January 2000 to September 2010.

http://ec.europa.eu/energy/observa-tory/oil/bulletin_en.htm.[10] The database reports separately the pre-tax retail price, the excises, the VAT rate and the pump price. Before 2000 the pump price is not provided but can be easily obtained by applying the VAT rate to the sum of the industrial price and the excises. The prices are expressed in ITL before the euro changeover date (January 2002), and in euro thereafter. All the prices before 2002 where expressed in ECU/EUR by applying the exchange rates provided in the database.

Heating gasoil price come from the *Energy Statistics* reported by the Italian Ministry of Industrial Development (2014) website.[11]

The average crude price in USD per barrel (C) comes from the 2013#1 CD-ROM edition of the International Financial Statistics, series "PETROLEUM:AVERAGE CRUDE PRICE" (00176AAZZF...).

From 1994:1 to 1998:12 we used the ECU/USD exchange rate. From 1999:1 onwards, the EUR/USD exchange rate. Both series come from the 2013#1 CD-ROM edition of the IFS.

The four variables (in logs) were tested for the presence of unit roots, using both the Dickey-Fuller (1981) and the Phillips-Perron (1988) test.[12] The results are summarized in Table 2. All the time series display a unit root.

[10] From 1994 to 2005 we used the data reported in the spreadsheet Italie.xls of the database per country (http://ec.europa.eu/energy/ob-servatory/oil/doc/time_series/time_series_country.zip). From 2006 to 2012 we used the *Oil bulletin price history* database (http://ec.europa.eu/energy/observa-tory/reports/Oil_Bulletin_Prices_History.xls). Since the weekly data are collected each Monday, there are missing observations (each Easter Monday is a bank holiday in Italy, and sometimes Christmas or New Year's Day fall on Monday). These calendar effects have been smoothed out by replacing the missing observation with the average of the preceding and following observation. The regular weekly series thus obtained has been converted to monthly frequency by taking the monthly averages of weekly data. Starting from 2012:1, the series come from the Energy statistics of the Italian Ministry of Economic Development: http://dgerm.sviluppoeconomico.gov.it/dgerm/prezzimedi.asp?prodcod=1&anno=YEAR (where YEAR is the desired year).

[11] From 1994 to 1995 we used the data coming from the EU Market Observatory for Energy (http://ec.europa.eu/energy/observatory/oil/bulletin_en.htm). Weekly data were converted to monthly data by taking simple averages.

[12] The number of lags in the ADF test was selected using the modified Hannan-Quinn information criterion as suggested by Ng and Perron (2001). The lag selection considered a maximum lag length of 14 months. The Phillips-Perron test was performed using the Bartlett kernel with the Newey-West (1994) bandwidth. The test for the variables in levels included a drift, and the tests for the variables in first differences were performed accordingly with no deterministic component.

4. Short- and long-run asymmetries

4.1 The model structure

The long-run relations between the fuel prices and the two explanatory variables (crude price and the EUR/USD exchange rate) were estimated using both the usual cointegration approach, and the asymmetric cointegration approach proposed by Shin *et al*. (2013).

In the standard cointegration approach the dependent variable responds in the same way to both increases and decreases in each explanatory variable. In our case, this approach leads to the following linear ECM model:

$$\Delta y_t = \rho\zeta_{t-1} + \sum_{j=1}^{p-1}\gamma_j \Delta y_{t-j} + \sum_{j=0}^{q-1}\left(\pi_{1j}\Delta c_{t-j} + \pi_{2j}\Delta er_{t-j}\right) + \varepsilon_t \qquad \textbf{(1)}$$

where small caps indicate logarithms, y is the dependent variable (i.e., either r, log of pre-tax gasoline retail price, or g, log of the heating gasoil price), c is the log of the crude price in dollars, er is the log of the EUR/USD exchange rate, ρ is the feedback coefficient (expected to be negative), γ_j and π_{ij} are coefficients (in particular, π_{10} and π_{20} are the impact elasticities of petrol price to crude price and the exchange rate respectively) and ζ_t is the cointegrating residual:

$$\zeta_t = y_t - \beta_1 c_t - \beta_2 er_t \qquad \textbf{(2)}$$

where the β_i are the long-run elasticities.[13]

The asymmetric cointegration approach proposed by Shin *et* al. (2013) uses a nonlinear auto-regressive distributed-lag (NARDL) model, whose structure derives from the ARDL model (Pesaran *et al.,* 2001), and whose nonlinearity derives from the fact that each explanatory variable is decomposed in two partial sum processes, one that cumulates positive changes, and the other one that cumulates negative changes. This approach leads to the following nonlinear error correction model

$$\Delta y_t = \rho\xi_{t-1} + \sum_{j=1}^{p-1}\gamma_j \Delta y_{t-j} + \sum_{j=0}^{q-1}\left(\pi_{1j}^{+}\Delta p_{t-j}^{+} + \pi_{1j}^{-}\Delta p_{t-j}^{-} + \pi_{2j}^{+}\Delta e_{t-j}^{+} + \pi_{2j}^{-}\Delta e_{t-j}^{-}\right) + \varepsilon_t \qquad \textbf{(3)}$$

[13] In order to keep notation as simple as possible, we assumed that in the error-correction representation the order of lags was equal for both explanatory variables. In practical applications this assumption can be easily relaxed.

where a "+" indicates positive changes and a "–" indicates negative changes in the explanatory variables, the short-run coefficients differ for positive and negative changes, and ξ_t is the cointegrating residual obtained from the static asymmetric relation

$$\xi_t = y_t - \beta_1^+ p_t^+ - \beta_1^- p_t^- - \beta_2^+ e_t^+ - \beta_2^- e_t^- \quad \textbf{(4)}$$

where in turn a "+" (respectively, a "–") over the variable indicates the partial sum of its positive (respectively, negative) changes, as follows:

$$x_t^+ = \sum_{j=1}^{t} \Delta x_j^+ = \sum_{j=1}^{t} \max\left(\Delta x_j, 0\right)$$

$$x_t^- = \sum_{j=1}^{t} \Delta x_j^- = \sum_{j=1}^{t} \min\left(\Delta x_j, 0\right)$$

and

$$x_t = x_t^+ + x_t^-$$

While in the standard ECM model the responses to a positive or a negative shock are perfectly symmetric, the NARDL model allows for both different short- and long-run elasticities, or, in other words, for different dynamic multipliers, following a positive or a negative shock to each explanatory variables.

Following Pesaran *et al.* (2001) the estimation of the NARDL, as well as the bounds testing for the existence of a long-run asymmetric relation between the variables, are performed in the following unrestricted ARDL parameterization

$$\begin{aligned} \Delta y_t = {} & \rho y_{t-1} + \theta_1^+ p_{t-1}^+ + \theta_1^- p_{t-1}^- + \theta_2^+ e_{t-1}^+ + \theta_2^- e_{t-1}^- + \sum_{j=1}^{p-1} \gamma_j \Delta y_{t-j} + \\ & + \sum_{j=0}^{q-1} \left(\pi_{1j}^+ \Delta p_{t-j}^+ + \pi_{1j}^- \Delta p_{t-j}^- + \pi_{2j}^+ \Delta e_{t-j}^+ + \pi_{2j}^- \Delta e_{t-j}^- \right) + \varepsilon_t \end{aligned} \quad \textbf{(6)}$$

where instead of the cointegrating residual we included the partial sums of each explanatory variables. The asymmetric long-run coefficients in Eq. (6) and (4) are tied by the following relationships:

$$\beta_i^+ = -\theta_i^+ / \rho;\ \beta_i^- = -\theta_i^- / \rho$$

4.2 Symmetric cointegration estimates

Table 3 and 4 report the estimates of the static cointegrating regression (2), along with the cointegration tests, respectively for the gasoline retail price R and for the heating gasoil price G. We used both Phillips and Hansen (1990) fully modified OLS (FMOLS) estimator, and Johansen (1988) maximum-likelihood estimator, and tested for the existence of a cointegrating relationship with both Engle and Granger (1987) residual-based test, and Johansen (1988) maximum likelihood-based trace test.

In both cases, both approaches reject strongly the hypothesis of non cointegration. The evidence indicates the existence of only one cointegrating relation among the variables, and the estimates of the long-run elasticities obtained with the two different estimation methods are very similar. As far as crude price is concerned, the long-run elasticity of retail gasoline price is equal to 0.64 both in the FMOLS and the ML estimates, and the long-run elasticity of heating gasoil ranges from 0.77 (FMOLS) to 0.78 (ML). As for the exchange rate, the elasticity of gasoline price is equal to 0.66 with FMOLS estimates and 0.67 with ML estimates, and the elasticity of heating gasoil price ranges from 0.62 to 0.65.

As far as the exchange rate pass-through is concerned, the impact of a 10% devaluation of the exchange rate will result in an increase by about 6.4% in the gasoline pre-tax retail, which in turn, given the structure of the Italian excise taxes, will bring about an increase of about 3% of the gasoline pump price (considering that excise taxes account for about a half of the pump price). The pass-through to heating gasoil price is of a similar order of magnitude.

Using the cointegrating residual obtained from the FMOLS regressions, we estimated ECM models like Eq. (1), with a maximum order of lags equal to 2. The estimation results are presented in Table 5 and 6 for . The estimates show a strong presence of error correcting behavior, with a very significant coefficient on the lagged cointegrating residual.

Both lags of the exchange rate and lag 2 of the dependent variable and of crude price are not statistically significant. Removing them leaves the equation properties almost unchanged (with the possible exception of a slight decrease in the lagged dependent variable coefficient). However, the equation's residuals are strongly non-normal, with a Jarque-Bera statistic equal to 72.33 in the restricted model. Owing to the large size of the sample, non-normality should not be a major problem (both because the implied t-statistics of the coefficients are large, and because we can invoke a central limit theorem to interpret them as asymptotic t's). A look at the residual graph (Figure 3) shows that the non-normality might depend on the presence of a single outlier in November 2011. This could be a consequence of the increase in excise taxes decided by government Monti. In fact, a look at the pump prices show that this increase was not immediately translated on them. In other words, the producers preferred to compress their margin in order not to lose market share. For this reason the model estimate of the industrial price growth rate is upward biased.

4.3 Asymmetric cointegration estimates

While the evidence provides strong support to the symmetric cointegration model, in principle this does not rule out the possible existence of asymmetries in both the short- and the long-run parameters, and ignoring these asymmetries may lead to biased short- and long-run elasticities.[14] In order to check for the existence of such asymmetries, we estimated the unrestricted NARDL (6), where, on the basis of the previous specification search, we considered a maximum order of lags equal to 2 (i.e., $q = p = 2$). The estimation results are presented in Table 7 and 8 for the gasoline and heating gasoil price respectively, along with the cointegration

[14] It is worth noting that the RESET test statistic for the heating gasoil equation points out the existence of nonlinearities in the specification.

tests, the diagnostic tests, and the tests for asymmetries on both the short- and long-run elasticities.

In the NARDL framework the existence of a significant long-run relation can be tested following two approaches: the first one tests with a *t*-statistic for the significance of the feedback coefficient ρ in Eq. (3), along the lines set out by Banerjee *et al.* (1998); the second one tests with a *F* statistic for the significance of the variables that enter Eq. (6) in level (in the same way as Pesaran *et al.*, 2001).

Both statistics are reported in Table 7 and 8 and are strongly significant for both equations, thus rejecting the null of non-cointegration.[15]

Coming to the properties of the equations, in both cases the short-run elasticities do not differ significantly from each other (i.e., there is no short-run asymmetry in the model response) and are generally close to those of the symmetric linear specification presented in Tables 5 and 6 respectively. For instance, the short-run elasticity of gasoline price to an increase (devaluation) in the exchange rate is equal to 0.37, the elasticity to a decrease is equal to 0.41, and in the symmetric case we had an impact elasticity equal to 0.39. The short-run elasticities in the heating gasoil equation display some more asymmetry, but in any case the *F* test is unable to reject the null hypothesis of absence of short-run asymmetries for both dependent and both explanatory variables.

The situation is completely reversed when it comes to the long-run, where the NARDL approach gives strongly asymmetric elasticities for both gasoline and heating gasoil, quite different from the ones obtained in the linear specification.

As far as gasoline price is concerned, the linear specification provides an elasticity to crude price equal to 0.64. The NARDL estimation suggests that these estimates are upward biased. The long-run elasticity in response to an increase of crude price is lower, at 0.44, and it differs from the response to a decrease, equal to

[15] Owing to the fact that the variables are decomposed into two partial sum processes, the tests statistics are non-standard and it is unclear what critical value should be used. However, Shin *et al.* (2013) propose to adopt the "bound testing" approach by Pesaran *et al.* (2001), and to use their tabulated critical values. A further problem arises because it is unclear what number of regressors to consider: whether the number of explanatory variable (in our case, 2), or the number of their partial sums (in our case, 4). Shin *et al.* (2013) remark that by considering the smallest number the tests becomes more conservative, which implies that if one happens to reject the null, this should provide a stronger evidence than that suggested by the nominal significance level of the test. In fact, the highest critical value (in absolute value) for models with unrestricted intercept and two regressors at the 1% significance level are 6.36 for the *F*-test and -4.1, which implies that in our case the null of non-cointegration is very strongly rejected.

0.60. The *F* test for asymmetry confirms that the two long-run elasticities differ (with a statistics equal to 5.47). In short, gasoline price reacts more to crude price decreases than it does to crude price increases. The reverse occurs with the exchange rate, where the asymmetric long-run elasticity for positive changes (devaluations) is equal to 0.90, while that for negative changes is equal to 0.23, and it is not significant. The *F* test for asymmetry confirms that the two elasticities are significantly different. In other words, there is evidence that a devaluation as a long-run positive effect on gasoline price, but there is no evidence of a similar long-run negative effect in case of an exchange rate appreciation.

A similar pattern applies to the heating gasoil long-run coefficients. In this case the long-run pass-through of an exchange rate devaluation is almost equal to 100% (0.998), while revaluation do not have a significant long-run effect.

5. Conclusions

Using the recent NARDL model by Shin *et al.* (2013) we investigate the asymmetries in fuel pricing in Italy, considering the pre-tax gasoline and heating gasoil retail prices. There is no previous evidence for heating gasoil, and the evidence for gasoline price is rather mixed, but generally signals the existence of asymmetries, possibly determined by oligopolistic behaviour in the "second stage" (retail market).

The results indicate a similar pattern for both gasoline and heating gasoil. In the short run there is no significant evidence of asymmetric behaviour, while in the long run there are strong asymmetries, negative for the crude oil price (where the long-run effect of a decrease is larger than that of an increase) and positive for the exchange rate (where the long-run effect of an increase - devaluation - is larger than that of a decrease - revaluation).

As far as the impact of crude price is concerned, similar results are found in the recent study by Atil *et al.* (2014), who also use the NARDL approach, as well as in some of the previous studies relative to the Italian market. As Atil *at al.* (2014) point out, this outcome may depend on the fact that crude price decreases occur often during downward economic conditions, in which downward price expectation spirals affecting fuel prices are likely to occur. The fact that previous studies (such as Romano and Scandurra, 2009) report a positive asymmetry could be determined

by their estimates mixing-up the effect of the exchange rate and of the crude price variation. The asymmetry in case of the exchange rate is strongly positive, with no long-run significant effect for exchange rate decreases (revaluation), and a long-run pass-through between 0.9 (gasoline) and 1 (heating gasoil).

Most Italian economists consider that the euro is currently too strong against the dollar and express the view that a 20% devaluation of the Eurozone currency would relieve it from the current crisis (Prodi, 2014). According to our estimates, a 20% devaluation of the euro would bring about in the long run a 18% increase in the pre-tax price of gasoline. Owing to the impact of the excises, this would result in a long-run increase of the pump price equal to about 9%. Since the current pump price of gasoline is around 1.75 euros, this would mean an increase of about 15 cents in the long run. This increase could easily be offset by reducing the excises, that since the beginning of the crisis have increased by about 16 cents. It has been found that using aggregated information (i.e., non-daily data) can lower the probability of finding price asymmetries (Perdiguero-García, 2013). This means that it is likely that our results on asymmetry could be even stronger when using weekly or daily data.

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Figures

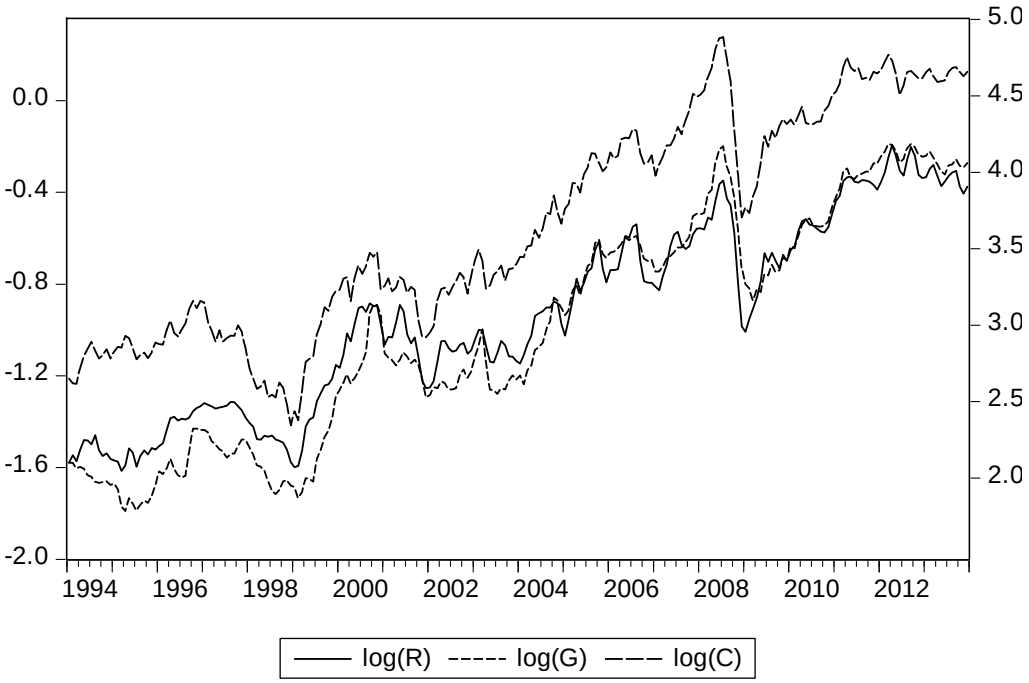


Figure 1 – The logarithms of pre-tax gasoline retail price (R, left-hand scale); heating gasoil price (G; left-hand scale), average crude price (C, right-hand scale).

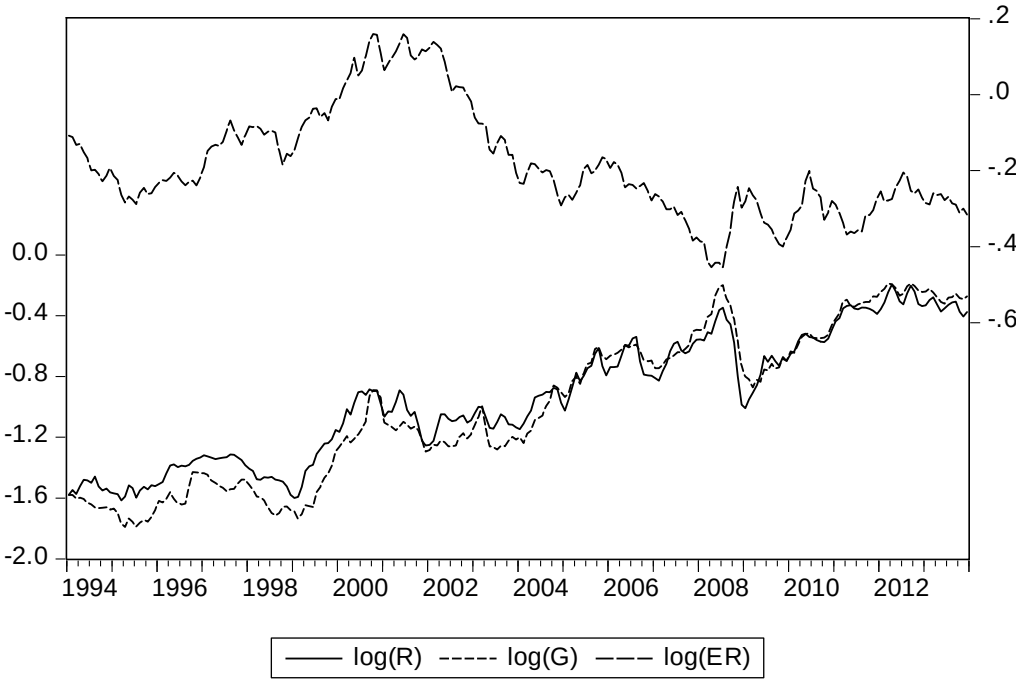


Figure 2 – The logarithms of pre-tax gasoline retail price (LR, left-hand scale); heating gasoil price (G; left-hand scale); nominal exchange rate (LP, right-hand scale).

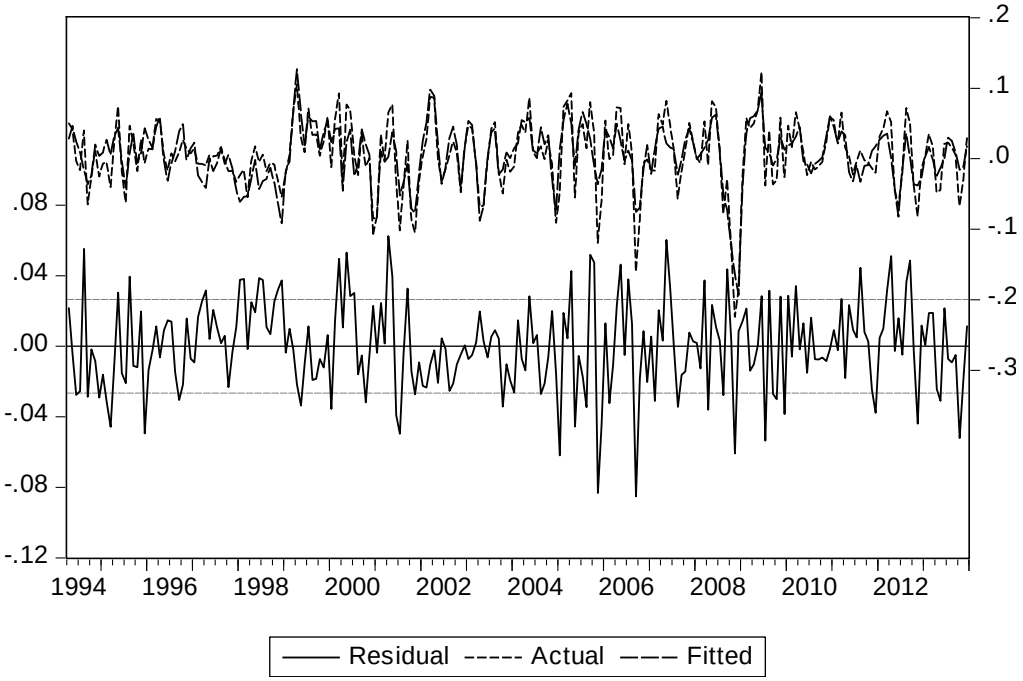


Figure 3.a – Actual, fitted and residual graph from the estimation of Eq. (2) reported in the second column of Table 3.a.

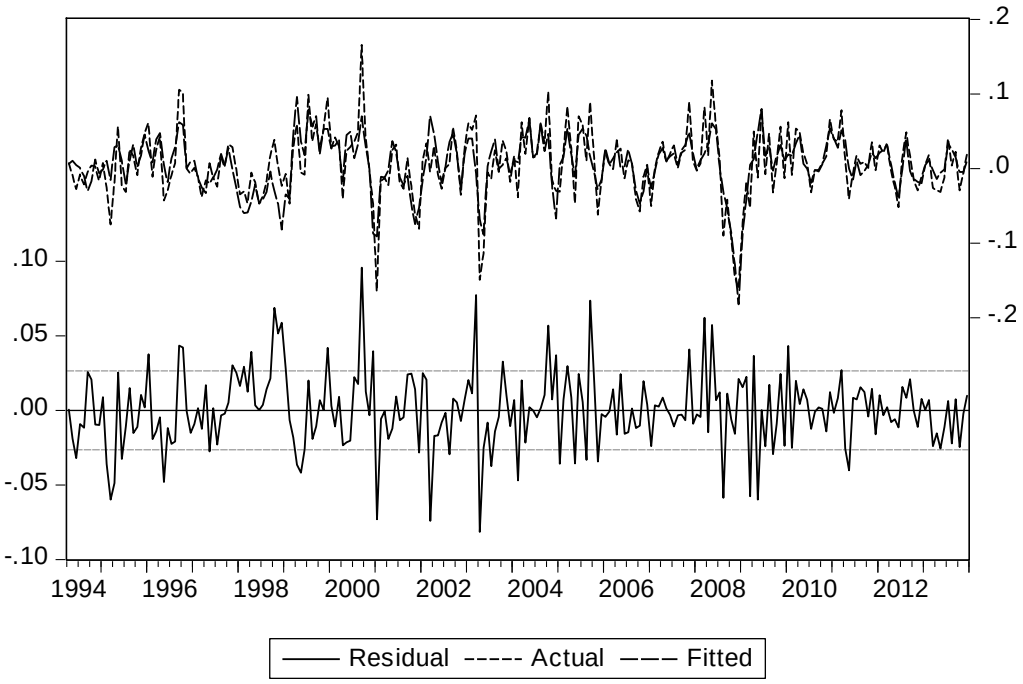


Figure 3.b – Actual, fitted and residual graph from the estimation of Eq. (2) reported in the second column of Table 3.b.

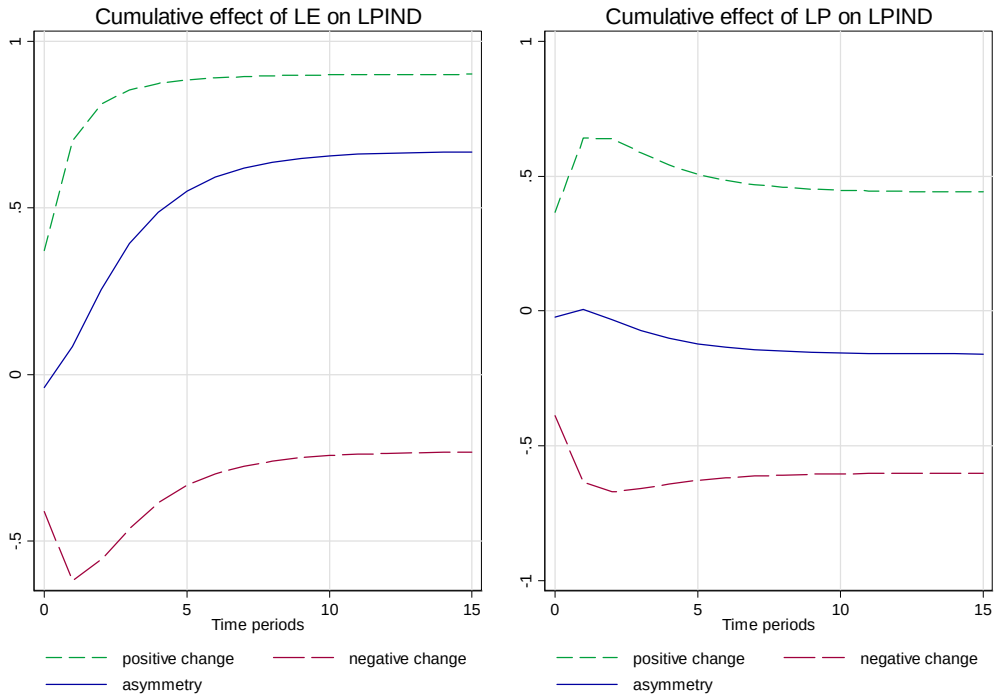


Figure 4.a – The dynamic multiplier for a positive and negative unit change in the explanatory variable, and the multiplier's asymmetry. The left-hand plot shows the response to the exchange rate (with a strong positive asymmetry), the right-hand plot the response to crude oil price (with a moderate negative asymmetry).

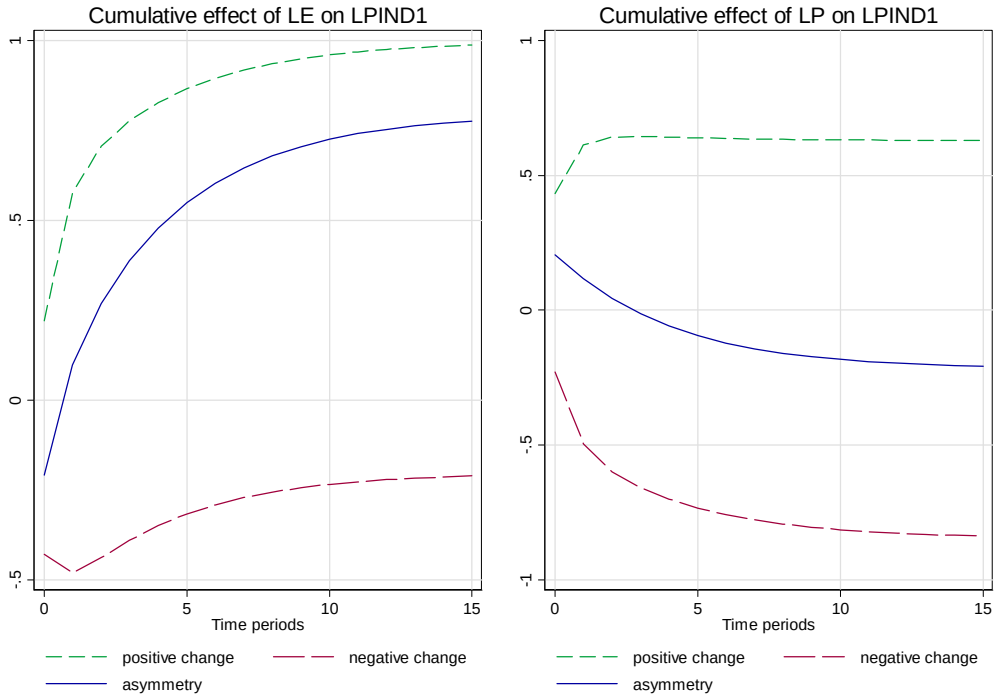


Figure 4.b – The dynamic multiplier for a positive and negative unit change in the explanatory variable, and the multiplier's asymmetry. The left-hand plot shows the response to the exchange rate (with a strong positive asymmetry), the right-hand plot the response to crude oil price (with a moderate negative asymmetry).

Tables

Table 1 – A summary of the previous empirical results

	crude price		exchange rate		error correction	
	positive	negative	positive	negative	positive	negative
Galeotti *et al.* (2003)	0.19	0.24	-0.06	0.46	-1.37	-1.36
Grasso and Manera (2007) [1]	0.26	0.26	0.09	0.68	-0.23	0.01
Grasso and Manera (2007) [2]	0.42	0.26	1.10	0.26	-0.20	0.15
Romano and Scandurra (2009)	0.16	0.12			-0.13	-0.08
Romano and Scandurra (2012)	0.09	0.14			-0.18	-0.12

The Table reports the short-run asymmetric coefficients, i.e., the impact elasticities to crude price and exchange rate, and the error correction coefficient (that measures the speed of adjustment towards the long-run relation). The long-run elasticities are not reported by the studies listed in the Table. Grasso and Manera [1] are A-ECM estimates, while Grasso and Manera [2] are TAR-ECM estimates.

Table 2 – Results of the unit root tests

Variable	transformation	ADF	*p*-val.	lags	PP	*p*-val.	b.width
log(R)	level	-1.27	0.64	2	1.25	0.65	2
	first difference	-9.96	0.00	0	9.21	0.00	9
log(G)	level	-0.85	0.80	1	-0.74	0.83	5
	first difference	-3.60	0.00	10	-10.14	0.00	2
log(C)	level	-1.14	0.70	1	-1.06	0.73	4
	first difference	-7.69	0.00	2	-11.89	0.00	0
log(ER)	level	-1.36	0.60	2	-1.43	0.57	4
	first difference	-11.39	0.00	0	-11.33	0.00	2

ADF is the augmented Dickey-Fuller (1981) test; PP is the Phillips-Perron (1988) test.

Table 3 – Results of the symmetric cointegration estimates, log(R)

	FMOLS		Maximum likelihood	
Var.	coeff.	S.E.	coeff.	S.E.
constant	-3.16	0.04	-3.17	
log(C)	0.64	0.01	0.64	0.02
log(ER)	0.66	0.06	0.67	0.08
R^2	0.98			
adj-R^2	0.98			
EG	-6.06 [0.00]			
Trace(0)			28.13 [0.08]	
Trace(1)			5.14 [0.79]	
Trace(2)			1.03 [0.31]	

EG is the Engle and Granger (1987) test for the null of non cointegration; *Trace(n)* is Johansen's (1988) test for the null hypothesis of at most *n* cointegrating relations among the variables.

Table 4 – Results of the symmetric cointegration estimates, log(G)

	FMOLS		Maximum likelihood	
Var.	coeff.	S.E.	coeff.	S.E.
constant	-3.67	0.05	-3.74	
log(C)	0.77	0.01	0.78	0.02
log(ER)	0.62	0.07	0.65	0.09
R^2	0.98			
adj-R^2	0.97			
EG	-4.88 [0.00]			
Trace(0)			40.3 [0.00]	
Trace(1)			4.41 [0.86]	
Trace(2)			0.70 [0.31]	

EG is the Engle and Granger (1987) test for the null of non cointegration; *Trace(n)* is Johansen's (1988) test for the null hypothesis of at most *n* cointegrating relations among the variables.

Table 5 – Dynamic linear estimation of the gasoline pre-tax retail price equation

Variable	Coeff.	S.E.	Coeff.	S.E.
constant	0.0001	0.0017	0.0001	0.0018
ζ_{-1}	-0.1723	0.0368	-0.1885	0.0352
Δr_{-1}	0.2398	0.0651	0.2494	0.0644
Δr_{-2}	-0.0508	0.0516		
Δc	0.3941	0.0223	0.3918	0.0244
Δc_{-1}	0.1403	0.0372	0.1319	0.0365
Δc_{-2}	-0.0622	0.0367	-0.0828	0.0242
Δer	0.3509	0.0793	0.3859	0.0654
Δer_{-1}	0.1366	0.0851		
Δer_{-2}	-0.0920	0.0825		
R^2	0.721		0.715	
adj-R^2	0.710		0.708	
SC(2)	2.11 [0.12]		0.98 [0.38]	
SC(24)	1.12 [0.32]		0.95 [0.54]	
HET	2.56 [0.08]		2.56 [0.02]	
NOR	5.50 [0.06]		6.47 [0.04]	
FF	0.46 [0.49]		0.18 [0.67]	

SC(n) is the Lagrange multiplier test for the hypothesis of no serial correlation up to order n, HET is the Lagrange multiplier test for the absence of heteroskedasticity (Breusch and Pagan, 1980), NOR is the test for the normality of the residuals, FF is Ramsey's *RESET* test for omitted variables and nonlinearity of the regression specification. All the tests are presented in their F form, with p-values in brackets.

Table 6 – Dynamic linear estimation of the pre-tax heating gasoil price equation

Variable	Coeff.	S.E.	Coeff.	S.E.
constant	0.0002	0.0018	0.000407	0.001739
ζ_{-1}	-0.1692	0.0283	-0.157158	0.027994
Δg_{-1}	0.1781	0.0626	0.162007	0.053660
Δg_{-2}	-0.1001	0.0524		
Δc	0.3273	0.0232	0.321318	0.027979
Δc_{-1}	0.0934	0.0350	0.087024	0.033428
Δc_{-2}	-0.0790	0.0327		
Δer	0.3427	0.0789	0.367598	0.094200
Δer_{-1}	0.1055	0.0851		
Δer_{-2}	-0.1069	0.0818		
R^2	0.660		0.647246	
adj-R^2	0.646		0.639643	
SC(2)	0.56 [0.57]		2.12 [0.12]	
SC(24)	1.30 [0.17]		1.53 [0.06]	
HET	1.96 [0.04]		1.54 [0.17]	
NOR	11.4 [0.00]		23.37 [0.00]	
FF	4.34 [0.04]		5.06 [0.03]	

SC(n) is the Lagrange multiplier test for the hypothesis of no serial correlation up to order n, HET is the Lagrange multiplier test for the absence of heteroskedasticity (Breusch and Pagan, 1980), NOR is the test for the normality of the residuals, FF is Ramsey's *RESET* test for omitted variables and nonlinearity of the regression specification. All the tests are presented in their F form, with p-values in brackets.

Table 7 – Dynamic nonlinear estimation of the pre-tax gasoline retail price equation

Variable	Coeff.	S.E.	*t*
constant	-0.397	0.065	-6.13
r_{-1}	-0.256	0.042	-6.14
c_{-1}^{+}	0.113	0.023	4.87
c_{-1}^{-}	0.154	0.028	5.54
er_{-1}^{+}	0.231	0.049	4.67
er_{-1}^{-}	0.059	0.038	1.55
Δr_{-1}	0.179	0.054	3.33
Δc^{+}	0.366	0.047	7.84
Δc^{-}	0.387	0.042	9.28
Δc_{-1}^{+}	0.191	0.053	3.64
Δc_{-1}^{-}	0.125	0.053	2.37
Δer^{+}	0.373	0.167	2.23
Δer^{-}	0.412	0.149	2.76
Δer_{-1}^{+}	0.179	0.146	1.23
Δer_{-1}^{-}	0.125	0.053	2.37
Short-run asymmetry F-tests			
c_t	0.194	[0.660]	
er_t	0.059	[0.808]	
Long-run coefficients			
c_{-1}^{+}	0.441	[0.000]	
c_{-1}^{-}	0.600	[0.000]	
er_{-1}^{+}	0.901	[0.000]	
er_{-1}^{-}	0.231	[0.100]	
Long-run asymmetry F-tests			
c_t	5.478	[0.020]	
er_t	7.932	[0.005]	
Model diagnostic			
t-BDM	-6.142		
F-PSS	7.655		
R^2	0.728		
adj-R^2	0.711		
SC(40)	36.720	[0.619]	
HET	7.227	[0.007]	
NOR	1.700	[0.428]	
FF	0.498	[0.684]	

SC(*n*) is the Lagrange multiplier test for the hypothesis of no serial correlation up to order *n*, HET is the Lagrange multiplier test for the absence of heteroskedasticity (Breusch and Pagan, 1980), NOR is the test for the normality of the residuals, FF is Ramsey's *RESET* test for omitted variables and nonlinearity of the regression specification. All the tests are presented in their *F* form, with *p*-values in brackets.

Table 8 – Dynamic nonlinear estimation of the pre-tax heating gasoil price equation

Variable	Coeff.	S.E.	*t*
constant	-0.313	0.053	-5.87
g_{-1}	-0.178	0.030	-5.87
c_{-1}^{+}	0.112	0.021	5.23
c_{-1}^{-}	0.151	0.025	6.12
er_{-1}^{+}	0.178	0.041	4.28
er_{-1}^{-}	0.036	0.036	0.99
Δg_{-1}	0.151	0.053	2.86
Δc^{+}	0.433	0.047	9.30
Δc^{-}	0.229	0.042	5.47
Δc_{-1}^{+}	0.079	0.052	1.52
Δc_{-1}^{-}	0.122	0.051	2.40
Δer^{+}	0.221	0.166	1.33
Δer^{-}	0.429	0.149	2.88
Δer_{-1}^{+}	0.186	0.167	1.11
Δer_{-1}^{-}	0.027	0.146	0.19
Short-run asymmetry F-tests			
c_t	2.428	[0.121]	
er_t	0.018	[0.893]	
Long-run coefficients			
c_{-1}^{+}	0.629	[0.000]	
c_{-1}^{-}	0.848	[0.000]	
er_{-1}^{+}	0.998	[0.000]	
er_{-1}^{-}	0.202	[0.314]	
Long-run asymmetry F-tests			
c_t	4.943	[0.027]	
er_t	5.468	[0.020]	
Model diagnostic			
t-BDM	-5.873		
F-PSS	7.930		
R^2	0.671		
adj-R^2	0.651		
SC(40)	73.330	[0.001]	
HET	0.121	[0.728]	
NOR	25.18	[0.000]	
FF	1.221	[0.303]	

SC(*n*) is the Lagrange multiplier test for the hypothesis of no serial correlation up to order *n*, HET is the Lagrange multiplier test for the absence of heteroskedasticity (Breusch and Pagan, 1980), NOR is the test for the normality of the residuals, FF is Ramsey's *RESET* test for omitted variables and nonlinearity of the regression specification. All the tests are presented in their *F* form, with *p*-values in brackets.