*Manuscript
Click here to view linked References

# LONG- AND SHORT-RUN PRICE ASYMMETRIES AND HYSTERESIS IN THE ITALIAN GASOLINE MARKET

**Alberto Bagnai**[a,b,c]

[a] Corresponding author: Department of Economics, University "Gabriele D'Annunzio", viale Pindaro 42, I-65127, Pescara (Italy), bagnai@unich.it, http://bagnai.org, phone: +39 347 7008977, fax +39 06 8081100

[b] a/simmetrie, Italian Association for the Study of Economic Asymmetries, via Filippo Marchetti 19, I-00199, Roma (Italy)

[c] INFER – International Network for Economic Research, http://www.infer-research.net/

**Christian Alexander Mongeau-Ospina**[b]

[b] a/simmetrie, Italian Association for the Study of Economic Asymmetries, via Filippo Marchetti 19, I-00199, Roma (Italy)

**Abstract:** Using monthly data from 1994 to 2013 we study the long-run relation of the pre-tax retail prices of gasoline 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, as well as the presence of hysteresis in the pricing behavior. The estimation results confirm the presence of asymmetry in the long-run elasticities, with significant differences between the crude price and the exchange rate, as well as the presence of hysteresis in the relation between the retail price of gasoline and crude oil price.

**J.E.L. Classification:** C22, D43, E31, L71, Q41.

**Keywords:** energy prices, asymmetric cointegration, asymmetric price adjustment, exchange rate pass-through, hysteresis.

# LONG- AND SHORT-RUN PRICE ASYMMETRIES AND HYSTERESIS IN THE ITALIAN GASOLINE MARKET

## 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. 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.

Price asymmetries have received a special attention in the market of crude-derived fuels, for several reasons: the relevance of these products for the general public, the large swings experienced by crude oil prices in the last decade, and the policy implications of the asymmetry. Asymmetry may indicate that the producers are exploiting their market power, or that the retailers are taking advantage of the consumers' search costs (Balke *et al.*, 2000). This would call for different policy responses, such as antitrust policies, or the obligation for the retailers to display prices.[1] However, Peltzman (2000) shows that "prices rise faster than they fall"

[1] On the Italian motorways, for instance, it is mandatory to display on billboards located at each toll booth the prices of gasoline and diesel oil in each station located in the next stretch, highlighting the retailer that offers the lowest price, in order to facilitate the consumers' search process.

even in competitive markets. In principle, if asymmetry may coexist with competitive market forces, one should weigh the advantages of addressing the (supposed) market imperfections with the costs determined by reducing the economies of scale. Tappata (2009) suggests that asymmetry does not necessarily imply collusive behaviour and can be determined by consumers' imperfect information. Moreover, other empirical work points out that the asymmetry may also depend on more benign causes, such as inventory management (Kaufmann and Laskowski, 2005), and refining adjustment costs (Balke *et al.*, 2000). At the same time, Perdiguero-García (2013) suggests that the asymmetry 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) is more likely to show price asymmetries, relative to the first segments, which face higher levels of competition in international markets.

Another relevant policy issue is related to the correct measurement of the impact on domestic gasoline prices of an increase in crude oil price or of exchange rate devaluation. The latter issue is especially important in the Southern countries of the Eurozone, owing to the ongoing debate on the possible inflationary consequences of a euro breakup. A recurrent argument against a segmentation of the Eurozone is that the expected nominal devaluation in Southern countries would lead to an immediate increase in gasoline prices and, through this channel, in the average inflation rate, with devastating effects on their economies.

Despite the large empirical literature (Grasso and Manera, 2007), a consensus on the causes, the size and the sign of asymmetries in the gasoline market has not been reached. This inconclusiveness could depend on three shortcomings of the previous empirical analyses: usually, they do not consider the possible presence of asymmetries in the long-run coefficients; quite often they do not assess separately the impact on domestic gasoline price of a variation in crude oil price and in the exchange rate; finally, they do not take into account the fact that the response of the domestic price may depend not only on the sign, but also on the size of the shock in the explanatory variables (i.e., they ignore the possible presence of hysteresis). The purpose of this study is to evaluate the empirical relevance of these issues by studying the relation of the pre-tax retail prices of gasoline with the crude price and the nominal exchange rate in the Italian market, using monthly data from 1994:1 to 2013:12.

As for the first point, Honarvar (2009) points out that previous empirical research relies mostly on "asymmetric ECM" (A-ECM; Granger and Lee, 1989) or "threshold ECM" (TAR-ECM; Hansen, 2000) approaches, where asymmetry is allowed only in the adjustment parameters (short-run elasticities and error correction parameter), not in the long-run elasticities (a recent exception is Atil *et al.*, 2014). If the underlying long-run relation has asymmetric parameters, a symmetric specification may lead to biased estimates, thus compromising the reliability of the long- and short-run parameters estimates. In order to address this issue, 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.[2]

As for the second point, some studies skip the estimation of the exchange rate pass-through to gasoline price by converting the price of crude oil in domestic currency. This amounts to imposing the constraint that the elasticities of gasoline price to both crude price and the exchange rate are equal, both in the short and in the long run. As Warmedinger (2004) points out, this restriction appears to be sensible only in the long run. However, the recent literature shows us that at an aggregate level the long-run exchange rate pass-through coefficients differ significantly from those of the other "shifters" variables (the variables used to proxy marginal costs, among which crude oil price; Campa and Goldberg, 2005), and are asymmetric in the long run (Delatte and López-Villavicencio, 2012). As far as the existing studies on gasoline prices are considered, whenever the crude price and the exchange rate are considered separately, their short-run coefficients are both asymmetric and different from each other. Once again, an untested assumption of equality between two long-run coefficients could lead to biased estimates. For this reason, in our NARDL model the crude price and the exchange rate are considered separately.

As for the third point, the size and distribution of the shocks to the exchange rate and to crude oil price have historically been very different, with variation in the crude oil price being larger and negatively skewed. Since Baldwin (1989), it is

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

known that the presence of sunk costs (such as marketing research, establishment of the distribution channels, and so on) may call for the adoption of *ad hoc* pricing policies, where the producers may pass-through to the consumer only the small variations in prices, while adopting a strategic behaviour in front of large positive or negative swings, in order to preserve (or expand) their market shares. This implies that pricing behaviour will be hysteretic, i.e., prices will depend not only on the level of the costs and the exchange rate, but also on the size of the (positive or negative) shocks to those variables. More specifically, there will be an "inaction band", defined by an upper and a lower threshold, within which the retailers will not modify their mark-up, thus translating into the final price all the (small) movements in the explanatory variables. Outside this band, the mark-up will be modified in order to compensate for the variations in the costs variables, thereby leading to smaller (and possibly asymmetric) long-run elasticities. The presence and extent of the hysteresis will depend on whether local currency pricing (LCP) or producer currency pricing (PCP) stabilization will prevail. Antoniades (2012) shows that in the Eurozone producer currency pricing (PCP) prevails, especially for products with low elasticity of substitution (as gasoline typically is: see Baumeister and Peersman, 2013). However, he does not take into account the possible presence of asymmetry, nor the presence of hysteresis (explored, among others, by Belke *et al*., 2013). Previous empirical testing of the PCP vs. LCP behaviour (e.g., Campa and Goldberg, 2005) found the evidence to be inconclusive (with the possible prevalence of PCP in the long run for some categories of goods). Inconclusive results may depend on the fact that by ignoring asymmetry and hysteresis, the estimated elasticities are actually mixing up the values that feature in the three different regimes of the "true" model: the "large positive shock", the "inaction band", and the "large negative shock" regime. In principle, if hysteretic behaviour prevails, we would expect the inaction band elasticities to be larger and not significantly different to one, while outside the inaction band we would expect smaller elasticities (possibly not different from zero), as the retailers compensate for swings in the marginal costs by adjusting the mark-up. In order to cope with this issue, we follow Fedoseeva and Werner (2014) by estimating a NARDL model that takes into account the possible existence of three regimes.

The remainder of the paper falls in five 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, setting out the methodology used to take into account both long-run asymmetry and hysteresis. Section 4 presents the estimation results which are discussed in Section 5. Finally, Section 6 concludes and draws policy implications of the paper's findings.

## 2. The Italian evidence

Before presenting our modelling approach, we briefly review the empirical literature on asymmetric price transmission in the Italian gasoline market, in order to assess whether some consistent stylized facts emerge that could help us in our specification strategy.

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 via an A-ECM, thus allowing 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).[3] The effects 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 setting. The results usually confirm the presence of asymmetric price adjustments, with gasoline price adjusting more rapidly to positive than to negative crude oil price variations. Moreover, the adjustment is found to be stronger in the second stage, which is attributed to a more competitive environment in the refining sector with 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 analysis. Results for Italy show that: 1) in the first stage there is short-run exchange rate asymmetry; 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.

[3] It should be stressed that in the framework of A-ECM estimation long-run asymmetry indicates asymmetric adjustment speed to a *unique* long-run equilibrium.

Grasso and Manera (2007) consider the same countries, using 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 TAR-ECM, and ECM with threshold cointegration (M-TAR ECM; Enders and Granger, 1998). Some conclusions hold in general: 1) when A-ECMs display price asymmetry this is mainly at the distribution level, where the adjustment occurs more gradually than in the first stage, which suggests the presence of an oligopolistic retail market; 2) only unfavourable movements in the exchange rate enter significantly in A-ECMs in the production level, while this evidence disappears in the second stage; 3) threshold ECMs increase the evidence for asymmetric pricing behaviour in the oil market with respect to A-ECMs; 4) ECMs with threshold cointegration are better at capturing long-run asymmetries with respect to A-ECMs.

Romano and Scandurra (2009) also adopt a two-stage setting. They use weekly data (from January 2000 to November 2008) and the framework is an A-ECM augmented by lags of the dependent variable. Exchange rate effects are not present, as the crude price is converted in local currency. The results show that both in the wholesale and retail market asymmetric effects exist only in the autoregressive coefficients, while impact elasticities are symmetric. Romano and Scandurra (2012) update their previous estimates, and extend the augmented A-ECM by including a volatility variable evaluated as the generalised autoregressive conditional heteroskedasticity (GARCH) estimates of the standard deviations of oil quotation and industrial price. Moreover, they estimate the models in two different subsamples based on a 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, where volatility is not significant. Mixed findings are found in the retail market: on the full sample, short-term asymmetry arises in the response of gasoline prices to wholesale prices; 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; volatility is always an important determinant of gasoline prices.

Table 1 summarizes the results of the previous single-stage analyses. 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. The speed of adjustment (resumed by the size of the error correction coefficient) is generally greater in case of positive shocks, but the response is in some cases greater to positive than to negative shocks, and in other cases 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". As it is often found in this research field, the results are mixed. Meyer and von Cramon-Taubadel (2004) point out that a major problem is that the debate still lacks of sound theoretical underpinnings that would allow the researcher to design tests able to discriminate between possible sources of asymmetry. In the following, we explore whether this inconclusiveness may depend on the reason suggested by Honarvar (2009), namely, by the fact that the studies surveyed do not allow for asymmetry in the long-run coefficients, and extend the modelling strategy in order to take into account also the possible presence of hysteretic pricing behaviour.

## 3. Estimation methods

*3.1 Asymmetry*

In the standard cointegration approach the dependent variable responds in the same way to both increases and decreases in each explanatory variable, as in the following error correction model (ECM):

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

where small caps indicate logarithms, $r$ is the log of pre-tax gasoline retail price in euros, $c$ is the log of the crude price in dollars, $er$ is the log of the EUR/USD exchange rate, $\rho$ is the feedback coefficient (expected to be negative), $\gamma_j$ and $\pi_{ij}$ are coefficients (in particular, $\pi_{10}$ and $\pi_{20}$ are the impact elasticities of price to crude price and exchange rate, respectively) and $\zeta_t$ is the cointegrating residual:

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

where $\beta_1$ and $\beta_2$ are the long-run elasticities.[4]

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

$$\Delta r_t = \rho \xi_{t-1} + \sum_{j=1}^{p-1} \gamma_j \Delta r_{t-j} + \sum_{j=0}^{q-1} \left( \pi_{1j}^{+} \Delta c_{t-j}^{+} + \pi_{1j}^{-} \Delta c_{t-j}^{-} + \pi_{2j}^{+} \Delta er_{t-j}^{+} + \pi_{2j}^{-} \Delta er_{t-j}^{-} \right) + \varepsilon_t \qquad \textbf{(3)}$$

where the superscripts "+" and "–" indicate, respectively, positive and negative changes in the explanatory variables, the short-run coefficients differ for positive and negative changes, and $\xi_t$ is the cointegrating residual obtained from the static asymmetric relation

$$\xi_t = r_t - \beta_1^{+} c_t^{+} - \beta_1^{-} c_t^{-} - \beta_2^{+} er_t^{+} - \beta_2^{-} er_t^{-} \qquad \textbf{(4)}$$

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

[4] 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.

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

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

and, by definition, the current value of variable $x_t$ is given by the sum of its initial value and the positive and negative partial sums:

$$x_t = x_0 + x_t^+ + x_t^-$$

While in the standard ECM model the responses to a positive or a negative shock are perfectly symmetric, the NARDL model allows for 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

$$\begin{aligned}\Delta r_t &= \rho r_{t-1} + \theta_1^+ c_{t-1}^+ + \theta_1^- c_{t-1}^- + \theta_2^+ er_{t-1}^+ + \theta_2^- er_{t-1}^- + \sum_{j=1}^{p-1} \gamma_j \Delta r_{t-j} + \\ &+ \sum_{j=0}^{q-1} \left(\pi_{1j}^+ \Delta c_{t-j}^+ + \pi_{1j}^- \Delta c_{t-j}^- + \pi_{2j}^+ \Delta er_{t-j}^+ + \pi_{2j}^- \Delta er_{t-j}^-\right) + \varepsilon_t\end{aligned} \qquad \textbf{(5)}$$

The asymmetric long-run coefficients in Eq. (5) and (4) are tied by the following relationships:

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

*3.2 Hysteresis*

Following Fedoseeva and Werner (2014), we extended the previous model in order to take into account the possible presence of an inaction band, where price setters do not modify the mark-up to offset swings in the explanatory variables. The model structure then becomes:

$$\begin{aligned}\Delta r_t &= \rho r_{t-1} + \theta_1^+ c_{t-1}^+ + \theta_1^0 c_{t-1}^0 + \theta_1^- c_{t-1}^- + \theta_2^+ er_{t-1}^+ + \theta_2^0 er_{t-1}^0 + \theta_2^- er_{t-1}^- + \sum_{j=1}^{p-1} \gamma_j \Delta r_{t-j} + \\ &+ \sum_{j=0}^{q-1} \left( \pi_{1j}^+ \Delta c_{t-j}^+ + \pi_{1j}^0 \Delta c_{t-j}^0 + \pi_{1j}^- \Delta c_{t-j}^- + \pi_{2j}^+ \Delta er_{t-j}^+ + \pi_{2j}^0 \Delta er_{t-j}^0 + \pi_{2j}^- \Delta er_{t-j}^- \right) + \varepsilon_t \qquad \textbf{(6)}\end{aligned}$$

where the partial sum decomposition of the explanatory variables is defined as:

$$x_t^+ = \sum_{j=1}^{t} \Delta x_j^+ = \sum_{j=1}^{t} \Delta x_j I\left(\Delta x_j \geq \Delta x^{upper}\right)$$

$$x_t^0 = \sum_{j=1}^{t} \Delta x_j^0 = \sum_{j=1}^{t} \Delta x_j I\left(\Delta x^{lower} < \Delta x_j < \Delta x^{upper}\right)$$

$$x_t^- = \sum_{j=1}^{t} \Delta x_j^- = \sum_{j=1}^{t} \Delta x_j I\left(\Delta x_j \leq \Delta x^{lower}\right)$$

where $I(.)$ is the indicator function, and $\Delta x^{upper}$ and $\Delta x^{lower}$ are the upper and lower thresholds of the inaction band.

The main issue in the implementation of this model is that there are no clear theoretical nor empirical directions on how these thresholds should be set. Verheyen (2013), in a study on exchange rate nonlinearities in EMU exports, defines the inaction band as the one comprised between the 30% and the 70% quantile. Fedoseeva and Werner (2014) define the inaction band as the one comprising the shocks lower than one standard error in absolute value. In both cases this strategy ensures that a comparable number of observations exist under the three regimes. However, the need to have "enough" observations in each regime contrasts with the need of studying the behaviour of firms facing extreme conditions. In our modelling strategy we adopt a data-based method to determine the size of the inaction band, by choosing the symmetric quantile interval, from $q$% to (100-$q$)%, which minimizes the sum of squared residuals of the model.

## 4. Estimation results

*4.1The data*

The gasoline pre-tax retail prices (*R*) come from the *Oil bulletin* of the European Commission Energy Market Observatory.[5] Prices are expressed in ITL before the euro changeover date (January 2002), and in euro thereafter. All the prices before 2002 were expressed in ECU/EUR by applying the exchange rates provided in the database. 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. [6]

[5] http://ec.europa.eu/energy/observatory/oil/bulletin_en.htm. From 1994 to 2005 we used the data reported in the spreadsheet Italie.xls of the database per country (http://ec.europa.eu/energy/observatory/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/-observatory/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).

[6] Starting from 2012:1 the crude oil price comes from the U.S. Energy Information Administration (http://www.eia.gov/forecasts/steo/query/index.cfm?periodType=MONTHLY-&startYear=1996&endYear=2015&formulas=x112x1x7x4x25x8) and the EUR/USD exchange rate from the Pacific Exchange Rate Service (http://fx.sauder.ubc.ca/data.html).

Figure 1 displays the three time series. The three variables (in logs) were tested for the presence of unit roots, using both the Dickey-Fuller (1981) and the Phillips-Perron (1988) test. All the time series display a unit root.[7]

*4.2Symmetric cointegration estimates*

In order to get a benchmark against which to evaluate the size of any possible asymmetry, we first estimated a symmetric cointegrating regression like Eq. (2). We applied the Phillips and Hansen (1990) fully modified OLS (FMOLS) estimator and obtained the following results (coefficients standard errors in parentheses):

$$\begin{array}{lllll} r_t = & -3.16 & +\ 0.64c_t & +\ 0.66er_t & R^2 = 0.98 \\ & (0.04) & (0.01) & (0.06) & EG = -6.06 \end{array} \qquad \textbf{(7)}$$

where $EG$ is the statistic of the Engle and Granger (1987) cointegration test.[8] The hypothesis of non cointegration is strongly rejected. The long-run elasticities are very similar: 0.64 for crude price, 0.66 for the exchange rate.

*4.3 Asymmetric cointegration estimates*

In order to check for the existence of long-run asymmetries, we estimated the unrestricted NARDL (Eq. 5) with a maximum order of lags equal to 2 (i.e., $q = p = 2$), chosen on the basis of the BIC information criterion.[9] The estimation results are

[7] The results are available upon request.

[8] Estimates using Johansen's (1988) maximum likelihood methods were almost identical. Results are available upon request.

[9] $q$ and $p$ were allowed to vary independently between 2 and 13.

presented in Table 2, along with the cointegration 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 $\rho$ in Eq. (3), along the lines set out by Banerjee *et al.* (1998); the second one tests with an $F$ statistic for the significance of the variables that enter Eq. (5) in levels (in the same way as Pesaran *et al.*, 2001). Both statistics are reported in Table 2, as $t$-BDM and $F$-PSS respectively, and are strongly significant, thus rejecting the null of non-cointegration.[10]

The short-run elasticities do not differ significantly from each other (i.e., there is no short-run asymmetry). The situation is completely reversed when it comes to the long-run, where the NARDL approach gives strongly asymmetric elasticities. 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 0.60. The $F$ test for asymmetry confirms that the two long-run elasticities differ at the 5% significance level. In short, gasoline price reacts more to crude price decreases than

[10] Since variables are decomposed into two partial sum processes, the tests statistics are non-standard and it is unclear what critical values should be used. Shin *et al.* (2013) propose to adopt the "bound testing" approach by Pesaran *et al.* (2001), using their tabulated critical values. However, it is unclear what number of regressors to consider: whether the number of explanatory variables (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 test 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 for the $t$-test, which implies that in our case the null of non-cointegration is very strongly rejected.

it does to crude price increases (negative asymmetry). The reverse (positive asymmetry) occurs with the exchange rate, where the long-run elasticity for positive changes (devaluations) is equal to 0.90 and strongly significant, while that for negative changes is equal to 0.23 and it is significant only at 10%. Also in this case, the *F* test for asymmetry confirms that the two elasticities are significantly different.

*4.4 Hysteresis*

The inaction band was defined in terms of quantiles (as in Verheyen, 2013), looking through grid search for the quantile that minimizes the sum of squared residuals of the estimated model (as in Greenwood-Nimmo *et al.*, 2011). This occurs for $q$=0.165. Table 3 reports the estimates of the model selected, along with the model diagnostic, and the *p*-values of the long-run asymmetry test.[11] Unlike in the NARDL estimates reported in Section 4.3, in this case the asymmetry tests give mixed results for the two explanatory variables. The hypothesis of coefficient equality among the three regimes is rejected for both variables. However, as far as the exchange rate is concerned, the test rejects the hypothesis that the long-run elasticity to positive shocks be equal to the elasticity to shocks within the inaction band, while the corresponding hypothesis for negative shocks is rejected only at the 10% level. In the case of crude oil both hypothesis are rejected. This suggests that the two-threshold specification is appropriate only for the latter variable.

[11] Short-run asymmetry tests always fail to reject the null hypothesis of symmetry.

We therefore repeated the estimation procedure by taking into account the existence of an inaction band only for crude price. In this case, the quantile that minimizes the sum of squared residuals is $q$=0.148. The resulting equation, reported in Table 4, displays strongly significant short- and long-run coefficients and passes all the diagnostic tests (with the only exception of the test for residual homoscedasticity). Figures 2 and 3 report the dynamic multipliers in response to a 1% shock in crude price and the exchange rate, respectively. The impact response (at time zero) is always symmetric. In case of a crude price shock (Figure 2), the lagged response to shocks falling within the inaction band (dotted line) is typically larger than that to extreme shocks (although in case of negative shocks this difference fades away after about five months). The long-run response to large shocks is negatively asymmetric, with a long-run elasticity of 0.46 to positive shocks, and of 0.61 to negative ones. In case of exchange rate (Figure 3), there is a strong positive asymmetry, with an elasticity of 0.81 to positive shocks and of 0.47 to negative ones. All the tests for equality between long-run coefficients reject the null, with the exception of the test of equality between the inaction-band and negative-shock elasticities in case of crude price.

## 5. Discussion

In this section we first discuss the main implications of the results, in the light of the methodology used in the paper, then we check their robustness, by taking into accounts several variables that have been proposed in the literature as possible sources of observed asymmetry.

*5.1 Asymmetric cointegration and the gasoline price*

Several features of the results need some discussion.

First, once the asymmetry in the long-run elasticities is taken into account, any evidence of asymmetry in the impact elasticities disappears, both in the two- and in the three-regimes specification. This suggests that the pervasive evidence of asymmetry in the impact elasticities and feedback coefficients found in the previous A-ECM and TAR-ECM studies, summarized in Table 1, was a spurious result caused by neglecting asymmetry in the long-run elasticities. In fact, asymmetry matters on the long run, i.e., it becomes relevant only in the presence of persistent changes in the explanatory variables.

Second, ignoring long-run asymmetry may lead to severely biased estimates. For instance, the long-run elasticity to exchange rate changes is equal to 0.66 in the linear specification (Eq. 7). The estimates of our preferred equation (Table 4) show that this value underestimates the long-run impact of a depreciation (equal to 0.81) and overestimates the long-run impact of an appreciation (equal to 0.47).

Third, the estimation results point out that, unlike in the case of crude price changes, the pricing behaviour in response to exchange rate changes is asymmetric but not hysteretic (there is no evidence of an “inaction band” where the exchange rate variation are fully passed through). As a matter of fact, the changes in exchange rate have been relatively smaller than those in crude oil price. Moreover, since the inception of the euro the perception of the Italian general public has been to have entered a strong and stable currency. Finally, exchange rate variations do obviously affect all the imported goods (not only energy products). It is likely that these facts have prevented the definition of a “pain threshold” (in the sense of Belke *et al.*, 2013) in the gasoline pricing behaviour with respect to exchange rate changes. On the contrary, the fact that crude price changes were sometimes very large (and obviously market-specific) called for some strategic behaviour in gasoline pricing.

These asymmetries in perception may help to explain another puzzling feature of the results, namely, the fact that the asymmetry is positive for the exchange rate, and negative for crude price. The fact that the euro has been perceived by the general public as a stable currency may have helped retailers to take full advantage of its depreciation, by passing it almost fully through to retail prices, while avoiding to adjust prices in response to an appreciation. On the contrary, crude price, owing to its large swings, has had a major role as a signal of changes in the inflation performance of a country. The negative asymmetry with respect to its changes is therefore consistent both with Atil *et al.* (2014) empirical results, and with Taylor (2000) endogenous mark-up model, where a low-inflation environment is found to reduce firms’ market power. At the same time, this feature may explain why the asymmetry to crude price was found to be positive by previous studies (e.g., Romano and Scandurra, 2009), where the crude price was first converted in national currency. This outcome could result from mixing-up the responses to the crude price and the exchange rate in the same coefficient, in a sample period in which exchange rate variability was comparatively larger than in ours.

### *5.2 Robustness check*

In order to check for the robustness of our results, we controlled for several variables whose role in asymmetric pricing has been explored in the literature, namely crude price volatility, inventory dynamics, the degree of capacity utilization, seasonal variation, and taxation.

The expected impact of higher oil price volatility on asymmetry depends on the underlying theoretical model (Radchenko, 2005). On the one hand, if asymmetry depends on oligopolistic coordination among the retailers, an increase in volatility leads to a reduction in observed asymmetry, since retailers fail to coordinate on a collusive price, because of the increase of uncertainty brought about by higher volatility. On the other hand, in a standard search model an increase in volatility leads to higher asymmetry, because it creates an additional signal-extraction problem for consumers, thereby reducing their search activity and increasing retailers' market power. In our robustness check, volatility was measured by the conditional standard deviation of the crude oil price estimated through a GARCH model, as in Romano and Scandurra (2012).

The role of inventory dynamics and capacity utilization was explored among others by Kaufmann and Laskowski, (2005). Building on Reagan and Weitzman (1982), they argue that a profit maximizing firm will respond to an unanticipated fall in demand by selling from inventories, with limited effects in prices, while unanticipated increases in demand will be dampened by higher prices. In order to control for these effects, we augmented the benchmark model with the log of gasoline inventories, and with a proxy of the degree of capacity utilization, given by the deviation of the industrial production index of the energy sector from its long-run component, the latter evaluated through a HP filter.[12]

Following Kaufmann and Laskowski (2005), we examined also the role of seasonal idiosyncratic factors by including in the model a set of seasonal dummy variables. Finally, following Polemis and Fotis (2014), we assessed the impact on asymmetry of the tax regime by taking the after-tax retail price (including both excise taxes and VAT) as the dependent variable in our preferred specification.[13]

The alternative models are labelled in Table 5 as Volatility, Inventories, Capacity, Seasonality, and Taxes, respectively. The Table reports the estimates of the long-run elasticities to crude price and exchange rate, along with the *p*-values of the long-run asymmetry tests under these alternative specifications.[14]

The size of the estimated elasticities is almost unaffected under the alternative specifications, with the only exception of the "Taxes" model (last column of Table 5), whose long-run elasticities are smaller. This is to be expected, since in Italy excise taxes account on average for 50% of the retail price, which implies that

[12] Inventories come from the *Oil Bulletin* ("Bollettino Petrolifero") published by the Italian Ministry of Economic Development: http://dgerm.sviluppoeconomico.gov.it/dgerm/-bollettino.asp (last accessed on 2014-11-06). The industrial production index of the energy sector was obtained from Istat: http://dati.istat.it/Index.aspx?DataSetCode=-DCSC_INDXPRODIND_1 (last accessed on 2014-11-06).

[13] The source of the after-tax retail price is the same used for the pre-tax price, described in Section 4.1.

[14] Complete estimation results are available upon request.

the response of the after-tax price to a shock in the exogenous variables is smaller than that of the pre-tax price.

As in the previous models, the short-run asymmetry tests (not reported) fail to reject the null of symmetry. As far as the long-run asymmetry tests are concerned, those with respect to crude prices variations are robust under any alternative specification. This confirms the evidence of negative asymmetry with respect to crude price changes. The evidence on exchange rate asymmetry is somewhat mixed. In the regressions augmented with volatility, stocks, or capacity utilization (first three columns of Table 5), the size of the elasticities is remarkably stable, but the tests fail to reject the null hypothesis of symmetry at the 5% level; however, symmetry is always rejected at least at the 10% level. The first result is consistent with previous findings by Romano and Scandurra (2012), where the degree of asymmetry tends to decrease in the presence of large volatility, and supports the hypothesis of oligopolistic coordination as a possible source of asymmetry, where an increase in volatility reduces the firms' market power by impeding their coordination. As for stocks and capacity utilization, our results are broadly in line with those of Kaufmann and Laskowski (2005). However, while in their A-ECM specification any evidence of asymmetry disappears once stocks and capacity utilization are included in the model, in our NARDL specification the long-run asymmetry hypothesis is still rejected at least at the 10% significance level.

## 6. Conclusions and policy implications

Using the recent NARDL model by Shin *et al.* (2013), as modified by Greenwood-Nimmo *et al.* (2011), we investigate the presence of asymmetry and hysteresis in pre-tax gasoline retail price in Italy. The results indicate that 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 and positive for the exchange rate. As for hysteresis, the results show that there is no evidence of an "inaction band" with respect to exchange rate changes, while the evidence with respect to the crude price changes is mixed. In the short run, relatively small changes in crude price are passed through almost fully to retail prices. In the long run, only the response to positive shocks seems to be size-dependent.

As pointed out in the discussion of the results, the negative asymmetry with respect to the crude oil price is consistent with Taylor's (2000) endogenous mark-up model, where firms lose market power in a low-inflation environment (see e.g., Campa and Goldberg, 2005, or Atil et al., 2014). Moreover, the negative impact of volatility on the degree of observed asymmetry, as signalled by the reduced significance of the long-run asymmetry test statistics, once volatility is taken into account, gives some further indirect support to the hypothesis that the observed asymmetry may depend on oligopolistic coordination, as hypothesized by Radchenko (2005). This evidence suggests that the existence and the sign of the observed asymmetries support Perdiguero-Garcia's (2013) hypothesis of non-competitive behaviour in gasoline retail markets. As a matter of fact, since for the Italian consumers falling crude prices are an important signal of a low inflation environment, when crude prices fall the retailers may find it more difficult to exercise their market power. The same does not apply to exchange rate movement, both because their amplitude has been comparatively smaller, and because since 1997 (when the Italian lira was pegged to the ECU at a parity close to the irrevocable ITL/EUR exchange rate) the general perception of the Italian consumers is to have adopted a stable currency. As a consequence, we can hypothesize that the up- and downward swings of the EUR/USD exchange rate had little impact on the consumers' inflation expectations, and on the pricing strategies of the retailers. In particular, since these swings were less clearly perceived, it was easier for retailers to incorporate them in final prices, by reacting to devaluations (exchange rate increases) with a positive asymmetry.

A first policy implication therefore is that increased competition in the retail market could be effective in reducing gasoline price dynamics. This intuition is confirmed by another source of empirical evidence. The Italian antitrust authority has ascertained that in the no-logo retailers network (the so-called "pompe bianche"), established in Italy with the legislative decree No. 32/1998, and further fostered by the law No. 111/2011, the consumers are able to save up to 13 cents per litre with respect to the price practised by branded service stations (Italian Competition Authority, 2012).

Another politically relevant issue is the impact on the gasoline retail price of an exchange rate depreciation, that is often invoked as a possible solution for the Eurozone crisis, and that would in any case occur in the Southern countries of the

Eurozone in case of euro breakup. As a cautionary note it is worth stressing that the two hypotheses (a fall of the euro, or a breakup of the euro) are likely to imply different responses, especially because a euro breakup is likely to bring about a dramatic increase in the consumers' inflation expectations. However, no matter how large the uncertainty about the second scenario can be, we believe that it is better to evaluate it on the basis of possibly unbiased estimates. The presence of positive asymmetry with respect to the exchange rate changes induces a downward bias in the symmetric estimate of the corresponding long-run elasticity. By adopting the NARDL specification, the long-run elasticity to a currency depreciation goes to 0.81, from the 0.66 value found in the symmetric estimation. As a consequence, the long-run impact of a 20% depreciation with respect to the USD on the pre-tax retail price would be an increase by about 16%. This is a sizeable change, and the shape of the multipliers reported in Figure 3 shows that it would occur fairly quickly, with a median lag of one month. However, owing to the structure of Italian taxation, where the excises, currently at 72 cents per litre, account for about 40% of the retail price, the long-run impact on the pump price would be of at most 7% (say, about 12 cents per litre). This order of magnitude is consistent with the historical experience (e.g., with the response of gasoline prices to the large devaluation experienced by the euro since its inception in 1999), and implies that the inflationary impact of a currency devaluation could be kept under control by a small reduction in the excises, that since the beginning of the Eurozone crisis have been raised by 16 cents.

More conclusive evidence on the nature and size of gasoline pricing asymmetries would require the application of the NARDL methodology in a two-stage setting, where the production and distribution stage are analysed separately. We leave this task for future research.

## Acknowledgments

We thank Svetlana Fedoseeva and the other participants to the 16th INFER annual conference (Pescara, May 29-31 2014) for their useful remarks, as well as Patrick Lynch for revising our English. The usual disclaimer applies. Financial support from the Italian Ministry of the Education, University and Research (60% funds), as well as from the Nando Peretti Foundation. The funding sources had no role in study design, in the collection, analysis and interpretation of data, in the writing of the report, nor in the decision to submit the article for publication.

## References

Antoniades, A., 2012. Local versus producer currency pricing: evidence from disaggregated data. International Economic Review, 53, 1229-1241.

Atil, A., Lahiani, A., Nguyen, D.K., 2014. Asymmetric and nonlinear pass-through of crude oil prices to gasoline and natural gas prices. Energy Policy, 65, 567-573.

Bacon, R.W., 1991. Rockets and feathers: the asymmetric speed of adjustment of UK retail gasoline prices to cost changes. Energy Economics, 13, 211–218.

Baldwin, R., 1989. Sunk-cost hysteresis. NBER Working Papers, No. 291, Washington, National Bureau of Economic Research.

Balke, N.S., Brown, S.P.A., Yücel, M.K., 2000. Crude oil and gasoline prices: an asymmetric relationship? Economic and Financial Policy Review, first quarter, 2-11.

Banerjee, A., Dolado, J., Mestre, R., 1998. Error-correction mechanism tests for cointegration in a single-equation framework. Journal of Time Series Analysis, 19, 267-283.

Baumeister, C., Peersman, G., 2013. The role of time-varying price elasticities in accounting for volatility changes in the crude oil market. Journal of Applied Econometrics, 28, 1087-1109.

Belke, A., Goecke, M., Guenther, M., 2013. Exchange rate bands of inaction and play hysteresis in German exports – sectoral evidence for some OECD destinations. Metroeconomica, 64, 152–79.

Breusch, T.S., Pagan, A.R., 1980. The Lagrange Multiplier Test and Its Applications to Model Specification in Econometrics. Review of Economic Studies, 47, 239-53.

Campa, J.M., Goldberg, L.S., 2005. Exchange Rate Pass-Through into Import Prices. Review of Economics and Statistics, 87, 679-690.

Delatte, A.-L., López-Villavicencio, A., 2012. Asymmetric exchange rate pass-through: Evidence from major countries. Journal of Macroeconomics, 34, 833-844.

Dickey, D., Fuller, W., 1981. Likelihood ratio statistics for autoregressive time series with a unit root. Econometrica, 49, 1057-1072.

Enders, W., Granger, C. W. J., 1998. Unit-Root Tests and Asymmetric Adjustment with an Example Using the Term Structure of Interest Rates," Journal of Business & Economic Statistics, 16, 304-11.

Engle, R.F., Granger, C.W.J., 1987. Co-integration and error correction: representation, estimation and testing. Econometrica, 55, 251-276.

Fedoseeva, S., Werner, L.M., 2014. Asymmetry and hysteresis: two perspectives on pricing-to-market nonlinearity. Paper presented at the 16th annual conference of the International Network for Economic Research, Department of Economics, Gabriele D'Annunzio University, Pescara (Italy), May 29th-31st 2014.

Frey, G., Manera, M., 2007. Econometric models of asymmetric price transmission. Journal of Economic Surveys, 21, 349–415.

Galeotti, M., Lanza, A., Manera, M., 2003. Rockets and feathers revisited: an international comparison on European gasoline markets. Energy Economics, 25, 175–190.

Granger, C.W.J, Lee, T.H., 1989. Investigation of Production, Sales and Inventory Relationships Using Multicointegration and Non-symmetric Error Correction Models," Journal of Applied Econometrics, 4(S), S145-59.

Grasso, M., Manera, M., 2007. Asymmetric error correction models for the oil–gasoline price relationship. Energy Policy, 35, 156–177.

Greenwood-Nimmo, M., Shin, Y., van Treeck, T., 2011. The Asymmetric ARDL Model with Multiple Unknown Threshold Decompositions: An Application to the Phillips Curve in Canada. mimeo, February, http://www.greenwoodeconomics.com/%28MG2%29IMK_multi.pdf, last accessed on 2014-08-01.

Hansen, B.E., 2000. Sample Splitting and Threshold Estimation. Econometrica, 68, 575-604.

Honarvar, A., 2009. Theoretical explanations for asymmetric relationships between gasoline and crude oil prices with focus on the US market. OPEC Energy Review, September, 205-224.

Italian Competition Authority, 2012. Cognitive survey on the Italian energy market, http://www.agcm.it/component/joomdoc/doc_download/3448-ic44-testo-indagine-28-dic-2012.html, last accessed on 2014-07-24.

Italian Ministry of Industrial Development, 2014. Energy Statistics, http://dgerm.sviluppoeconomico.gov.it/dgerm/prezzimedi.asp?prodcod=3&anno=2014, last accessed on 2014-07-15.

Johansen, S., 1988. Statistical analysis of cointegration vectors. Journal of Economic Dynamics and Control, 12, 231-254.

Kaufmann, R.K., Laskowski, C., 2005. Causes for an asymmetric relation between the price of crude oil and refined petroleum product. Energy Policy, 33, 1587-1596.

Meyer, J., von Cramon-Taubadel, S., 2004. Asymmetric price transmission: a survey. Journal of Agricultural Economics, 55, 581-611.

Newey, W., West, K., 1994. Automatic Lag Selection in Covariance Matrix Estimation. Review of Economic Studies, 61, 631-653.

Ng, S., Perron, P., 2001. Lag Length Selection and the Construction of Unit Root Tests with Good Size and Power. Econometrica, 69, 1519-1554.

Peltzman, S., 2000. Prices rise faster than they fall. Journal of Political Economy, 108, 466-502.

Perdiguero-García, J., 2013. Symmetric or asymmetric oil prices? A meta-analysis approach. Energy Policy, 57, 389–397.

Pesaran, M.H., Shin, Y., Smith, R.J., 2001. Bounds testing approaches to the analysis of level relationships. Journal of Applied Econometrics, 16, 289-326.

Phillips, P.C.B., Hansen, B., 1990. Statistical inference in instrumental variables regression with I(1) processes. Review of Economic Studies, 57, 99-125.

Phillips, P.C.B., Perron, P., 1988. Testing for a Unit Root in Time Series Regression. Biometrika, 75, 335–346.

Polemis, M.L., Fotis, P.N., 2014. The taxation effect on gasoline price asymmetry nexus: Evidence from both sides of the Atlantic. Energy Policy, 73, 225-233.

Radchenko, S., 2005. Oil price volatility and the asymmetric response of gasoline prices to oil price increases and decreases. Energy Economics, 27, 708-730.

Reagan, P.B., Weitzman, M.L., 1982. Asymmetries in price and quantity adjustments by the competitive firm. Journal of Economic Theory, 27, 410-420.

Romano, A.A., Scandurra, G., 2009. Price asymmetries in the Italian retail and wholesale gasoline markets. Proceedings of the SIS meeting on Statistical Methods for the Analysis of Large data-sets, University G. D'Annunzio – Chieti Pescara, 23-25 September.

Romano, A.A., Scandurra, G., 2012. Price asymmetries and volatility in the Italian gasoline market. OPEC Energy Review, 36, 215–229.

Shin, Y., Yu, B., Greenwood-Nimmo, M., 2013. Modelling asymmetric cointegration and dynamic multipliers on a nonlinear ARDL framework. In: Horrace, W.C., Sickles, R.C. (Eds.), Festschrift in Honor of Peter Schmidt, forthcoming, available at http://dx.doi.org/10.2139/ssrn.1807745, last accessed 2014-07-15.

Tappata, M., 2009. Rockets and feathers: Understanding asymmetric pricing. RAND Journal of Economics, 40, 673-687.

Taylor, J., 2000. Low inflation, pass-through, and the pricing power of firms. European Economic Review, 44, 1389-1408, June.

Verheyen, Florian, 2013. Exchange rate nonlinearities in EMU exports to the US. Economic Modelling, 32(C), 66-76.

Warmedinger, T., 2004. Import prices and pricing-to-market effects in the Euro area. European Central Bank Working Paper, No. 299, January.

# Tables

Table 1 – A summary of the previous empirical results[a]

| | 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 | 0.06 |
| Romano and Scandurra (2009) | 0.16 | 0.12 | | | -0.13 | -0.08 |
| Romano and Scandurra (2012) | 0.09 | 0.14 | | | -0.18 | -0.12 |

[a] 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 – Dynamic nonlinear estimation of the pre-tax gasoline retail price equation[a]

| *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 |
| $\Delta r_{-1}$ | 0.179 | 0.054 | 3.33 |
| $\Delta c^{+}$ | 0.366 | 0.047 | 7.84 |
| $\Delta c^{-}$ | 0.387 | 0.042 | 9.28 |
| $\Delta c_{-1}^{+}$ | 0.191 | 0.053 | 3.64 |
| $\Delta c_{-1}^{-}$ | 0.125 | 0.053 | 2.37 |
| $\Delta er^{+}$ | 0.373 | 0.167 | 2.23 |
| $\Delta er^{-}$ | 0.412 | 0.149 | 2.76 |
| $\Delta er_{-1}^{+}$ | 0.179 | 0.146 | 1.23 |
| $\Delta er_{-1}^{-}$ | 0.125 | 0.053 | 2.37 |
| *Short-run asymmetry F-tests* | | | |
| *c* | 0.194 | | |
| *er* | 0.059 | | |
| *Long-run coefficients* | | | |
| $c^{+}$ | 0.441 *** | | |
| $c^{-}$ | 0.600 *** | | |
| $er^{+}$ | 0.901 *** | | |
| $er^{-}$ | 0.231 * | | |
| *Long-run asymmetry F-tests* | | | |
| *c* | 5.478 ** | | |
| *er* | 7.932 *** | | |
| *Model diagnostic* | | | |
| *t-BDM* | -6.14 | | |
| *F-PSS* | 7.66 | | |
| $R^2$ | 0.73 | | |
| adj-$R^2$ | 0.71 | | |
| SC(12) | 1.27 | | |
| SC(24) | 0.83 | | |
| HET | 1.92 ** | | |
| NOR | 1.70 | | |
| FF | 0.49 | | |

[a] SC($n$) is the Lagrange multiplier test for the hypothesis of no serial correlation up to order $n$. This and the following tests are presented in their $F$ form, asterisks stand for significance at 10% (*), 5% (**) or 1% (***).

[b] Lagrange multiplier test for the absence of heteroskedasticity (Breusch and Pagan, 1980).

[c] Test for the normality of the residuals.

[d] *RESET* test for omitted variables and nonlinearity of the regression specification.

Table 3 – Dynamic nonlinear estimation of the pre-tax gasoline retail price equation with two thresholds

| *Variable* | *Coeff.* | *S.E.*[a] | *t* | *Long-run coefficients* | |
|---|---|---|---|---|---|
| constant | -0.500 | 0.083 | -6.03 | | |
| $r_{-1}$ | -0.321 | 0.053 | -6.10 | | |
| $er_{-1}^{+}$ | 0.236 | 0.050 | 4.69 | 0.734 | [0.000] |
| $er_{-1}^{-}$ | 0.252 | 0.076 | 3.30 | 0.783 | [0.000] |
| $er_{-1}^{0}$ | 0.133 | 0.037 | 3.64 | 0.414 | [0.000] |
| $c_{-1}^{+}$ | 0.177 | 0.033 | 5.31 | 0.549 | [0.000] |
| $c_{-1}^{-}$ | 0.192 | 0.033 | 5.78 | 0.597 | [0.000] |
| $c_{-1}^{0}$ | 0.250 | 0.053 | 4.76 | 0.778 | [0.000] |
| $\Delta r_{-1}$ | 0.190 | 0.049 | 3.85 | | |
| $\Delta er^{+}$ | 0.287 | 0.133 | 2.15 | | |
| $\Delta er^{-}$ | 0.435 | 0.129 | 3.36 | | |
| $\Delta er^{0}$ | 0.460 | 0.179 | 2.58 | | |
| $\Delta er_{-1}^{+}$ | 0.101 | 0.155 | 0.65 | | |
| $\Delta er_{-1}^{-}$ | 0.085 | 0.122 | 0.70 | | |
| $\Delta er_{-1}^{0}$ | 0.043 | 0.169 | 0.26 | | |
| $\Delta c^{+}$ | 0.408 | 0.039 | 10.36 | | |
| $\Delta c^{-}$ | 0.383 | 0.039 | 9.80 | | |
| $\Delta c^{0}$ | 0.393 | 0.059 | 6.62 | | |
| $\Delta c_{-1}^{+}$ | 0.145 | 0.047 | 3.06 | | |
| $\Delta c_{-1}^{-}$ | 0.079 | 0.052 | 1.52 | | |
| $\Delta c_{-1}^{+}$ | 0.193 | 0.062 | 3.12 | | |
| *Model diagnostic* | | | | | |
| *t-BDM* | -6.100 | | | | |
| *F-PSS* | 6.778 | | | | |
| $R^2$ | 0.748 | | | | |
| *adj-*$R^2$ | 0.725 | | | | |
| *SC(12)*[b] | 1.190 | [0.292] | | | |
| *SC(24)*[b] | 0.860 | [0.656] | | | |
| *HET*[c] | 1.490 | [0.040] | | | |
| *NOR*[d] | 0.193 | [0.909] | | | |
| *FF*[e] | 3.216 | [0.042] | | | |

| *Long-run asymmetry tests* | *P=N=Z*[f] | *P=N*[g] | *P=Z*[h] | *N=Z*[i] |
|---|---|---|---|---|
| *er* | [0.016] | [0.841] | [0.028] | [0.063] |
| *c* | [0.002] | [0.313] | [0.001] | [0.033] |

[a] S.E. are heteroskedasticity consistent standard errors.
[b] SC(n) is the LM test for the hypothesis of no serial correlation up to order *n*. We report *p*-values in brackets for this and the other tests.
[c] LM test for the absence of heteroskedasticity (Breusch and Pagan, 1980).
[d] Test for the normality of the residuals.

[e] Ramsey's *RESET* test for omitted variables and nonlinearity of the regression.
[f] Test of coefficient equality in the three regimes.
[g] Test of coefficient equality in the positive and negative regime.
[h] Test of coefficient equality in the positive regime and the inaction band.
[i] Test of coefficient equality in the negative regime and the inaction band.

Table 4 – Dynamic nonlinear estimation of the pre-tax gasoline retail price equation with one threshold for the exchange rate and two for the crude price

| *Variable* | *Coeff.* | *S.E.*[a] | *t* | *Long-run coefficients* | |
|---|---|---|---|---|---|
| constant | -0.468 | 0.070 | -6.71 | | |
| $r_{-1}$ | -0.297 | 0.044 | -6.72 | | |
| $er_{-1}^{+}$ | 0.241 | 0.041 | 5.82 | 0.812 | [0.000] |
| $er_{-1}^{-}$ | 0.140 | 0.049 | 2.85 | 0.469 | [0.000] |
| $c_{-1}^{+}$ | 0.135 | 0.025 | 5.47 | 0.455 | [0.000] |
| $c_{-1}^{-}$ | 0.181 | 0.028 | 6.37 | 0.610 | [0.000] |
| $c_{-1}^{0}$ | 0.186 | 0.038 | 4.87 | 0.625 | [0.000] |
| $\Delta r_{-1}$ | 0.195 | 0.052 | 3.75 | | |
| $\Delta er^{+}$ | 0.327 | 0.153 | 2.13 | | |
| $\Delta er^{-}$ | 0.449 | 0.147 | 3.05 | | |
| $\Delta er_{-1}^{+}$ | 0.082 | 0.176 | 0.47 | | |
| $\Delta er_{-1}^{-}$ | 0.144 | 0.130 | 1.10 | | |
| $\Delta c^{+}$ | 0.407 | 0.040 | 10.12 | | |
| $\Delta c^{-}$ | 0.383 | 0.043 | 8.89 | | |
| $\Delta c^{0}$ | 0.367 | 0.052 | 7.07 | | |
| $\Delta c_{-1}^{+}$ | 0.171 | 0.046 | 3.74 | | |
| $\Delta c_{-1}^{-}$ | 0.085 | 0.050 | 1.70 | | |
| $\Delta c_{-1}^{+}$ | 0.210 | 0.053 | 3.95 | | |
| *Model diagnostic* | | | | | |
| *t-BDM* | -6.720 | | | | |
| *F-PSS* | 7.538 | | | | |
| $R^2$ | 0.743 | | | | |
| *adj-*$R^2$ | 0.723 | | | | |
| *SC(12)*[b] | 0.955 | [0.493] | | | |
| *SC(24)*[b] | 0.651 | [0.893] | | | |
| *HET*[c] | 1.588 | [0.027] | | | |
| *NOR*[d] | 0.453 | [0.797] | | | |
| *FF*[e] | 2.586 | [0.078] | | | |

| *Long-run asymmetry tests* | *P=N=Z*[f] | *P=N*[g] | *P=Z*[h] | *N=Z*[i] |
|---|---|---|---|---|
| *er* | | [0.041] | | |
| *c* | [0.000] | [0.005] | [0.003] | [0.845] |

[a] S.E. are heteroskedasticity consistent standard errors.
[b] SC(n) is the LM test for the hypothesis of no serial correlation up to order *n*. We report *p*-values in brackets for this and the other tests.
[c] LM test for the absence of heteroskedasticity (Breusch and Pagan, 1980).
[d] Test for the normality of the residuals.
[e] Ramsey's *RESET* test for omitted variables and nonlinearity of the regression.
[f] Test of coefficient equality in the three regimes.
[g] Test of coefficient equality in the positive and negative regime.
[h] Test of coefficient equality in the positive regime and the inaction band.
[i] Test of coefficient equality in the negative regime and the inaction band.

Table 5 – Robustness check on the preferred specification

| *Long-run coefficients*[a] | | | | | |
|---|---|---|---|---|---|
| | Volatility | Inventories | Capacity | Seasonality | Taxes |
| $er_{-1}^{+}$ | 0.80*** | 0.80*** | 0.80*** | 0.81*** | 0.62*** |
| $er_{-1}^{-}$ | 0.48*** | 0.51*** | 0.48*** | 0.45*** | 0.02 |
| $c_{-1}^{+}$ | 0.46*** | 0.48*** | 0.45*** | 0.46*** | 0.03 |
| $c_{-1}^{-}$ | 0.60*** | 0.61*** | 0.59*** | 0.62*** | 0.33*** |
| $c_{-1}^{0}$ | 0.62*** | 0.64*** | 0.62*** | 0.62*** | 0.33*** |
| *Long-run asymmetry tests (p-values)* | | | | | |
| exchange rate | | | | | |
| $P=N$[b] | 0.061 | 0.082 | 0.057 | 0.028 | 0.001 |
| crude price | | | | | |
| $P=N=Z$[c] | 0.003 | 0.000 | 0.000 | 0.000 | 0.000 |
| $P=N$[d] | 0.020 | 0.020 | 0.010 | 0.003 | 0.000 |
| $P=Z$[e] | 0.008 | 0.002 | 0.003 | 0.005 | 0.000 |
| $N=Z$[f] | 0.811 | 0.687 | 0.725 | 0.972 | 0.967 |

[a] Asterisks stand for significance at 10% (*), 5% (**) or 1% (***).
[b] Test of coefficient equality for positive and negative variations.
[c] Test of coefficient equality in the three regimes.
[d] Test of coefficient equality in the positive and negative regime.
[e] Test of coefficient equality in the positive regime and the inaction band.
[f] Test of coefficient equality in the negative regime and the inaction band.

Figure(s)

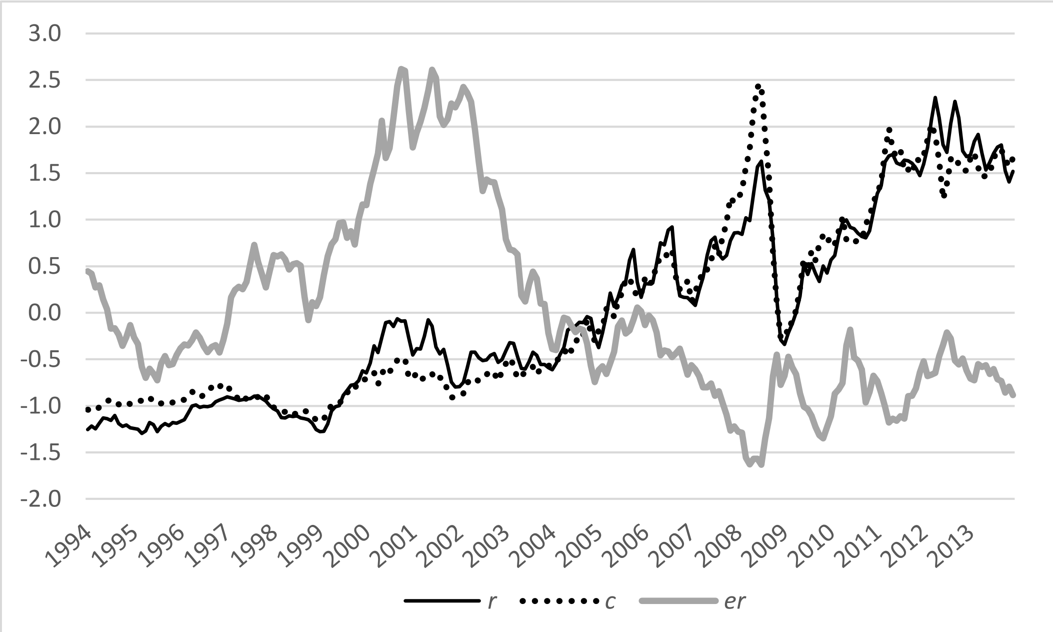


Figure 1 – Standardized values of pre-tax gasoline retail price (*r*), average crude price (*c*), and nominal exchange rate (*er*).

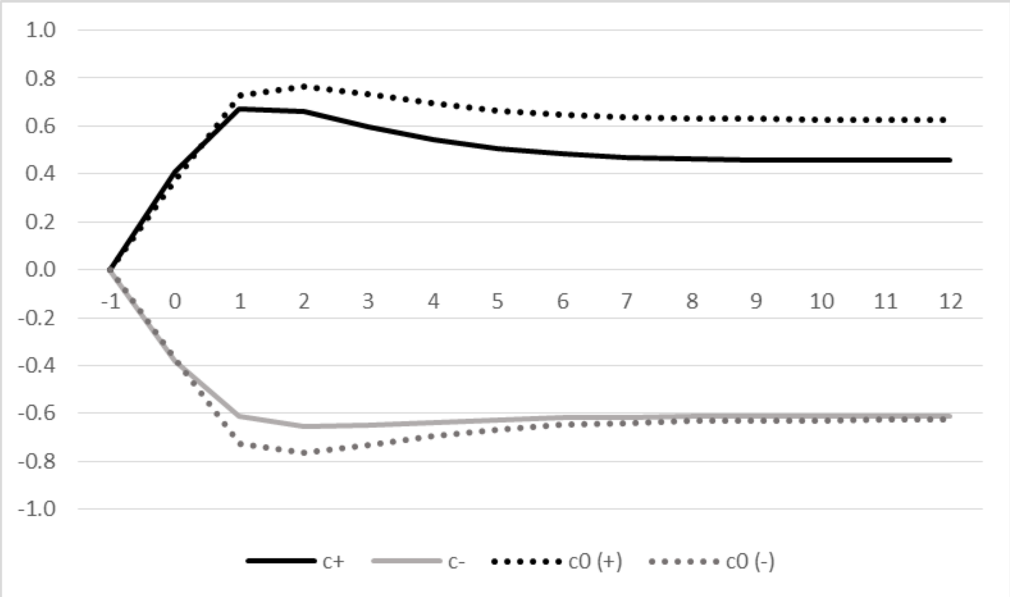


Figure 2 – Dynamic multipliers for 1% crude price shocks. The dotted line plots the response to a shock falling within the inaction band.

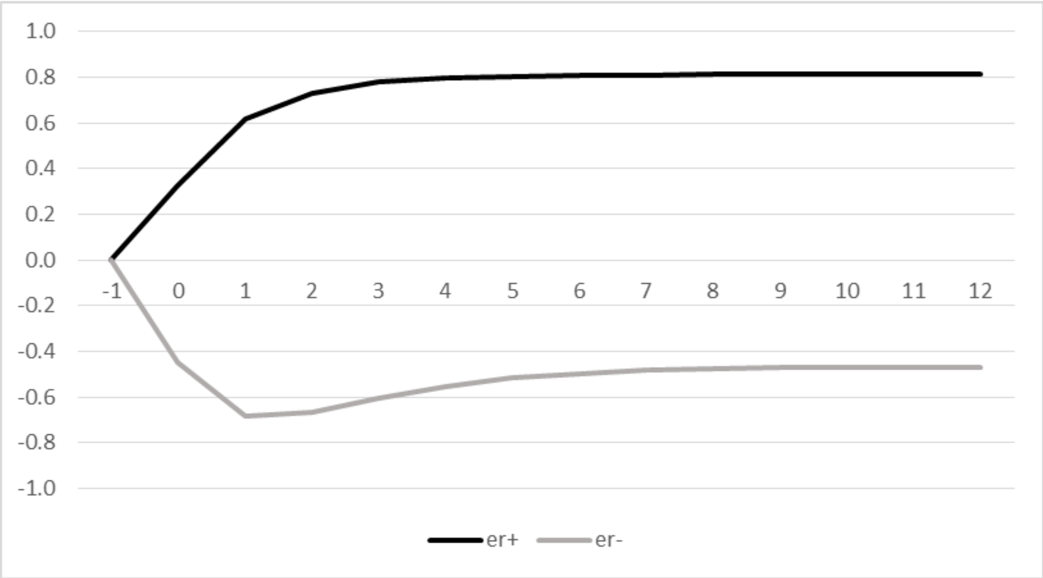


Figure 3 – Dynamic multipliers for 1% exchange rate shocks.