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PRICE ASYMMETRIES IN THE EUROPEAN GASOLINE MARKET

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Abstract – Building on the well-established "rockets and feathers" literature, and on the recently developed nonlinear autoregressive distributed lag (NARDL) modelling, we investigate the asymmetries in gasoline pricing on a comprehensive sample of monthly data from twelve Eurozone countries running from 1994:1 to 2014:12. The empirical results feature two robust patterns. Firstly, while the effects of exchange rate variations display a positive asymmetry (i.e., devaluations have a greater impact with respect to revaluations), crude price variations induce negative asymmetry (i.e., reductions in the price of crude oil have a greater impact than price rises). Secondly, the positive asymmetry to exchange rate changes is much stronger in core Eurozone countries. The negative asymmetry with respect to crude oil prices confirms the results of recent empirical research and theoretical models.

JEL classification: C22, D43, D82, E31, L71, Q41.

Keywords: asymmetric cointegration, asymmetric price adjustment, pass-through, gasoline price, European gasoline market, signaling.

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1. Introduction

Gasoline price has been subject to numerous empirical studies, usually falling into one of the following categories (Eckert, 2013): crude oil or wholesale price pass-through; Edgeworth cycles; the impacts of mergers or regulation and price dispersion; price differentials across individual stations. As for the first point, on which this paper will focus, the pervasiveness of asymmetry in the gasoline market has been recently documented by Perdiguero-García (2013) through an extensive meta-analysis. However, asymmetric price adjustment is not peculiar to the gasoline market: Peltzman (2000) studies 242 products (77 consumer goods and 165 producer goods) and finds asymmetric price reaction for the majority of them; Frey and Manera (2007) carry out a meta-analysis on the econometric models of asymmetric price transmission in different markets (gasoline, agriculture, alimentary) and show that only a small fraction of the estimated models presents no asymmetry.

Several reasons explain why price asymmetries have received a special attention in the market of crude-derived fuels: 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. The widespread perception among the public at large is that asymmetry in gasoline pricing follows a "rockets and feathers" pattern (from Bacon, 1991), i.e., that prices rise faster in response to costs increases than they fall in response to costs decreases (positive asymmetry). The "rockets and feathers" hypothesis (RFH) finds some support in the empirical literature[1]: Peltzman (2000) says that it is a "stylized fact", while Bastianin et al. (2014) consider it an important factor that should be taken into account in forecasting gasoline prices. However, evidence on the RFH is actually quite "mixed and sometimes contradictory", as Contín-Pilart et al. (2009) put it, and it is fair to conclude that a consensus on the causes, the size and the sign of asymmetries in gasoline pricing has not been reached.

As far as the four main Eurozone gasoline markets are concerned, Karagiannis et al. (2014) find that gasoline and diesel prices adjust symmetrically to cost shocks. However, their study shares with the previous literature a crucial feature, namely, it allows for asymmetric responses only in the short-run. More recent analyses resort to the asymmetric cointegration approach by Shin et al. (2014) that allows for asymmetry in both the short- and long-run responses. These studies find a negative asymmetry with respect to crude price (e.g. Atil et al., 2014). While disproving the RFH, this result is consistent with endogenous mark-up models à la Taylor (2000): the theoretical implication is that asymmetry does not originate from competitive behavior, as in Peltzman (2000), but is rather the consequence of an oligopolistic market structure. When the analysis is pushed further, by disentangling the impact of crude price from that of the exchange rate, "asymmetric asymmetries" emerge: while the asymmetry to

[1] A summary can be found in Table 1 of Kristoufek and Lunackova (2015).

crude price (in USD) is confirmed to be *negative*, the asymmetry to exchange rate variations is *positive*, i.e., gasoline prices in local currency responds more to a national currency devaluation (namely, an increase in the domestic cost of foreign currency), than to a revaluation.

Irrespective of its causes, the presence of "asymmetric asymmetry" implies that the inconclusiveness of previous studies might depend on two sources of bias. Firstly, the neglect of long-run asymmetries (a point already raised by Honarvar, 2009), which implies that the estimated models generally impose the untested assumption that the long-run elasticities are symmetrical (i.e., equal for positive and negative shocks). Secondly, the fact that some of these studies express both crude oil price and gasoline price in a common currency, be it local currency or USD, before the estimation.[2] In so doing, they force two different asymmetries to conflate in the same parameters, or, to put it in another way, they impose the untested hypothesis 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. Regardless of the presence of "asymmetric asymmetry", this hypothesis appears to be warranted only in the long-run (Warmedinger, 2004), and it is disproved by studies on the aggregate pass-through behavior (such as Campa and Goldberg, 2005), where the "shifters" (i.e., the proxy of marginal costs) are usually found not to have the same long-run coefficients as the exchange rate.

In order to assess the actual impact of these sources of misspecification on the analysis of gasoline pricing, a more extensive empirical study is needed. In this paper, we apply a recently proposed asymmetric cointegration estimator (Shin et al., 2014), by considering separately the effects of changes in the crude oil price and the exchange rate on the pre-tax retail gasoline price in twelve Eurozone countries, using monthly data from 1994:1 to 2014:12.[3] This will allow us to verify whether the negative asymmetry with respect to crude oil price is confirmed and to assess whether "asymmetric asymmetries" feature only in the Italian market or are a more widespread phenomenon. Moreover, since the Eurozone countries in general depend in a relatively similar way on foreign sources of fossil fuels,[4] but come from very different historical experiences, in particular as far as inflation and the management of their previous national currencies are concerned, it is of some interest to investigate whether the pass-through from exchange rate to gasoline prices follows the same pattern in both "strong" and "weak" Eurozone members, or instead if some patterns emerge that could shed some light on the role of the consumers' perception of the currency strength.

[2] See e.g., Meyler (2009), Rodrigues (2009), Clerides (2010), Silva et al., (2013), Venditti (2013), Karagiannis et al. (2014), and Kristoufek and Lunackova (2015).

[3] We refer to the EA12 definition: Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal and Spain.

[4] The percentage of imports of crude oil on total petroleum consumption in the EA-12 countries was on average 83% during the period 1994-2013 (elaboration based on thousands of barrels per day obtained from the U.S. Energy Information Administration: http://www.eia.gov/countries/data.cfm).

The remainder of the paper is structured as follows. Section 2 provides an overview of previous findings with a focus on European markets. Section 0 presents the data and methodology used to take into account asymmetries. Section 0 contains the estimation results, which are then discussed in Section 0. Finally, Section 0 concludes and draws some policy implications on the grounds of the paper's findings.

2. Evidence on European markets

While research on asymmetric gasoline price adjustment has mainly focused on the U.S. market, relatively less attention has been dedicated to European countries as can be inferred from the summaries in Frey and Manera (2007),[5] Wlazlowski et al. (2008; Table 1), Clerides (2010; Table 1), Polemis (2012; Table 1), Perdiguero-García (2013; Table 1); Kristoufek and Lunackova (2015; Table 1).

Table 1 summarizes the previous results on asymmetric price adjustment in major European countries obtained by Galeotti et al. (2003), Grasso and Manera (2007) and Clerides (2010).[6] The reasons why we are referring to these three studies are: a) Galeotti et al. (2003) is the first study that analyzes the relationship between the price of gasoline, the price of crude oil and the exchange rate in European countries using an asymmetric error correction framework; b) Grasso and Manera (2007) update Galeotti et al. (2003) by using more recent data and econometric methodology;[7] c) Clerides (2010) undertakes a very complete analysis on European markets, covering all the EU-27 country members.[8]

[Table 1 around here]

It can be clearly seen in the previous Table that the estimates are very heterogeneous. The estimates by Galeotti et al. (2003) display asymmetries which come from different responses to crude oil prices (except in Italy, where no price asymmetry is present) and, to some extent (in two out of four countries), to the error correction term. While asymmetry to prices shows the expected signs, the speed of adjustment coefficients do

[5] Frey and Manera's work, besides gasoline, includes agricultural and alimentary markets. 30 out of 114 studies on gasoline deal with European markets, as results from their Tables 6, 7, 8 and 9.

[6] More studies on European markets, can be found in Karagiannis et al. (2014), Kristoufek and Lunackova (2015), Polemis and Fotis (2013), Polemis and Fotis (2015), Rodrigues (2009), Venditti (2013), Wlazlowski et al. (2009). Individual country studies, or analyses with a focus on a particular country, can be found, among others, in Asane-Otoo and Schneider (2015; Germany), Bermingham and O' Brien (2010; Ireland and UK), Bettendorf et al. (2003; Netherlands), Lamotte et al. (2013; France), Polemis (2012; Greece), Silva et al. (2013; Portugal).

[7] Besides asymmetric-ECM (A-ECM), they use threshold-ECM and ECM with threshold cointegration. Moreover, in the A-ECM estimates, they include richer short term dynamics (i.e., autoregressive terms and delays of the differences of explanatory variables), not present in Galeotti et al. (2003).

[8] In Table 1 results for the United Kingdom are not reported (which are available in all the three studies), as our focus consists in the Euro area main countries (EA-12). For the same reason, we omit Clerides' estimates for the remaining EU-27 countries that are not EA-12 countries.

not (coefficients below -1 implies an "over-correction") and the magnitude of responses to disequilibrium is the opposite in the two countries for which a significant and asymmetric error correcting term exists (it is bigger in France for positive shocks, while it is bigger in Germany for negative shocks). No tests for asymmetry is reported for the exchange rate, though we can notice that the signs are different from what we would expect.

The two sets of results by Grasso and Manera (2007), obtained respectively with an asymmetric-ECM (model 1) and a threshold-ECM (model 2), are markedly different from each other. Asymmetries to prices are present only in Spain with model 1, while they exist only in France with model 2; the exchange rate impacts differently (and in general with unexpected signs) if a devaluation or appreciation occurs with model 1, but this vanishes for France and Spain in model 2; a symmetric error correcting term is obtained for all countries with model 1, while asymmetry is obtained for the majority of countries in model 2.

Results by Clerides (2010) are not strictly comparable with those in the rest of the Table as the exchange rate is not explicitly taken into account. Indeed, the author estimates the gasoline price-crude oil price relationship after having converted both prices in euros. Thus, the price elasticities are a convolution of the elasticities to crude oil price and to the exchange rate. However, we reported these estimates because, as said before, they represent one of the most complete studies on European markets, as far as the number of countries are considered. Clerides' findings indicate very few cases in which some asymmetry is present (namely, in both the short-run and the error correction parameters in Finland and Portugal, and only in the error correcting term in Belgium). We believe that, by not including the exchange rate as an additional explanatory variable, Clerides' results could incorporate some bias. This is somehow confirmed by looking at the results in the table which included the exchange rate. We will investigate further this hypothesis in Section 4.

3 Material and methods

3.1 Data

Gasoline pump prices in euros (net of duties and taxes) were obtained from European Commission (2014).[9] Except for Austria, Finland and Luxembourg, the gasoline price series start in 1994:1.[10]

[9] Before the euro changeover (1st January 2002), the prices were reported in national currencies. We converted them in euros by using the ECU-EUR/national currency exchange rate. Some data for Portugal were integrated using the database provided by the Portuguese Directorate General for Energy and Geology (Ministry of Environment, Spatial Planning and Energy, 2014).

[10] The starting date is 1995:1 for Austria and Finland, and 1996:11 for Luxembourg.

The crude oil price series is the Brent price expressed in dollars per barrel. Its source is the U.S. Energy Information Administration and was obtained through FRED (Federal Reserve Economic Data, 2014).

The exchange rates were computed from data available in the PERS (Pacific Exchange Rate Service, 2014) database. For each country, we took the last available observation and retropolated it with its rate of variation in period *t*-1 of its local currency unit (LCU) exchange rate versus the US dollar up to the first observation in 1994, i.e., $er_t = er_{t-1}(er_t^n/er_{t-1}^n - 1)$, for $t = T-1, T-2, \ldots, 1$, where er_t is the USD/EUR exchange rate, $er_T = 0.81$, and er_t^n is the USD/LCU exchange rate. According to PERS, the latter "[p]rior to euro adoption […] are historical series, and following euro adoption these are pseudo rates imputed by applying the euro locking rate to the current euro exchange rate. Pre-1999 [2001 for Greece] data for the euro are the official ECU basket rates, not imputed pseudo rates." It follows that we obtained 12 USD/EUR-ECU exchange rates that are the same after the euro adoption and present some differences during the period 1994-1998 that reflect variations in the exchange rate of national currencies towards the USD.

Figure 1 displays the cross-country average of gasoline prices (along with a ±2 standard deviation interval) and the price of crude oil; Figure 2 also reports the cross country average gasoline prices (and intervals) along with the average of the USD/EUR-ECU exchange rates (and ±2 s.d. interval). Table 2 presents some basic descriptive statistics of the series.

[Figure 1 around here]

[Figure 2 around here]

[Table 2 around here]

An interesting feature of the data can be gathered by the statistics on increases and reductions (indicated on the Table by "% $\lessgtr X$", where X is a threshold). While the percentage of positive gasoline price variations is nearly 54% (the highest being 61% in Spain and the lowest 50.2% in Austria), the percentage of increases of the crude oil price is almost 60%. Moreover, observations that can be considered as outliers (above *UIF* and below *LIF*; see footnote a, Table 2) are preeminently price decreases: the percentage of consistent reductions is on average (2.6%) more than twice the percentage of strong increases (1.1%). This pattern in the data can be easily seen from Figure 3.

[Figure 3 around here]

3.2 Methodology

3.2.1 Cointegration and error correction representation

We will start by assuming that the long term relationship between the log of pre-tax gasoline retail price in euros, r_t, the log price of crude oil in dollars, c_t, and the log of the USD/EUR exchange rate, er_t, can be expressed with the following linear model:

$$r_t = \alpha + \beta_1 c_t + \beta_2 er_t + \zeta_t \qquad \textbf{(1)}$$

where β_1 and β_2 are coefficients that represent the long-term elasticities.[11]

In order for Eq. (1) to represent a meaningful long term relationship the standard approach to cointegration requires that all the variables involved have the same order of integration and that the residual term, ζ_t, is stationary. Once both conditions are satisfied (thus, ζ_t is the cointegrating residual), the following error correction model (ECM) can be estimated

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

where 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 price to crude price and exchange rate, respectively).[12]

The previous two-step approach gives us the long-run relationship, Eq. (1), and associated short-run dynamics, Eq. (2). It is possible to use the auto-regressive distributed-lag (ARDL) modelling approach of Pesaran and Shin (1999) to have an estimate of both the long-run and the short-run relationships in a single step. Doing so consists in estimating a conditional ECM by means of an ARDL model which can take the following form

$$\Delta r_t = \alpha + \rho r_{t-1} + \theta_1 c_{t-1} + \theta_2 er_{t-1} + \sum_{j=1}^{p-1} \gamma_j \Delta r_{t-j} + \sum_{j=0}^{q-1} \left(\pi_{1j} \Delta c_{t-j} + \pi_{2j} \Delta er_{t-j} \right) + \varepsilon_t \qquad \textbf{(3)}$$

where ρ, γ_j and π_{ij} have the same meaning as before, and the long-term elasticities can be obtained as

[11] Thus, we are estimating what Galeotti et al. (2003) call "single stage" model.

[12] While we used the same order of lags for both explanatory variables in the ECM, Eq. (2), this assumption can be easily relaxed in practical applications.

$$\beta_1 = -\theta_1 / \rho; \quad \beta_2 = -\theta_2 / \rho \tag{4}$$

Estimating Eq. (3) is more convenient with respect to the estimating Eq. (1) and Eq. (2), as it consists of a single step. Moreover, when there is not a clear cut distinction in the order of integration of the variables, i.e., whether they are *I*(1) or *I*(0) processes, one can rely on the bounds testing approach of Pesaran et al. (2001) who tabulate critical values for different combinations of *I*(1) and *I*(0) variables.

3.2.2 Asymmetry

In the standard cointegration approach the dependent variable responds in the same way to both increases and decreases in each explanatory variable. In Eq. (3), for instance, the impact effect of a variation of the crude oil price is the same both for positive and negative shocks, π_{10}; a similar conclusion can be stated about the long-term effect, β_1.

The asymmetric cointegration approach proposed by Shin et al. (2014) uses a nonlinear ARDL (NARDL) model, 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.[13] This approach leads to the following formulation

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

where the superscripts "+" and "–" indicate, respectively, positive and negative changes in the variables, and the short- and long-run coefficients differ for positive and negative changes. Explanatory variables are expressed as partial sum processes of positive and negative changes, respectively:

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

[13] The NARDL framework has been applied in various fields where asymmetries can play an important role: exchange rate pass-through to consumer prices (Delatte and López-Villavicencio, 2012) and export prices (Fedoseeva and Werner, 2015); the response of housing prices to macroeconomic fundamentals (Katrakilidis and Trachanas, 2012); the relationship between government expenditures and revenues (Athanasenas et al. 2014).

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

where x_t indicates a generic variable (in our case it represents c_t or er_t). 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:

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

Thus, while in the standard ECM the responses to positive or negative shocks 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 variable.

The generic asymmetric long-run coefficients in Eq. (5) are calculated as:

$$\beta_i^+ = -\theta_i^+ / \rho; \ \beta_i^- = -\theta_i^- / \rho \qquad \textbf{(6)}$$

3.2.3 Testing for pass-through and asymmetry

The (N)ARDL model allows the investigation of a series of interesting economic hypothesis such as the degree of pass-through of cost shocks to prices, and the existence of a short- or long-run asymmetric price adjustment.

By considering the *i-th* variable (where, in our case, 1 = crude oil and 2 = exchange rate), a full pass-through of costs shocks in the long run in the symmetric case can be assessed by testing whether the following hypothesis holds:

$$H_0: \ \beta_i = 1$$

where β_i is obtained as in Eq. (4). A similar test can be constructed for the NARDL model, though in this case given that the pass-through can explicitly follow from positive or negative cost shocks, in the previous hypothesis β_i must be replaced with β_i^+ or β_i^-, respectively, which are calculated as in Eq. (6).

Short-run asymmetry exists if there is no evidence for rejecting the following hypothesis:

$$H_0: \quad \sum_{j=0}^{q-1} \pi_{ij}^{+} = \sum_{j=0}^{q-1} \pi_{ij}^{-}$$

A long-run asymmetric price adjustment can be assessed by testing whether the following hypothesis holds:

$$H_0: \quad \beta_i^{+} = \beta_i^{-}$$

In the previous tests, the alternative hypothesis is always two-sided. To save space we will not carry out one-sided tests like, e.g., the overshooting of depreciations (H_1: $\beta_i^{+}>1$). A simple overview of the coefficients' magnitude will give an idea of the appropriate alternative.

4 Results

While it is not strictly necessary to assess whether the variables involved in the estimates are *I*(0) or *I*(1), as we use the bounds testing approach of Pesaran et al. (2001) to cointegration, their order of integration should not exceed one. Unit root tests are reported in Table 3 and indicate that all the variables are *I*(1).

[Table 3 around here]

4.1 Symmetric cointegration estimates

Estimates of Eq. (3) are reported in Table 4, along with the cointegration tests, and the diagnostic tests.[14]

In the ARDL 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

[14] We experimented with values for *p* and *q* in the interval (1,…,6) and found that the couple of values which balanced the parsimony of the model (by minimization of the BIC criterion) and the whiteness of residuals is $p = 1$, $q = 2$. This said, in some models some traces of residual autocorrelation and heteroskedasticity are present. In order to cope with this issue, all models have been calculated by using HAC (Newey-West) covariances. It should be said that adding additional lags did not provide a significant improvement and, most importantly, results are not dramatically affected by the choice of lags.

feedback coefficient ρ 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. (3) in levels, in the same way as Pesaran et al., 2001. Both statistics are reported in Table 4, indicated respectively as *t_BDM* and *F_PSS*, and are strongly significant, except for Portugal, thus rejecting the null of non-cointegration. Given that our data contain exclusively *I*(1) variables, we also estimated Eq. (1) by using the Phillips and Hansen's (1990) Fully Modified OLS (FMOLS) estimator and calculated Engle and Granger's CRADF statistic: in this case, the null of no-cointegration is found to have no support at any meaningful level of confidence in all countries but Austria where the null is rejected at 1%.

In order to exclude the presence of multiple cointegrating relationships among r_t, c_t and er_t, Johansen's trace and maximum eigenvalue tests have been carried out and are indicated in the Table, respectively, as *Trace* and *Max. Eig.*: while no country is found to have a multiple cointegrating vectors, for Austria both tests say that there is not a cointegrating relation.

[Table 4 around here]

In all cases the equations present expected signs: the speed of adjustment is negative and impact coefficients are positive.

Long term elasticities obtained as in Eq. (4), are reported in Table 5, along with tests for unitary elasticity. In all cases, the elasticities have the expected (positive) sign. Unitary elasticities to crude price are rejected in all countries, while unitary elasticities to exchange rate cannot be rejected in Greece and Portugal.

[Table 5 around here]

Estimates in Table 4 assume that short- and long-term effects are the same for positive and negative variations. We open for to the possible presence of asymmetric price adjustment.

4.2 Asymmetric cointegration estimates

Estimates of Eq. (5) are reported in Table 6 along with cointegration tests, long term elasticities, diagnostic tests, and tests for asymmetries on both short- and long-run elasticities.

[Table 6 around here]

In all countries but Portugal a cointegration relationship is confirmed to exist as can be inferred by the *t_BDM* and *F_PSS* statistics present in the previous Table.[15] Asymmetry tests in the (lower panel) indicate that while there is not an asymmetric short-run response to shocks to explanatory variables at any reasonable confidence level, long-run asymmetries are likely to be present both for the exchange rate and crude oil price, the only exceptions being Greece, Netherlands and Portugal. This evidence can be better represented by looking at the dynamic multipliers reported in Figure 4 and Figure 5, which are, respectively, the responses of a 1% shock in the price of crude oil and in the exchange rate for countries where an asymmetric cointegrating relationship holds.

[Figure 4 around here]

[Figure 5 around here]

5 Discussion

The estimates in the previous sections show that a stable cointegrating relation between gasoline price, crude oil price, and exchange rate exists in almost all Eurozone countries, the only exception being Portugal. In two of these countries (Greece and Portugal) there is no sign of any asymmetric price adjustment, while in the remaining nine we found that shocks to the crude oil price and to the exchange rate are symmetric in the short-run but asymmetric on the long-run.

As for the adjustment of gasoline prices to crude oil prices, in Figure 6 we report the average and country specific long term effects. On average, a 1% positive shock

[15] Shin et al. (2014) propose to adopt the "bounds testing" approach by Pesaran et al. (2001), using their tabulated critical values. However, given that variables are decomposed into two partial sum processes, 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. (2014) remark that by considering the smallest number of regressors the *F*-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. By contrast, the *t*-test is more conservative by considering the highest number. The highest critical value (in absolute value) for models with unrestricted intercept and no trend at the 1% significance level are 6.36 for the *F*-test (two regressors) and −4.6 for the *t*-test (four regressors), which implies that in our case the null of non-cointegration is very strongly rejected, the only exceptions being the *F*-test for Netherlands which is, however, significant at the 5% (the critical value is 4.85), and both the *F*-test and *t*-test for Portugal, neither significant even at the 5%.

generates an increase of 0.48%, while a 1% negative shock results in a reduction of 0.65%. To put it simply, in the long-run there is a negative asymmetry: the effect of reductions of the price of crude oil is stronger than that of increases. As the Figure shows, this result is relatively common to all countries.

[*Figure 6* around here]

With respect to the exchange rate, Figure 7 displays the average and the country specific long term effects. It is evident that the long term effect of an increase (i.e., a devaluation) is strongly different from that of a decrease, and the difference between positive and negative long term effects is higher than that associated to crude oil price variations: a 1% devaluation implies an average pass-through to gasoline prices of 0.88%, while a 1% revaluation leads to a 0.40% reduction. Again, and allowing for a greater heterogeneity, this is an overall consistent result across countries.

[Figure 7 around here]

An interesting result comes from the different effects that the exchange rate exerts on different groups of countries. To keep the exposition simple, we consider only the four main gasoline markets: France and Germany are grouped in the "Core" group, while Italy and Spain are included in the "Periphery" group.[16] The two group's effects, as well as the average effects, is shown in Figure 8. While both the short-run dynamics and the long term effect to an appreciation of the exchange rate is not substantially different in Core and Periphery countries, they differ consistently when it comes to a depreciation: in the Core there is a slight overshooting in the short-to-medium term after the depreciation occurs (almost 1.1 on average from the second to the fifth month) to converge to an almost unitary long-term elasticity (0.95), implying a full long-run pass-through of a devaluation; in the Periphery the effect is 0.56 after the month following the shock, and it converges smoothly to its long-term value of 0.73, indicating an incomplete long-run pass-through.

[Figure 8 around here]

[16] These four countries accounted on average for the 76% of the total EA12 gasoline consumption during the period 2000-2013. Source: elaboration on http://www.eia.gov/cfapps/ipdbproject/IEDIndex3.cfm?tid=5&pid=5&aid=2 (last accessed: 2015-06-07).

Another interest finding is that our results confirm that estimates of price response to shocks ignoring the presence of asymmetries may lead to highly biased analysis. Consider Belgium, for instance. Its long-term symmetric elasticities to crude oil price and exchange rate are, respectively, 0.62 and 0.52 (see Table 5). By allowing for asymmetries, we estimated that the long-term elasticity to crude oil price is 0.51 in response to a positive change, while it is 0.7 in response to a price reduction (Table 6); as for the long-term elasticity to exchange rate, the only significant effect is that of a devaluation which is almost 0.9.

In order to shed some light on why some previous studies are not supportive of an asymmetric price adjustment in the European gasoline market (e.g., Clerides, 2010), we estimated Eq. (5) for the nine countries where we found asymmetry by removing the exchange rate and by converting the price of crude oil in euros.[17] Results show that long-term asymmetry becomes and exception as only in Greece, Italy and Spain the coefficients differ significantly. Moreover, short-term asymmetries emerge in Austria, Germany, Ireland (in this case significant only at 10%) and Spain.

5.1 Robustness check

We will now check the robustness of the results obtained in the previous Section by controlling for factors that might influence the price of gasoline. We re-estimated Eq. (5) by including: 1) volatility, measured by the conditional standard deviation of the crude oil price estimated through a GARCH(1,1) model (see Romano and Scandurra, 2012); 2) seasonal idiosyncratic factors, given by monthly dummy variables (see Kaufmann and Laskowski, 2005);[18] 3) consumer's willingness to spend, proxied by the consumer confidence indicator; 4) a specific demand-related variable, measured by the index of passenger car registrations; 5) an economy-wide supply indicator, given by the index of industrial production; 6) the industry petroleum stocks (see Kaufmann and Laskowski, 2005). The alternative models are labeled, respectively, as *Volatility*, *Seasonality*, *Demand 1*, *Demand 2*, *Supply* and *Stocks* in Table 7 and Table 8.[19]

[Table 7 around here]

[Table 8 around here]

[17] To save space we do not report the estimates, but they are available upon request.

[18] The model that accounts for seasonal factors has been estimated also in two alternative ways: 1) by seasonal adjusting the original variables (*r*, *c*, and *er*); 2) with quarterly dummies instead of monthly ones. In both cases, results remain qualitatively similar to the model with monthly dummies.

[19] In *Demand 1*, *Demand 2*, *Supply* and *Stocks* the control variable is seasonal adjusted and has been included both in log-levels and in first log-differences. The consumer confidence indicator (amplitude adjusted), index of passenger car registrations, and the index of industrial production were drawn from the OECD statistical database (http://stats.oecd.org/). The industry petroleum stocks were taken from the EIA (http://www.eia.gov/cfapps/ipdbproject/iedindex3.cfm?tid=50&pid=70&aid=5&cid=&syid=2010&eyid=2014&freq=M&unit=MBBL)

As the previous Tables show, the results are generally unaffected to the different specifications. The main differences are: 1) the inclusion of supply and sectoral specific demand factors (models *Supply* and *Demand 2*) supports the evidence of long-run asymmetry for both crude oil and exchange rate variations in Greece and Netherlands; 2) the asymmetric long term response to crude oil prices is less evident in France, especially when stocks are included.

6 Conclusions and policy implications

Using a recently developed econometric methodology, the NARDL model proposed by Shin et al. (2014), we investigate the presence and nature of asymmetries in the adjustment of pre-tax gasoline retail prices in twelve European countries (Euro area 12). The adoption of a more flexible modelling strategy, which carefully distinguish between the short- and the long-run behaviour, allows us to reconcile some puzzling evidence in this field and to explain some apparent contradictions. Firstly, our study confirms the recent finding by Karagiannis et al. (2014) that the short-run adjustment of gasoline price in the major European markets is symmetric. Secondly, in the vast majority of cases (nine countries, including the four major European markets) NARDL estimates show clearly that symmetry does *not* apply to the long-run response. Thirdly, the negative asymmetry of gasoline price to crude oil changes, found in the US market using the same methodology by Atil et al. (2014), applies, in the long-run, in all the countries considered, and is therefore confirmed to be a widespread phenomenon. Fourthly, the "asymmetric asymmetry", i.e., the coexistence of the negative asymmetry with respect to crude price, with a positive asymmetry with respect to the exchange rate, features in all the countries considered.

These results call for three considerations.On empirical grounds, the consistency and robustness of the results across different markets indicates that the inconclusiveness of previous studies about the existence and size of the asymmetries could actually be determined by the limitations of the methodology adopted. As argued above, by ignoring long-run asymmetries and (in some cases) by expressing gasoline and crude oil prices in the same currency, these studies impose a host of untested assumptions. Since these assumptions generally appear to be rejected by the data, it is extremely likely that imposing them may have biased the conclusions of some previous studies. A similar reasoning can explain why another remarkable empirical feature found in this study, i.e., the widespread presence of "asymmetric asymmetry" is relatively absent in the applied literature.

On theoretical grounds, while contrary to the general public perception, the negative price asymmetry is consistent with a number of theoretical models grounded on oligopolistic or monopolistic competitive behaviour. For instance, in Taylor's (2000) endogenous mark-up model firms lose market power in a low-inflation environment,

which implies that when crude price falls (and inflation declines) the retailers are compelled to compete by adjusting prices more quickly in order to keep their market shares. While further theoretical research is needed, the mechanism outlined by Taylor offers some insights on the "asymmetric asymmetry" behavior. In fact, while the euro has generally been perceived as a stable currency, the price of crude oil, owing to its large swings, has played a major role as a signal of changes in the inflation performance of a country. This asymmetry in consumers' perceptions may explain why retailers have been able to take full advantage of nominal depreciations, by passing them almost fully through to retail prices, while avoiding to adjust prices in response to an appreciation. A further hint in this direction is given by the fact that in the short-run the response to an exchange rate depreciation is more intense in Eurozone core countries — that come from a long tradition of "strong" national currencies — which may have biased the consumers' awareness of the role of exchange rate changes (see Figure 8).

Coming to the policy implications, our results are consistent with the hypothesis that asymmetries stem from oligopolistic behaviour (Balke et al. 2000), as recently confirmed by Perdiguero-García (2013). This implies that in the countries considered by our study the absence of competition in the gasoline retail market is still a relevant policy issue. Another point that our results raise is that the different magnitude of the exchange rate pass-through to gasoline prices in different countries imply a high degree of heterogeneity in the response to crude oil price shocks. Given that the countries considered in this study share the same currency, and conditionally on its survival, this result would call for an improvement in the fiscal coordination within the area.

Our work can be extended in various directions. For instance, an interesting development may be to open for the presence of hysteretic effects in the price adjustment mechanism. Indeed, our estimates consider the existence of only two regimes: positive and negative changes. It could be the case that small or high input costs changes are passed through differently. In order to explore this size dependency, the variations of input costs could be categorized as "large positive shocks", "inaction band" and "large negative shocks", thus producing different sets of short- and long-run elasticities. We leave this extension for future research.

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References

Asane-Otoo, E. and Schneider, J. (2015) "Retail fuel price adjustment in Germany: A threshold cointegration approach", *Energy Policy*, 78, 1-10.

Athanasenas, A., Katrakilidis, C., Trachanas, E. (2014) "Government spending and revenues in the Greek economy: evidence from nonlinear cointegration", *Empirica*, 41(2), 365-376.

Atil, A., Lahiani, A., Mguyen, 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(3), 211-218.

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.

Bastianin, A., Galeotti, M., Manera, M. (2014) "Forecasting the oil–gasoline price relationship: Do asymmetries help?", *Energy Economics*, 46 (Supplement 1), S44-S56.

Bermingham, C., O' Brien, D. (2010) "Testing for asymmetric pricing behaviour in Irish and UK petrol and diesel markets", *Central Bank of Ireland Research Technical Papers* No. 3/RT/10.

Bettendorf, L., van der Geest, S.A., Varkevisser, M. (2003) "Price asymmetry in the Dutch retail gasoline market", *Energy Economics*, 25(6), 669-689.

Campa, J.M., Goldberg, L.S. (2005) "Exchange rate pass-through into import prices", *Review of Economics and Statistics*, 87, 679-690.

Campa, J.M., Goldberg, L.S., González-Mínguez, J.M. (2005) "Exchange-Rate Pass-Through to Import Prices in the Euro Area", *NBER Working Paper* No. 11632.

Clerides, S. (2010) "Retail fuel price response to oil price shocks in EU countries", *Cyprus Economic Policy Review*, 4(1), 25-45.

Contín-Pilart, I., Correljé, A.F., Palacios, M.B. (2009) "Competition, regulation, and pricing behaviour in the Spanish retail gasoline market", *Energy Policy*, 37(1), 219-228.

Davidson, R., MacKinnon, J.G. (2004) *Econometric Theory and Methods*, Oxford University Press: New York.

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

Eckert, A. (2013), "Empirical studies of gasoline retailing: a guide to the literature", *Journal of Economic Surveys*, 27(1), 140-166.

Engle, R.F., Granger, C.W.J. (1987) "Co-integration and error correction: representation, estimation, and testing", *Econometrica*, 55(2), 251-276.

European Commission (2014) *Oil Bulletin*, https://ec.europa.eu/energy/en/statistics/weekly-oil-bulletin (last accessed: 2014-12-26).

Federal Reserve Economic Data (2014) http://research.stlouisfed.org/fred2/series/WCOILBRENTEU (last accessed: 2014-12-26).

Fedoseeva, S., Werner, L.M. (2015) "How linear is pricing-to-market? Empirical assessment of hysteresis and asymmetry of PTM", *Empirical Economics*, doi:10.1007/s00181-015-0957-4

Frey, G., Manera, M. (2007) "Econometric models of asymmetric price transmission", *Journal of Economic Surveys*, 21(2), 349–415.

Galeotti, M., Lanza, A., Manera, M. (2003) "Rockets and feathers revisited: An international comparison on European gasoline markets", *Energy Economics*, 25(2), 175-190.

Grasso, M., Manera, M. (2007) "Asymmetric error correction models for the oil–gasoline price relationship", *Energy Policy*, 35(1), 156–177.

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.

Karagiannis, S., Panagopoulos, Y., Vlamis, P. (2014) "Are unleaded gasoline and diesel price adjustments symmetric? A comparison of the four largest EU retail fuel markets", *Economic Modelling*, 48, 281-291.

Katrakilidis, C., Trachanas, E. (2012) "What drives housing price dynamics in Greece: New evidence from asymmetric ARDL cointegration", *Economic Modelling*, 29(4), 1064-1069.

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.

Kristoufek, L. Lunackova, P. (2015) "Rockets and feathers meet Joseph: Reinvestigating the oil-gasoline asymmetry on the international markets", *Energy Economics*, 49, 1-8.

Lamotte, O., Porcher, T., Schalck, C., Silvestre, S. (2013) "Asymmetric gasoline price responses in France", *Applied Economics Letters*, 20(5), 457-461.

Meyler, A. (2009) "The pass through of oil prices into euro area consumer liquid fuel prices in an environment of high and volatile oil prices", *Energy Economics*, 31(6), 867-881.

Ministry of Environment, Spatial Planning and Energy (2014) http://www.dgeg.pt/wwwbase/wwwinclude/ficheiro.aspx?tipo=0&id=13672&ambiente=WebSiteMenu (last accessed: 2014-12-26).

Ng, S., Perron, P. (2001) "Lag length selection and the construction of unit root tests with good size and power", *Econometrica*, 69(6), 1519-1554.

Pacific Exchange Rate Service (2014) http://fx.sauder.ubc.ca/data.html (last accessed: 2014-12-26).

Peltzman, S. (2000) "Prices rise faster than they fall", *Journal of Political Economy*, 108(3), 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. (1999) "An autoregressive distributed lag modelling approach to cointegration analysis", in: Strom S. (Ed.), *Econometrics and Economic Theory in the 20th Century: The Ragnar Frisch Centennial Symposium*, Cambridge University Press, Cambridge, chapter 11.

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(1), 99-125.

Polemis, M.L. (2012) "Competition and price asymmetries in the Greek oil sector: An empirical analysis on gasoline market", *Empirical Economics*, 43(2), 789-817.

Polemis, M.L., Fotis, P.N. (2013) "Do gasoline prices respond asymmetrically in the euro zone area? Evidence from cointegrated panel data analysis", *Energy Policy*, 56, 425-433.

Polemis, M.L., Fotis, P.N. (2015) "Rent seeking oligopolistic behaviour in European gasoline markets", *Economics Bulletin*, 35(1), 827-833.

Rodrigues, J. (2009) "Asymmetries in the adjustment of motor diesel and gasoline pump prices in Europe", *Portuguese Competition Authority* Working Paper N. 37.

Romano, A.A., Scandurra, G. (2012) "Price asymmetries and volatility in the Italian gasoline market", *OPEC Energy Review*, 36(2), 215–229.

Shin, Y., Yu, B., Greenwood-Nimmo, M. (2014) "Modelling asymmetric cointegration and dynamic multipliers in a nonlinear ARDL framework", in: Horrace, W.C., Sickles, R.C. (Eds.), *Festschrift in Honor of Peter Schmidt*, Springer, New York.

Silva, F., Batista, M., Elias, N. (2013) "Fuel price transmission mechanisms in Portugal", *Applied Economics Letters*, 20(1), 72-75.

Taylor, J. (2000) "Low inflation, pass-through, and the pricing power of firms", *European Economic Review*, 44(7), 1389-1408.

Venditti, F. (2013) "From oil to consumer energy prices: How much asymmetry along the way?", *Energy Economics*, 40, 468-473.

Warmedinger, T. (2004) "Import prices and pricing-to-market effects in the Euro area", *European Central Bank Working Paper*, N. 299, January.

Wlazlowski, S., Giulietti, M., Binner, J., Milas, C. (2008) "Smooth transition models in price transmission", *The Rimini Centre for Economic Analysis Working Paper* N. 04-08.

Wlazlowski, S., Giulietti, M., Binner, J., Milas, C. (2009) "Price dynamics in European petroleum markets", *Energy Economics*, 31(1), 99-108.

Figures

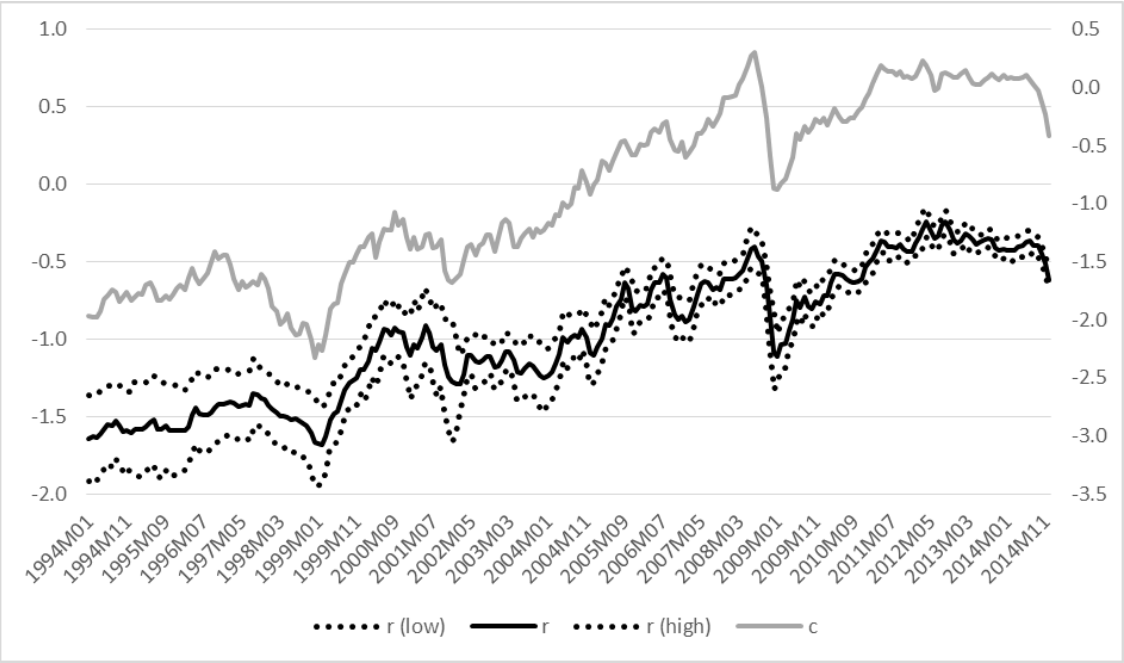


Figure 1 – Logarithms of the cross-country averages of pre-tax gasoline retail prices (r, left-hand scale) and crude oil price (c, right-hand scale); "low"/"high" is given by the average of the corresponding series less/plus twice its standard deviation.

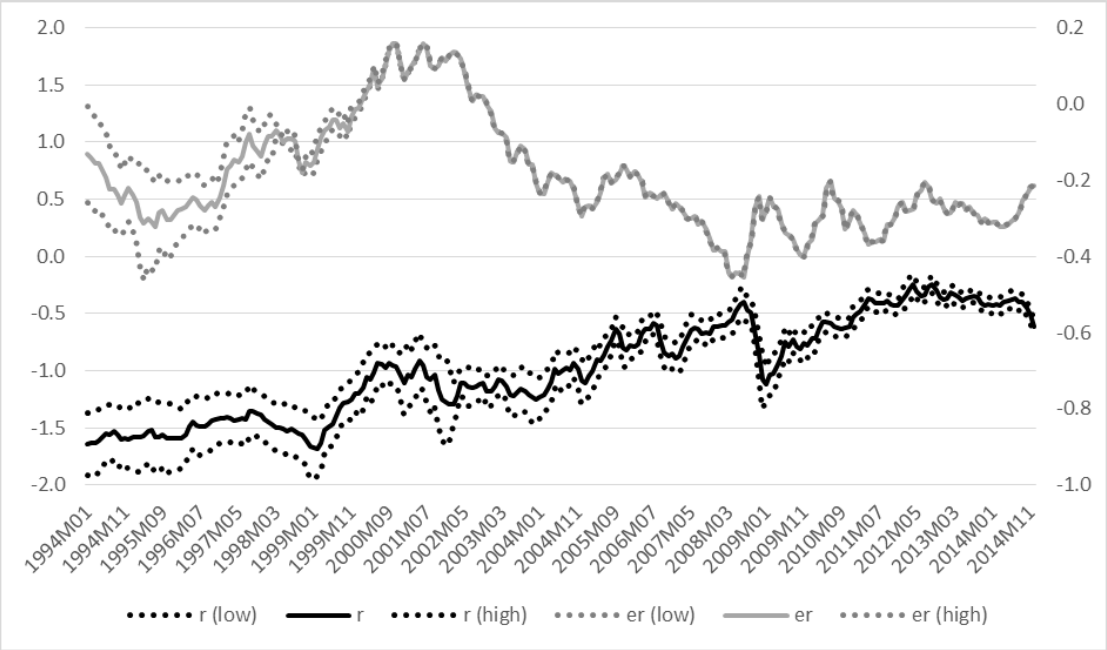


Figure 2 – Logarithms of the cross-country average of pre-tax gasoline retail prices (r, left-hand scale) and USD/EUR-ECU exchange rates (er, right-hand scale); "low"/"high" is given by the average of the corresponding series less/plus twice its standard deviation.

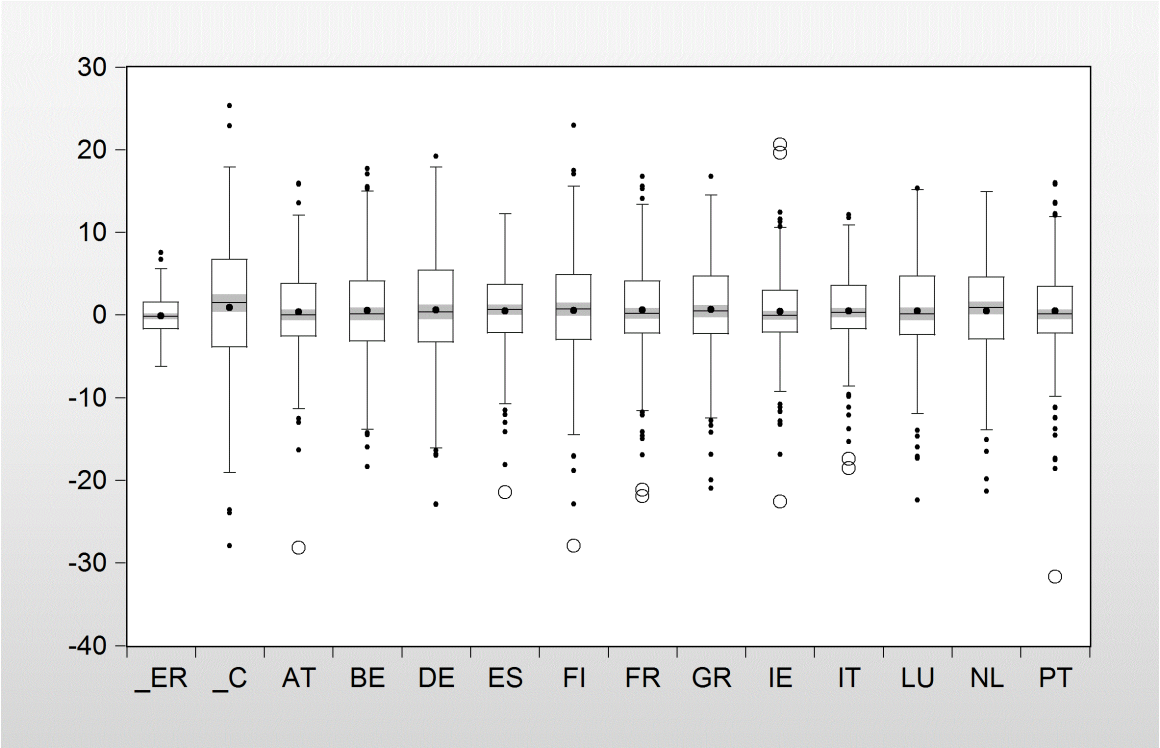


Figure 3 – Boxplot of percentage variations of the exchange rate (_ER), crude oil price (_C), and gasoline pump prices (net of duties and taxes; indicated with the country's acronym); mild and extreme outliers are represented by filled and empty circles, respectively (see footnote a, Table 2).

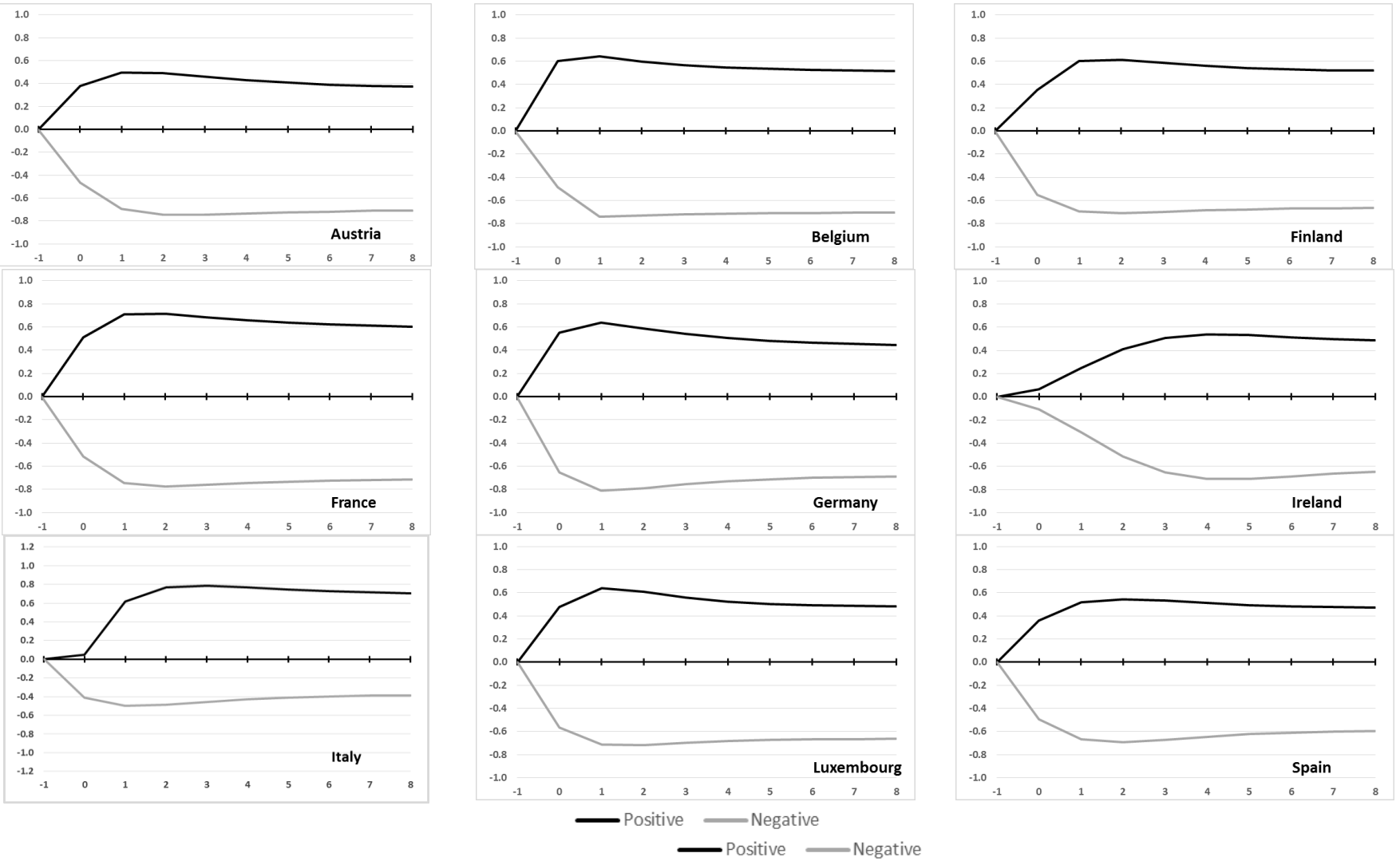


Figure 4 – Dynamic multipliers for 1% crude oil price shock.

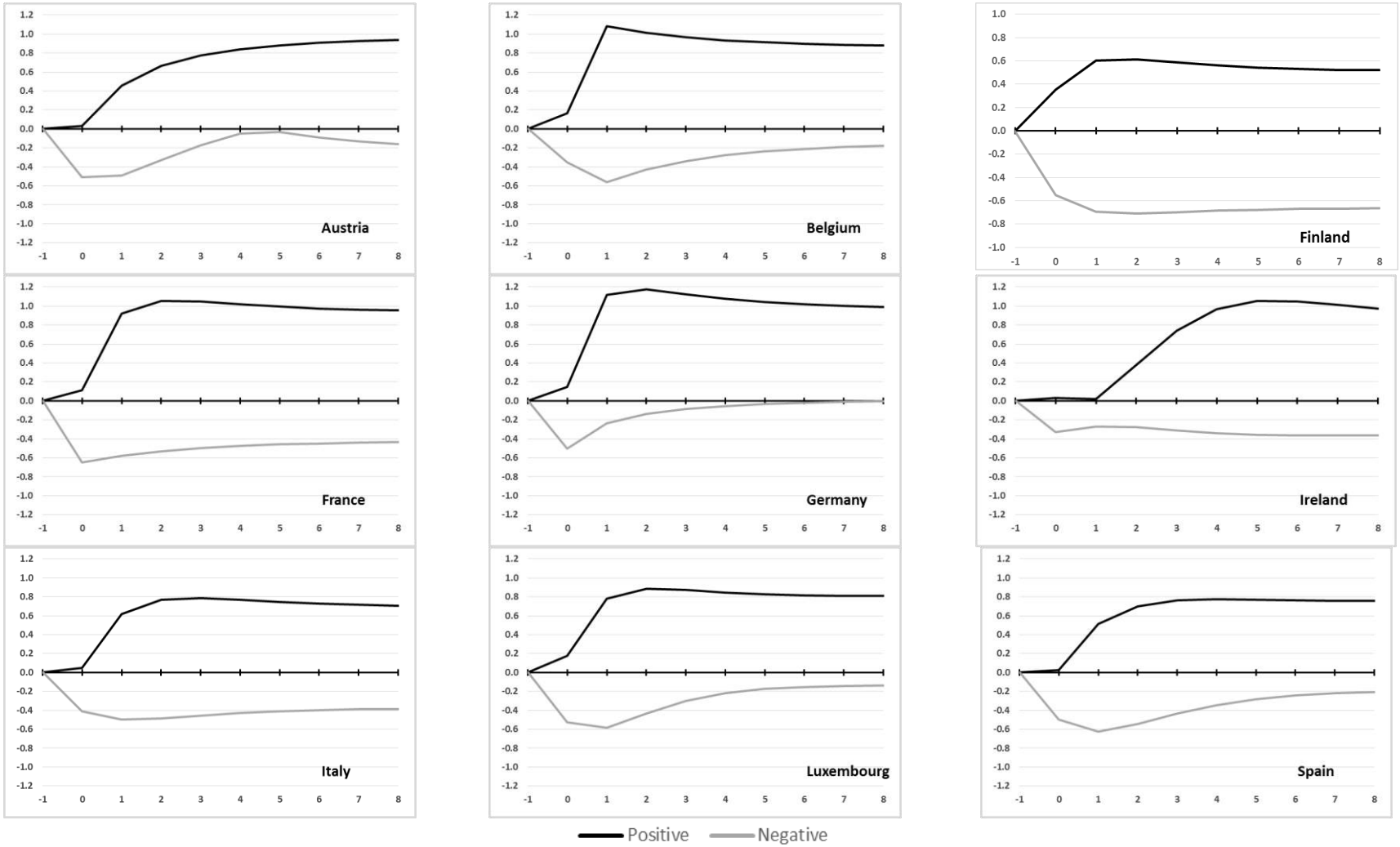


Figure 5 – Dynamic multipliers for 1% exchange rate shock.

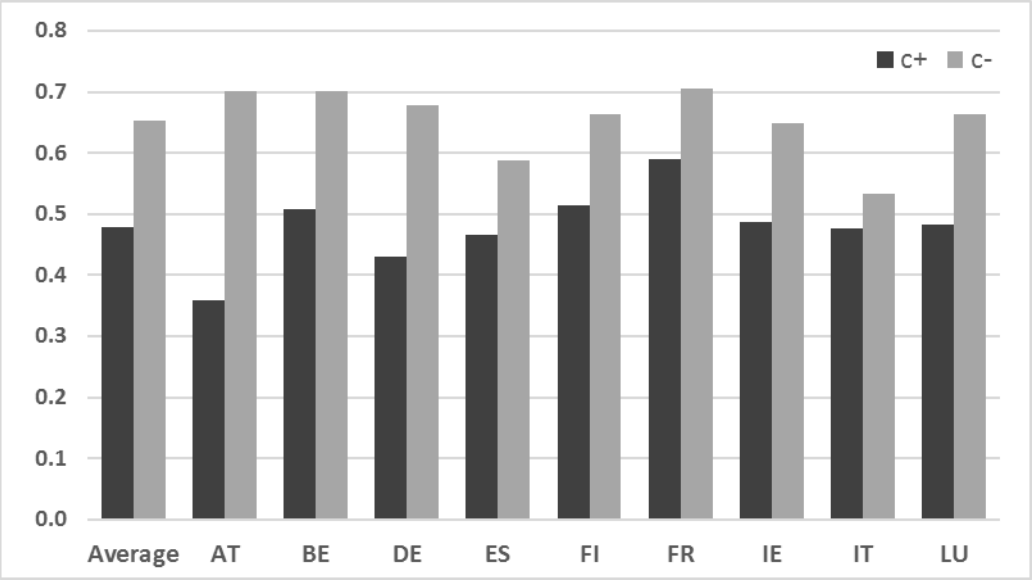


Figure 6 – Long term effect on gasoline prices of a 1% positive (c+) and negative (c−) variations of crude oil price.

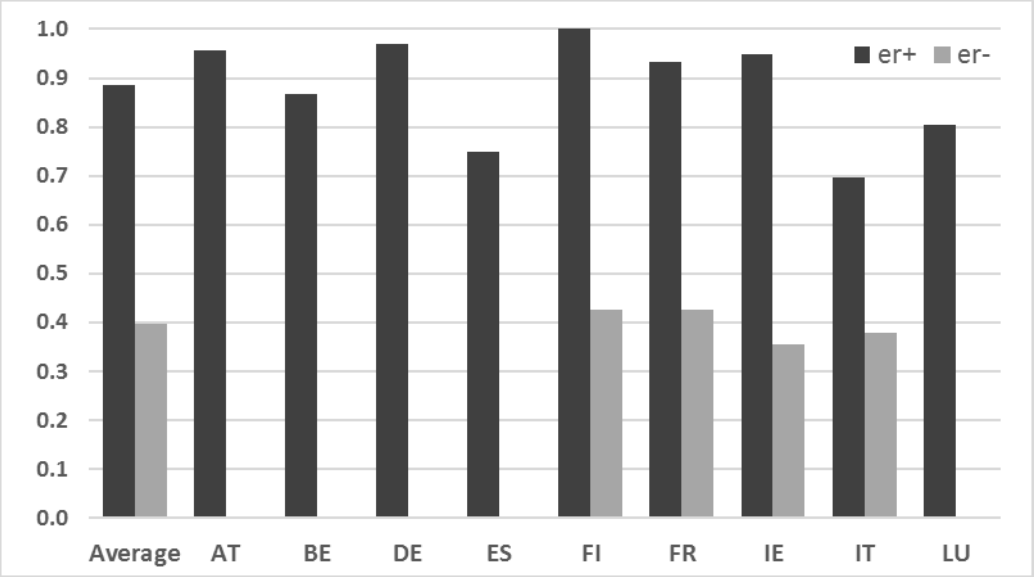


Figure 7 – Long term effect on gasoline prices of a 1% positive (er+) and negative (er−) variations of the exchange rate. Coefficients not significant at 5% are omitted.

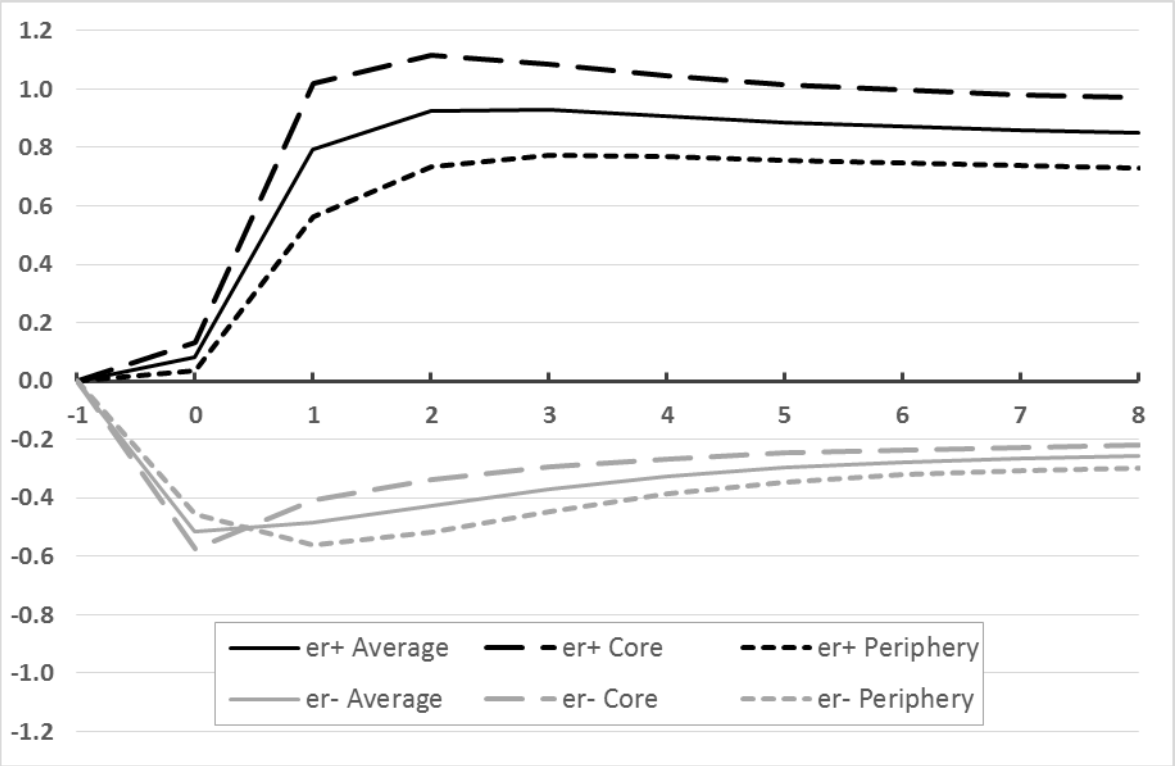


Figure 8 – Dynamic multipliers for 1% positive (er+) and negative (er-) exchange rate shocks: Average, Core and Periphery.

Tables

Table 1 – Previous results on the response of gasoline prices to crude oil price and exchange rate variations in European countries.

	Crude oil price		Exchange rate		Error correction	
	positive	*negative*	*positive*	*negative*	*positive*	*negative*
Galeotti et al. (2003)						
France	**0.56**	**0.16**	*0.32*	*0.69*	**-1.06**	**-0.68**
Germany	**0.79**	**0.55**	*0.25*	*0.41*	**-0.74**	**-1.13**
Italy	0.20	0.24	*-0.06*	*0.46*	-1.37	-1.36
Spain	**0.24**	**0.16**	*0.16*	*0.14*	-1.08	-1.01
Grasso and Manera (2007) [1]						
France	0.44	-0.01	**0.51**	**0.61**	-0.45	-0.18
Germany	0.41	0.38	**-0.22**	**0.50**	-0.41	-0.31
Italy	0.26	0.26	**0.09**	**0.68**	-0.23	0.01
Spain	**0.18**	**0.11**	**0.20**	**0.18**	-0.24	-0.17
Grasso and Manera (2007) [2]						
France	**0.10**	**0.51**	0.54	0.75	-0.38	-0.28
Germany	0.37	0.51	**0.45**	**-1.02**	**-0.30**	**-0.55**
Italy	0.42	0.26	**1.10**	**0.25**	**-0.20**	**-0.06**
Spain	0.33	0.20	0.43	0.25	**-0.78**	**-0.25**
Clerides (2010)						
Austria	0.38	0.49	n.a.	n.a.	-0.03	-0.10
Belgium	0.60	0.44	n.a.	n.a.	**-0.14**	**-0.39**
Germany	0.53	0.61	n.a.	n.a.	-0.12	-0.18
Spain	0.17	0.32	n.a.	n.a.	-0.07	-0.04
Finland	**0.52**	**-0.10**	n.a.	n.a.	**-0.37**	**-0.01**
France	0.41	0.44	n.a.	n.a.	-0.10	-0.01
Greece	0.42	0.48	n.a.	n.a.	-0.14	-0.13
Ireland	-0.12	-0.05	n.a.	n.a.	-0.18	-0.08
Italy	0.31	0.29	n.a.	n.a.	-0.09	-0.09
Luxembourg	0.39	0.68	n.a.	n.a.	-0.10	-0.12
Netherlands	0.26	0.61	n.a.	n.a.	-0.01	-0.11
Portugal	**-0.14**	**0.19**	n.a.	n.a.	**0.00**	**-0.09**

Notes: The Table reports 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); coefficients for which the positive and negative effect are significantly different are reported in bold; Galeotti et al. (2003) do not report the significance of exchange rate coefficients (in italics); Clerides (2010) express both prices of gasoline and crude oil in euros, thus no coefficients for the exchange rate are available; long-run elasticities are not reported by Galeotti et al. (2003) and Grasso and Manera (2007), while Clerides (2010) reports symmetric long-run elasticities; results from Galeotti et al. (2003) and Grasso and Manera (2007) refer to the "single stage" model and coefficients' significance is based on their simulated F-tests (with rejection frequencies greater than 15%, as they indicate on their respective papers); [1] asymmetric-ECM estimates; [2] threshold-ECM estimates (negative estimates refer to the value of the positive coefficient plus the "differential" effect).

Table 2 – Descriptive statistics.

	er	*c*	*AT*	*BE*	*DE*	*ES*	*FI*	*FR*	*GR*	*IE*	*IT*	*LU*	*NL*	*PT*
Mean	-0.06	0.97	0.43	0.61	0.66	0.53	0.61	0.64	0.69	0.48	0.53	0.55	0.56	0.54
First quartile	-1.63	-3.84	-2.49	-3.09	-3.21	-2.08	-2.94	-2.15	-2.22	-2.04	-1.60	-2.34	-2.88	-2.12
Median	-0.11	1.51	0.08	0.20	0.42	0.70	0.76	0.23	0.52	0.01	0.35	0.19	0.92	0.16
Third quartile	1.57	6.75	3.85	4.17	5.43	3.76	4.90	4.14	4.76	3.03	3.65	4.72	4.62	3.52
Maximum	7.60	25.34	15.97	17.78	19.27	12.28	22.97	16.81	16.82	20.67	12.15	15.38	14.95	16.06
Minimum	-6.17	-27.88	-28.10	-18.34	-22.91	-21.37	-27.87	-21.87	-20.95	-22.56	-18.51	-22.37	-21.29	-31.65
Std. deviation	2.36	8.30	5.50	6.65	6.94	5.16	6.65	6.02	5.83	5.26	4.73	6.11	5.88	5.87
Skewness	0.03	-0.42	-0.53	0.10	-0.29	-0.67	-0.42	-0.32	-0.37	-0.13	-0.67	-0.35	-0.39	-0.81
Kurtosis	3.16	3.73	6.10	3.18	3.68	4.53	5.13	4.35	4.09	5.78	4.79	4.05	3.91	6.96
Normality	0.86	0.00	0.00	0.69	0.02	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
% > 0	46.44	58.57	50.21	51.79	54.58	60.96	54.81	51.79	55.38	51.39	54.98	53.00	53.78	52.99
% > UIF	0.84	0.80	1.26	1.99	0.40	0.00	1.26	1.59	0.40	2.79	0.80	0.46	0.00	2.39
% < LIF	0.00	1.59	2.09	1.59	1.99	2.39	2.09	3.59	2.39	3.98	3.19	2.76	1.59	3.98
Observations	239	251	239	251	251	251	239	251	251	251	251	217	251	251

Notes: er and c represent, respectively, the percentage change of the USD/EUR exchange rate and of the price of crude oil, and the remaining columns represent the percentage change of the gasoline pump prices in euros (net of duties and taxes) in country XX, where XX is the country code; "Normality" is the p-value of the Jarque-Bera statistic; % > X and % < X represent the percentage of observations that are, respectively, above or below the threshold given by X; LIF is the "lower inner fence" while UIF is the "upper inner fence", i.e., LIF = Q1−1.5(Q3−Q1) and UIF = Q3+1.5(Q3−Q1), where Q1 and Q3 are, respectively, the first and third quartile.[a]

[a] *LIF and UIF give a rough indication of extreme values: values below LIF or above UIF are considered mild outliers. Extreme outliers are indicated by values below the "lower outer fence" (LOF) and "upper outer fence" (UOF), which are obtained, respectively, as LOF = Q1−3(Q3−Q1) and UOF = Q1−3(Q3−Q1).*

Table 3 – Unit root tests.

	Levels			First differences		
Series	*ADF*	*p-value*	*lags*	*ADF*	*p-value*	*lags*
er	-1.37	0.599	2	-11.54	0.000	0
c	-1.53	0.516	1	-7.41	0.000	2
r_AT	-2.05	0.264	1	-10.56	0.000	0
r_BE	-1.63	0.467	1	-13.63	0.000	0
r_DE	-1.78	0.391	1	-12.66	0.000	0
r_ES	-1.50	0.531	2	-9.87	0.000	0
r_FI	-1.49	0.535	4	-8.78	0.000	1
r_FR	-1.55	0.506	2	-10.98	0.000	0
r_GR	-1.72	0.418	2	-11.25	0.000	0
r_IE	-1.63	0.467	3	-9.76	0.000	0
r_IT	-1.67	0.445	2	-9.83	0.000	0
r_LU	-1.79	0.384	2	-10.50	0.000	0
r_NL	-1.83	0.365	1	-12.62	0.000	0
r_PT	-1.68	0.440	2	-11.40	0.000	0

Notes: The exchange rate (er), crude oil price (c) and the gasoline price in country XX (r_XX) are expressed in logarithms; the test for the variables in levels includes a drift, and the tests for the variables in first differences were performed accordingly with no deterministic component; 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 6 months.

Table 4 – Symmetric cointegration estimates. Dependent variable: Δ(r).

	AT	*BE*	*DE*	*ES*	*FI*	*FR*	*GR*	*IE*	*IT*	*LU*	*NL*	*PT*
constant	**-0.05**	**-0.08**	**-0.06**	**-0.05**	**-0.06**	**-0.06**	**-0.03**	**-0.10**	**-0.05**	**-0.08**	**-0.06**	-0.02
r_{-1}	**-0.11**	**-0.24**	**-0.18**	**-0.16**	**-0.25**	**-0.19**	**-0.18**	**-0.30**	**-0.18**	**-0.26**	**-0.20**	**-0.17**
er_{-1}	0.04	**0.13**	**0.10**	**0.08**	**0.19**	**0.14**	**0.16**	**0.18**	**0.10**	**0.13**	**0.12**	**0.17**
c_{-1}	**0.05**	**0.15**	**0.12**	**0.10**	**0.17**	**0.14**	**0.13**	**0.17**	**0.11**	**0.15**	**0.12**	**0.11**
$\Delta(r_{-1})$	**0.16**	-0.06	0.03	**0.18**	**0.12**	0.11	0.07	**0.40**	**0.16**	0.11	0.09	**0.22**
$\Delta(er)$	**0.33**	**0.32**	**0.40**	**0.30**	**0.54**	**0.42**	**0.41**	0.14	**0.23**	**0.37**	**0.43**	**0.08**
$\Delta(er_{-1})$	0.15	**0.52**	0.27	0.20	0.13	0.23	0.16	-0.20	**0.20**	0.22	0.04	0.26
$\Delta(c)$	**0.43**	**0.54**	**0.61**	**0.45**	**0.46**	**0.52**	**0.51**	**0.08**	**0.40**	**0.52**	**0.55**	**0.32**
$\Delta(c_{-1})$	**0.11**	**0.16**	0.10	0.07	**0.09**	**0.12**	**0.12**	0.02	**0.10**	0.07	-0.02	0.09
Adj. R^2	0.62	0.61	0.62	0.69	0.52	0.68	0.66	0.47	0.71	0.65	0.63	0.40
t_BDM	**-3.84**	**-5.34**	**-4.26**	**-5.11**	**-6.88**	**-4.97**	**-3.82**	**-5.92**	**-4.09**	**-5.82**	**-5.00**	-2.88
F_PSS	**5.12**	**9.60**	**6.38**	**9.00**	**15.80**	**9.08**	**5.59**	**11.80**	**7.14**	**11.79**	**8.73**	4.16
CRADF	-4.26	-6.31	-4.95	-4.82	-5.78	-5.68	-5.70	-7.58	-5.87	-5.53	-5.23	-5.26
Trace	0	1	0	1	1	1	1	1	1	1	1	1
Max-Eig.	0	1	1	1	1	1	1	1	1	1	1	1
*SC(4) **	0.691	0.590	0.132	0.029	0.013	0.121	0.072	0.000	0.063	0.284	0.198	0.791
*SC(12) **	0.082	0.068	0.002	0.001	0.000	0.004	0.001	0.002	0.003	0.010	0.020	0.373
*HET **	0.000	0.004	0.003	0.019	0.102	0.023	0.130	0.000	0.170	0.025	0.125	0.000
*NOR **	0.001	0.931	0.249	0.034	0.000	0.307	0.063	0.000	0.000	0.675	0.021	0.000
*FF **	0.003	0.182	0.182	0.006	0.091	0.626	0.626	0.058	0.436	0.160	0.445	0.755

*Notes: * indicates the p-value of the associated statistic; the sample starts in 1994:1, except for Autria (1995:1), Finland (1995:1) and Luxembourg (1996:11), and ends in 2014:12; coefficients in bold are significant at 5%; t_BDM and F_PSS are, respectively, Banerjee et al. (1998) and Pesaran et al. (2001) cointegration statistics; the 5% bounds for t_BDM are -2.86 and -3.53, and for F_PSS are 3.79 and 4.85 CRADF is the Engle and Granger (1987) cointegrating regression ADF statistic computed on Eq. (1), and its 5% critical value (with 200 observations) is -3.78; Trace and Max.Eig. are, respectively, Johansen's trace and maximum eigenvalue tests and indicate the number of cointegrating relationships in the system (p_t, c_t, er_t); SC(i), HET, NOR and FF are p-values of the serial correlation LM statistic with i lags, SC(i), White's heteroskedasticity test, HET, Jarque-Bera normality test, NOR, and Ramsey's functional form test, FF.*

Table 5 – Long-run elasticities in the symmetric model and unitary elasticity tests.

	AT	*BE*	*DE*	*ES*	*FI*	*FR*	*GR*	*IE*	*IT*	*LU*	*NL*	*PT*
Long-run elasticities												
c	0.48	0.62	0.65	0.64	0.66	0.73	0.69	0.56	0.61	0.60	0.58	0.65
er	0.35	0.52	0.54	0.53	0.77	0.72	0.88	0.60	0.58	0.51	0.59	1.00
Unitary long-run elasticity tests, H_0*:* $\beta_i = 1$												
c	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
er	0.000	0.000	0.000	0.000	0.033	0.010	0.275	0.000	0.000	0.000	0.000	0.996

Notes: elasticities are calculated as in Eq. (4) using the estimates in Table 4; the coefficients are all significant at 5% (significance levels have been calculated with the Delta method: for an overview see Davidson and MacKinnon, 2004).

Table 6 – Dynamic nonlinear estimation. Dependent variable: Δ(r).

	AT	BE	DE	ES	FI	FR	GR	IE	IT	LU	NL	PT
constant	**-0.32**	**-0.53**	**-0.49**	**-0.42**	**-0.49**	**-0.44**	**-0.46**	**-0.54**	**-0.40**	**-0.46**	**-0.32**	**-0.26**
r_{-1}	**-0.23**	**-0.33**	**-0.28**	**-0.26**	**-0.30**	**-0.24**	**-0.25**	**-0.37**	**-0.26**	**-0.37**	**-0.21**	**-0.17**
er^{+}_{-1}	**0.22**	**0.28**	**0.27**	**0.19**	**0.31**	**0.22**	**0.22**	**0.35**	**0.18**	**0.30**	**0.16**	**0.21**
er^{-}_{-1}	-0.05	0.05	0.00	0.05	**0.13**	**0.10**	**0.17**	**0.13**	**0.10**	0.05	**0.09**	**0.15**
c^{+}_{-1}	**0.08**	**0.17**	**0.12**	**0.12**	**0.16**	**0.14**	**0.14**	**0.18**	**0.12**	**0.18**	**0.11**	**0.11**
c^{-}_{-1}	**0.16**	**0.23**	**0.19**	**0.15**	**0.20**	**0.17**	**0.14**	**0.24**	**0.14**	**0.24**	**0.13**	**0.13**
$\Delta(r_{-1})$	**0.22**	0.00	0.10	**0.26**	**0.16**	**0.16**	**0.12**	**0.41**	**0.23**	**0.17**	0.11	**0.22**
$\Delta(er^{+})$	0.03	0.16	0.15	0.03	**0.59**	0.11	**0.37**	-0.03	0.05	0.18	0.06	-0.05
$\Delta(er^{-})$	**0.51**	0.35	**0.50**	**0.50**	0.46	**0.65**	**0.49**	0.33	**0.41**	**0.53**	**0.74**	0.16
$\Delta(er^{+}_{-1})$	0.21	**0.69**	**0.72**	0.30	0.31	**0.59**	0.28	-0.30	**0.39**	0.34	0.24	0.36
$\Delta(er^{-}_{-1})$	0.03	0.27	-0.18	0.08	-0.04	-0.12	0.00	-0.20	0.00	0.12	-0.14	0.12
$\Delta(c^{+})$	**0.38**	**0.60**	**0.55**	**0.36**	**0.35**	**0.51**	**0.51**	0.07	**0.38**	**0.48**	**0.53**	**0.36**
$\Delta(c^{-})$	**0.47**	**0.49**	**0.65**	**0.49**	**0.55**	**0.52**	**0.49**	0.11	**0.40**	**0.57**	**0.58**	**0.30**
$\Delta(c^{+}_{-1})$	0.04	0.07	0.07	0.04	0.14	0.10	**0.15**	0.00	0.10	0.08	0.05	0.07
$\Delta(c^{-}_{-1})$	0.07	**0.19**	0.09	0.02	0.02	0.10	0.07	-0.04	0.06	0.02	-0.09	0.10
Adj. R^2	0.64	0.62	0.65	0.72	0.53	0.69	0.67	0.50	0.73	0.68	0.64	0.39
t_BDM	**-5.98**	**-6.27**	**-5.99**	**-6.17**	**-8.11**	**-5.55**	**-5.54**	**-6.31**	**-5.70**	**-7.96**	**-5.00**	-2.92
F_PSS	**8.22**	**8.04**	**8.23**	**8.20**	**14.00**	**6.40**	**6.97**	**9.14**	**8.81**	**13.26**	**5.27**	3.03
*SC(4) **	0.55	0.41	0.28	0.03	0.00	0.23	0.31	0.02	0.20	0.26	0.08	0.78
*SC(12) **	0.39	0.20	0.03	0.01	0.00	0.02	0.02	0.17	0.08	0.02	0.01	0.41
*HET **	0.00	0.00	0.01	0.08	0.55	0.00	0.01	0.00	0.00	0.00	0.00	0.00
*NOR **	0.19	0.82	0.03	0.31	0.00	0.27	0.02	0.00	0.00	0.01	0.02	0.00
*FF **	0.00	0.85	0.03	0.01	0.35	0.16	0.05	0.31	0.03	0.09	0.30	0.73
Long-run elasticities												
c^{+}	**0.36**	**0.51**	**0.43**	**0.47**	**0.51**	**0.59**	**0.55**	**0.49**	**0.48**	**0.48**	**0.53**	**0.66**
c^{-}	**0.70**	**0.70**	**0.68**	**0.59**	**0.66**	**0.71**	**0.56**	**0.65**	**0.53**	**0.66**	**0.62**	**0.79**
er^{+}	**0.96**	**0.87**	**0.97**	**0.75**	**1.04**	**0.93**	**0.86**	**0.95**	**0.70**	**0.80**	**0.75**	**1.24**
er^{-}	-0.21	0.15	-0.01	0.19	**0.43**	**0.42**	**0.69**	**0.36**	**0.38**	0.14	**0.40**	**0.87**
Long-run asymmetry *												
c	0.000	0.001	0.000	0.004	0.010	0.035	0.862	0.000	0.091	0.000	0.149	0.299
er	0.000	0.000	0.000	0.000	0.002	0.008	0.310	0.000	0.008	0.000	0.141	0.354
Short-run asymmetry *												
c	0.458	0.995	0.491	0.249	0.665	0.950	0.427	0.981	0.941	0.853	0.459	0.863
er	0.416	0.677	0.286	0.470	0.361	0.668	0.646	0.361	0.937	0.795	0.536	0.963

*Notes: * indicates the p-value of the associated statistic; the sample starts in 1994:1, except for Autria (1995:1), Finland (1995:1) and Luxembourg (1996:11), and ends in 2014:12; coefficients in bold are significant at 5% (significance levels for long-run elasticities and long-run asymmetry tests have been calculated with the Delta method: for an overview see Davidson and MacKinnon, 2004); t_BDM and F_PSS are, respectively, Banerjee et al. (1998) and Pesaran et al. (2001) cointegration statistics; SC(i), HET, NOR and FF are p-values of the serial correlation LM statistic with i lags, SC(i), White's heteroskedasticity test, HET, Jarque-Bera normality test, NOR, and Ramsey's functional form test, FF; long-run elasticities are obtained as indicated in Eq. (6); long- and short-run asymmetry are, respectively, the p-values of the test of equality of the long term coefficients and of the sum of the short-run coefficients for positive and negative variations of the corresponding explanatory variable.*

Table 7 – Robustness check on the dynamic nonlinear estimation. Long term elasticities.

		Volatility	*Seasonality*	*Demand 1*	*Demand 2*	*Supply*	*Stocks*
AT	c^+	**0.36**	**0.36**	**0.33**	**0.35**	**0.34**	**0.31**
	c^-	**0.63**	**0.70**	**0.71**	**0.70**	**0.70**	**0.73**
	er^+	**0.87**	**0.95**	**0.98**	**0.96**	**0.97**	**1.10**
	er^-	-0.05	-0.19	-0.33	-0.22	-0.25	-0.35
BE	c^+	**0.51**	**0.51**	**0.49**	**0.49**	**0.49**	**0.50**
	c^-	**0.68**	**0.70**	**0.71**	**0.71**	**0.70**	**0.70**
	er^+	**0.85**	**0.87**	**0.89**	**0.89**	**0.90**	**0.88**
	er^-	0.20	0.16	0.11	0.08	0.12	0.15
DE	c^+	**0.43**	**0.43**	**0.42**	**0.42**	**0.40**	**0.45**
	c^-	**0.65**	**0.68**	**0.69**	**0.68**	**0.65**	**0.69**
	er^+	**0.94**	**0.97**	**0.99**	**0.99**	**0.95**	**1.04**
	er^-	0.06	-0.01	-0.06	-0.04	-0.05	0.16
ES	c^+	**0.47**	**0.47**	**0.45**	**0.48**	**0.48**	**0.46**
	c^-	**0.58**	**0.60**	**0.59**	**0.58**	**0.58**	**0.59**
	er^+	**0.74**	**0.76**	**0.79**	**0.73**	**0.74**	**0.77**
	er^-	0.21	0.20	0.18	0.25	0.23	0.17
FI	c^+	**0.52**	**0.51**	**0.50**	**0.54**	**0.57**	**0.55**
	c^-	**0.65**	**0.66**	**0.64**	**0.66**	**0.66**	**0.69**
	er^+	**1.02**	**1.03**	**1.01**	**1.03**	**0.97**	**1.10**
	er^-	**0.46**	**0.43**	**0.41**	**0.52**	**0.59**	**0.56**
FR	c^+	**0.59**	**0.60**	**0.56**	**0.60**	**0.60**	**0.57**
	c^-	**0.72**	**0.72**	**0.70**	**0.70**	**0.71**	**0.63**
	er^+	**0.94**	**0.94**	**0.91**	**0.90**	**0.93**	**0.80**
	er^-	**0.40**	**0.44**	**0.30**	**0.47**	**0.45**	**0.63**
GR	c^+	**0.55**	**0.56**	**0.50**	**0.45**	**0.43**	**0.52**
	c^-	**0.51**	**0.57**	**0.56**	**0.62**	**0.61**	**0.62**
	er^+	**0.81**	**0.86**	**0.93**	**1.09**	**1.05**	**1.03**
	er^-	**0.79**	**0.70**	**0.51**	0.30	0.22	**0.57**
IE	c^+	**0.49**	**0.49**	**0.46**	**0.47**	**0.50**	**0.47**
	c^-	**0.59**	**0.64**	**0.65**	**0.65**	**0.62**	**0.66**
	er^+	**0.88**	**0.93**	**1.01**	**0.99**	**0.95**	**1.00**
	er^-	**0.48**	**0.35**	**0.32**	**0.33**	**0.45**	**0.32**
IT	c^+	**0.48**	**0.49**	**0.47**	**0.41**	**0.42**	**0.45**
	c^-	**0.52**	**0.54**	**0.53**	**0.59**	**0.53**	**0.55**
	er^+	**0.68**	**0.68**	**0.73**	**0.90**	**0.77**	**0.77**
	er^-	**0.41**	**0.39**	**0.40**	0.12	0.23	**0.30**
LU	c^+	**0.48**	**0.48**	**0.58**	**0.45**	**0.48**	**0.48**
	c^-	**0.65**	**0.67**	**0.71**	**0.67**	**0.66**	**0.67**
	er^+	**0.79**	**0.80**	**0.91**	**0.84**	**0.81**	**0.82**
	er^-	0.18	0.14	**0.46**	0.04	0.14	0.13
NL	c^+	**0.53**	**0.53**	**0.50**	**0.51**	**0.43**	**0.53**
	c^-	**0.57**	**0.64**	**0.59**	**0.61**	**0.60**	**0.62**
	er^+	**0.70**	**0.77**	**0.70**	**0.70**	**0.77**	**0.74**
	er^-	**0.51**	**0.39**	**0.35**	0.30	0.20	**0.41**
PT	c^+	**0.66**	**0.67**	**0.65**	**0.59**	**0.67**	**0.66**
	c^-	**0.88**	**0.80**	**0.80**	**0.83**	**0.76**	**0.76**
	er^+	**1.35**	**1.24**	**1.30**	**1.43**	**1.26**	**1.23**
	er^-	**0.67**	**0.88**	**0.80**	0.61	**1.00**	**0.94**

Note: Significant coefficients at 5% are reported in boldface (significance levels have been calculated with the Delta method: for an overview see Davidson and MacKinnon, 2004).

Table 8 – Long term asymmetry tests on the alternative specifications.

	Volatility		*Seasonality*		*Demand 1*		*Demand 2*		*Supply*		*Stocks*	
	c	*er*	*c*	*er*	*c*	*er*	*c*	*er*	*c*	*er*	*c*	*er*
AT	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
BE	0.009	0.004	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.002	0.002
DE	0.004	0.001	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
ES	0.025	0.002	0.002	0.000	0.002	0.000	0.041	0.009	0.038	0.011	0.030	0.007
FI	0.072	0.022	0.011	0.002	0.019	0.005	0.015	0.004	0.188	0.096	0.006	0.001
FR	0.053	0.014	0.044	0.015	0.011	0.002	0.202	0.105	0.061	0.027	0.309	0.365
GR	0.468	0.942	0.895	0.405	0.237	0.048	0.012	0.002	0.003	0.000	0.124	0.034
IE	0.028	0.013	0.000	0.000	0.000	0.000	0.000	0.000	0.043	0.015	0.003	0.003
IT	0.271	0.037	0.152	0.021	0.078	0.008	0.001	0.000	0.004	0.000	0.072	0.026
LU	0.001	0.000	0.000	0.000	0.036	0.033	0.001	0.000	0.000	0.000	0.000	0.000
NL	0.554	0.451	0.144	0.156	0.170	0.101	0.068	0.059	0.006	0.008	0.184	0.174
PT	0.088	0.112	0.313	0.386	0.205	0.213	0.136	0.128	0.540	0.595	0.522	0.580

Note: p-values of the tests of equality of long term elasticities (significance levels have been calculated with the Delta method: for an overview see Davidson and MacKinnon, 2004).

a/ working paper series

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