“ASYMMETRIC ASYMMETRIES” IN EUROZONE MARKETS GASOLINE PRICING

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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 1999:1 to 2015:12. The empirical results feature two robust patterns. Firstly, while the effects of exchange rate variations display a positive asymmetry (i.e., depreciations have a greater effect with respect to appreciations), crude price variations induce negative asymmetry (i.e., reductions in the price of crude oil have a greater effect than price rises). Secondly, the positive asymmetry to exchange rate changes is stronger in core Eurozone countries. The negative asymmetry with respect to crude oil prices confirms the results of recent empirical research and theoretical models. The different behavior between Eurozone core and periphery provides further insights in the nature of pricing asymmetries.

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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). While the “rockets and feathers” hypothesis (RFH) finds some support in the empirical literature (Kristoufek and Lunackova, 2015), the evidence 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.

This applies also to the relatively under-investigated Eurozone gasoline markets, where some studies find that gasoline and diesel prices adjust symmetrically to cost shocks (e.g., Karagiannis et al., 2014), while other find asymmetries of different sign.[1] However, these studies share with the previous literature a crucial feature, namely, they allow for asymmetric responses only in the short-run. The neglect of long-run asymmetries amounts at imposing the untested assumption that the long-run elasticities are equal for positive and negative shocks (a point already raised by Honarvar, 2009). More recent analyses of the US market 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.[2] Bagnai and Mongeau-Ospina (2015) apply this method to the Italian gasoline market, by disentangling the impact of crude price from that of the exchange rate. In their study "asymmetric asymmetries" emerge: while the asymmetry to crude price (in USD) is confirmed to be *negative*, as in Atil et al. (2014), the asymmetry to exchange rate variations is *positive*, i.e., gasoline prices in local currency responds more to a national currency depreciation (namely, an increase in the domestic cost of foreign currency), than to an appreciation.

Irrespective of its causes, the presence of "asymmetric asymmetry" suggests that the inconclusiveness of previous studies related to the Eurozone markets

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[1] A survey of the recent research is provided in Section 2.

[2] However, recent research by Chang and Serletis (2016), using a longer sample, and a GARCH-M VAR specification that accounts for the role of uncertainty, finds positive asymmetry in the US gasoline market.

might depend on another source of bias. As a matter of fact, a number of these studies express both crude oil price and gasoline price in a common currency, be it local currency or USD, before the estimation. 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.[3] Such an 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 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 1999:1 to 2015:12.[4] 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,[5] 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

[3] See e.g., Meyler (2009), Rodrigues (2009), Karagiannis et al. (2014), and Kristoufek and Lunackova (2015).

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

[5] 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).

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.

The remainder of the paper is structured as follows. Section 2 provides an overview of previous findings with a focus on European markets. Section 3 presents the data and methodology used to take into account asymmetries. Section 4 contains the estimation results, which are then discussed in Section 5. Finally, Section 6 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), Polemis and Fotis (2013), Perdiguero-García (2013; Table 1); Kristoufek and Lunackova (2015; Table 1).[6]

Table 1 summarizes the previous results on asymmetric price adjustment in major European countries obtained by Galeotti et al. (2003) and Grasso and Manera (2007). We focus on these studies because unlike most of the previous empirical literature, they include both crude oil price and the USD/EUR exchange rate in their models, thus estimating separate elasticities for the crude prices in USD and the USD/EUR exchange rate, and as such they provide results comparable to ours.[7]

[6] 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.

[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).

[Table 1 around here]

Table 1 clearly shows that the estimates obtained in these studies 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 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 depreciation 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.

3 Material and methods

3.1 Data

As shown above, the empirical studies on the European gasoline markets, which feature separate estimates of the elasticities to crude oil and the USD exchange rate, rely mostly on monthly time series. Therefore, while higher frequency data would in principle be preferable for our investigation, the results we would obtain by adopting them would not be comparable with the existing ones.[8] This would prevent us from assessing whether our econometric approach actually helps to reconcile the previous conflicting evidences. We chose therefore to use monthly data on a sample running from 1999:1 to 2015:12. Gasoline pump prices in euros (net of duties and taxes) were obtained from the European Commission (2016).[9] The crude oil price series is the Brent price expressed in dollars per barrel, its source is the U.S. Energy Information Administration (Federal Reserve Economic Data, 2016). The USD/EUR exchange rate was obtained from the Pacific Exchange Rate Service (2016); a currency depreciation for European countries consists in an increase of the exchange rate and vice versa for an appreciation.

3.2 Methodology

Our analysis is based on the auto-regressive distributed-lag (ARDL) modelling approach of Pesaran and Shin (1999).[10] The ARDL approach allows the researcher to obtain an estimate of both the long-run and the short-run relationships in a single step using a conditional ECM model, which in the symmetric case takes the following form

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

where r_t is the log of pre-tax gasoline retail price in euros, c_t is the log price of crude oil in dollars, er_t is the log of the USD/EUR exchange rate, ρ is the feedback coefficient (expected to be negative), γ_j and π_{ij} are coefficients (in particular, π_{10} and

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8 Killian (2010) points out that the observed asymmetries may depend on the frequency of observation, and advocates further study of asymmetries using monthly data.

9 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 Although price asymmetries in energy markets can be investigated also through nonlinear VAR (e.g., Rahman and Serletis, 2010), our results show that in the countries considered, with the possible exception of Netherlands, there is only one cointegrating vector among the three variables considered in the model (see Table 2 below). This justifies the use of a single equation estimator.

π_{20} are the impact elasticities of price to crude price and exchange rate, respectively).[11] The long-run elasticities can be obtained as

$$\beta_1 = -\theta_1 / \rho; \ \ \beta_2 = -\theta_2 / \rho \qquad \textbf{(2)}$$

The existence of a significant long-run relation can be tested in Eq. (1) following two approaches: a *t* test for the significance of the feedback coefficient ρ, along the lines set out by Banerjee et al. (1998), indicated as *t_BDM*; or an *F* test for the joint significance of the variables that enter Eq. (1) in levels, in the same way as Pesaran et al. (2001), indicated as *F_PSS*. An interesting feature of the ARDL approach is that these tests can be performed even in the cases in which there are no clear-cut indications about the order of integration of the variables, i.e., it is unclear whether they are generated by $I(1)$ or $I(0)$ processes. In these cases one can rely on the bounds testing method by Pesaran et al. (2001), who tabulate critical values for different combinations of $I(1)$ and $I(0)$ variables.

In Eq. (1) the dependent variable responds in a symmetric way to both increases and decreases in the explanatory variables. For instance, the impact of a variation of the crude oil price is the same both for positive and negative shocks, π_{10}, and the same applies to the long-run multiplier, β_1. Such parsimony, however, may lead to simplistic specifications, as "linear models are limited in flexibility and explanation of nonlinear behavior" (Ahn and Lee, 2012).

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. This approach leads to the following formulation

[11] While we used the same order of lags for both explanatory variables in the conditional ECM (1), this assumption can be easily relaxed in practical applications.

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

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:

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

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

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 long-run coefficients in Eq. (3) are calculated as:

$$\beta_i^+ = -\theta_i^+ / \rho; \; \beta_i^- = -\theta_i^- / \rho \qquad \textbf{(4)}$$

Short-run asymmetry exists if the following hypothesis is not rejected:

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

whereas long-run asymmetry can be assessed by testing whether the following hypothesis holds:

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

4 Results

While the bounds testing approach of Pesaran et al. (2001) allows for the presence of both $I(0)$ and $I(1)$ variables, the order of integration of the variables involved in the model should not exceed one. Unit root tests indicate that all the variables involved are $I(1)$, with the possible exception of gasoline price in Netherlands and Portugal, where the null of a unit root is rejected for the series in levels, thus suggesting a $I(0)$ data generating process.[12]

4.1 Symmetric cointegration estimates

Estimates of Eq. (1) are reported in Table 2, along with the cointegration tests, and the diagnostic tests.[13] Both the *t_BDM* and the *F_PSS* are strongly significant, thus rejecting the null of non-cointegration, except for Portugal and Spain.[14]

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[12] Results available upon request.

[13] 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 significantly influenced by the choice of lags.

[14] We also estimated a linear cointegrating relationship using the Phillips and Hansen's (1990) Fully Modified OLS (FMOLS) estimator and calculated Engle and

In order to exclude the presence of multiple cointegrating relationships among r_t, c_t and er_t, which would imply the need to use a VAR estimator, Johansen's trace and maximum eigenvalue tests have been carried out and are indicated in Table 2, respectively, as *Trace* and *Max. Eig.*. With the only exception of Netherlands, where the trace statistic indicates that there may exist two cointegrating vectors, in all the remaining countries both statistics conclude for a single cointegration vector is indicated by both statistics.

[Table 2 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. (2), are also reported in Table 2, and in all cases present the expected (positive) sign.[15]

4.2 Asymmetric cointegration estimates

Estimates of Eq. (3) are reported in Table 3 along with cointegration tests, long term elasticities, diagnostic tests, and tests for asymmetries on both short- and long-run elasticities.

[Table 3 around here]

The *t_BDM* and *F_PSS* statistics confirm the absence of cointegration for Portugal, but this hypothesis is strongly rejected in all the remaining countries, including Spain.[16] Asymmetry tests, in the lower panel of Table 3, indicate that the

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Granger's (1987) CRADF statistic: in this case, the null of no-cointegration was rejected at any meaningful level of confidence in all countries. Results available upon request.

[15] The tests for the hypothesis that the long-run elasticities are equal to one (full pass-through) consistently reject the null (results available upon request).

[16] 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

short-run response to shocks to both the crude oil and the exchange rate is symmetric at any reasonable confidence level. As for the long-run multipliers, long-run asymmetries features in the response to exchange rate shocks, or crude oil price shocks, or both, with the exceptions of France and Portugal, where long-run symmetry cannot be rejected, and mixed evidence for Netherlands, where it is rejected only at 10%. This evidence can be better represented by looking at the dynamic multipliers reported in Figure 1 and Figure 2, calculated as the responses to a 1% shock in the price of crude oil and in the exchange rate, respectively. For each country and each shock, the Figures report the dynamic multiplier to a positive shock (solid line), to a negative shock (dashed line), and the difference between the two multipliers (dotted line), along with its 5% confidence interval (shaded area). Rejection of long-run asymmetry implies that in the long run zero does not fall in this confidence interval.

[Figure 1 around here]

[Figure 2 around here]

5 Discussion

The estimates in the previous sections show that a cointegrating relation between gasoline price, crude oil price, and exchange rate exists in almost all Eurozone countries, the only exception being Portugal. In one of these countries (France) there is no sign of any asymmetric price adjustment, while in the

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 exception being Portugal where both the F and t statistics are not significant even at the 10%.

remaining ten we found that shocks to the crude oil price and/or to the exchange rate are symmetric in the short-run but asymmetric on the long-run.

As for the size of the adjustment of gasoline prices to crude oil price changes, Figure 3 summarizes the long-run effects. On average, a 1% positive shock generates an increase of almost 0.60%, while a 1% negative shock results in a reduction of nearly 0.70%. In other words, in the long-run negative asymmetry prevails: the effect of reductions in the price of crude oil is stronger than that of increases. As the Figure shows, this result is relatively uniform across all countries.

[Figure 3 around here]

Figure 4 displays the long-run effects of an exchange rate variation. The effect of an increase (i.e., a depreciation) 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% depreciation implies an average pass-through to gasoline prices of 0.94%, while a 1% appreciation leads to a 0.52% reduction, thereby showing the presence of a strong positive asymmetry in the long run. Again, and allowing for a greater heterogeneity, this is an overall consistent result across countries.

[Figure 4 around here]

An interesting pattern emerges once we split the sample into Eurozone *Core* and *Periphery* (the latter includes Greece, Ireland, Italy and Spain, while the former includes the remaining seven countries).[17] The two group's average dynamic

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[17] Portugal is not considered, given the lack of cointegration. With this qualification, our definition of Periphery therefore coincides with the widespread notion of PIIGS countries (the countries hit by the Eurozone sovereign debt crisis).

multipliers with respect to a 1% depreciation are shown in Figure 5.[18] The effects of an exchange rate depreciation differs between these groups. Core countries feature on average an almost full pass-through of a depreciation in the long run, with a short-run overshooting (the multiplier exceeds 1.1 on average from the second to the fifth month after the shock). In the Periphery the effect converges to a long-term value equal to 0.8, indicating an incomplete long-run pass-through, following a relatively more gradual adjustment pattern, with a moderate medium term overshooting. Thus, in Eurozone core countries the effect of a depreciation is on average 50% greater with respect to periphery countries in the months after the shock, with a long-run consolidated difference of about 20%.

[Figure 5 around here]

Our results confirm that estimates of price response ignoring the presence of asymmetries may lead to highly biased analysis. Consider Austria, for instance. Its long-term symmetric elasticities to crude oil price and exchange rate are, respectively, 0.68 and 0.97 (see Table 3). By allowing for asymmetries, we estimated that the long-term elasticity to crude oil price is 0.49 in response to a positive change, while it is 0.70 in response to a price reduction (Table 4); as for the long-term elasticity to exchange rate, a depreciation imparts a more than proportional pass-through (1.1) while an appreciation is passed-through by one third.

6 Conclusions

Using a recently developed econometric methodology, the NARDL model proposed by Shin et al. (2014), we investigate the presence and nature of

[18] Country specific responses were weighted with their within-group average share of gasoline consumption during the period 1999-2013. Source: http://www.eia.gov/cfapps/ipdbproject/IEDIndex3.cfm?tid=5&pid=5&aid=2 (last accessed: 2015-06-07).

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 behavior, 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 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 most of the countries considered, and is therefore confirmed to be a widespread phenomenon. Fourthly, a positive asymmetry with respect to the exchange rate, features in almost all the countries analyzed. We call "asymmetric asymmetry" the coexistence of a negative asymmetry with respect to crude price, and a positive asymmetry with respect to nominal exchange rate. Fifthly, the pass through from an exchange rate depreciation into a gasoline price increase is larger in core Eurozone countries with respect to peripheral Eurozone members.

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. We leave this extension for future research.

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References

Ahn, E.S., & Lee, J.M. (2012). The performance of nonlinearity tests on asymmetric nonlinear time series. *Journal of Economic Asymmetries*, 9(2), 11-44.

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.

Bagnai, A., & Mongeau Ospina, C.A. (2015). Long- and short-run asymmetries and hysteresis in the Italian gasoline market. *Energy Policy*, 78, 41-50.

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.

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

Chang, D., & Serletis, A. (2016). Oil, uncertainty, and gasoline prices. *Macroeconomic Dynamics,* forthcoming.

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.

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

Ellingsen, T., Friberg, R., & Hassler, J. (2006). Menu costs and asymmetric price adjustment. *CEPR Discussion Paper* No. 5749.

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

European Commission (2016) *Oil Bulletin*, https://ec.europa.eu/energy/en/statistics/weekly-oil-bulletin (last accessed: 2016-03-01).

Federal Reserve Economic Data (2016) http://research.stlouisfed.org/fred2/series/WCOILBRENTEU (last accessed: 2016-03-01).

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.

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

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).

Pacific Exchange Rate Service (2016) http://fx.sauder.ubc.ca/data.html (last accessed: 2016-03-01).

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., & 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., & 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.

Rahman, S., & Serletis, A. (2010). The asymmetric effects of oil price and monetary policy shocks: A nonlinear VAR approach. *Energy Economics*, 32, 1460-1466.

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

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.

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

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. (2009). Price dynamics in European petroleum markets. *Energy Economics*, 31(1), 99-108.

Figures

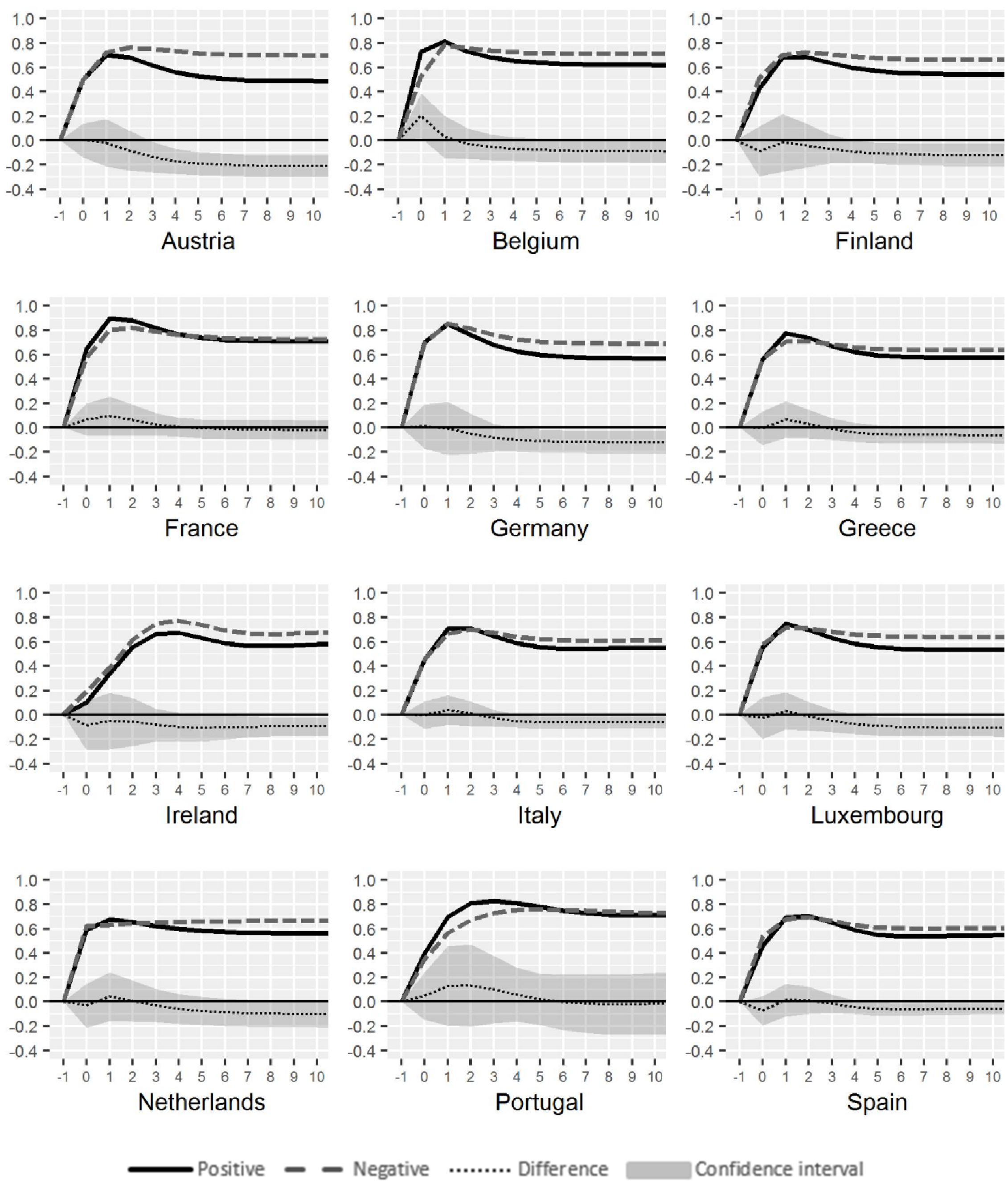

Figure 1 – Dynamic multipliers for a 1% crude oil price shock, calculated with stochastic simulation (1000 repetitions).

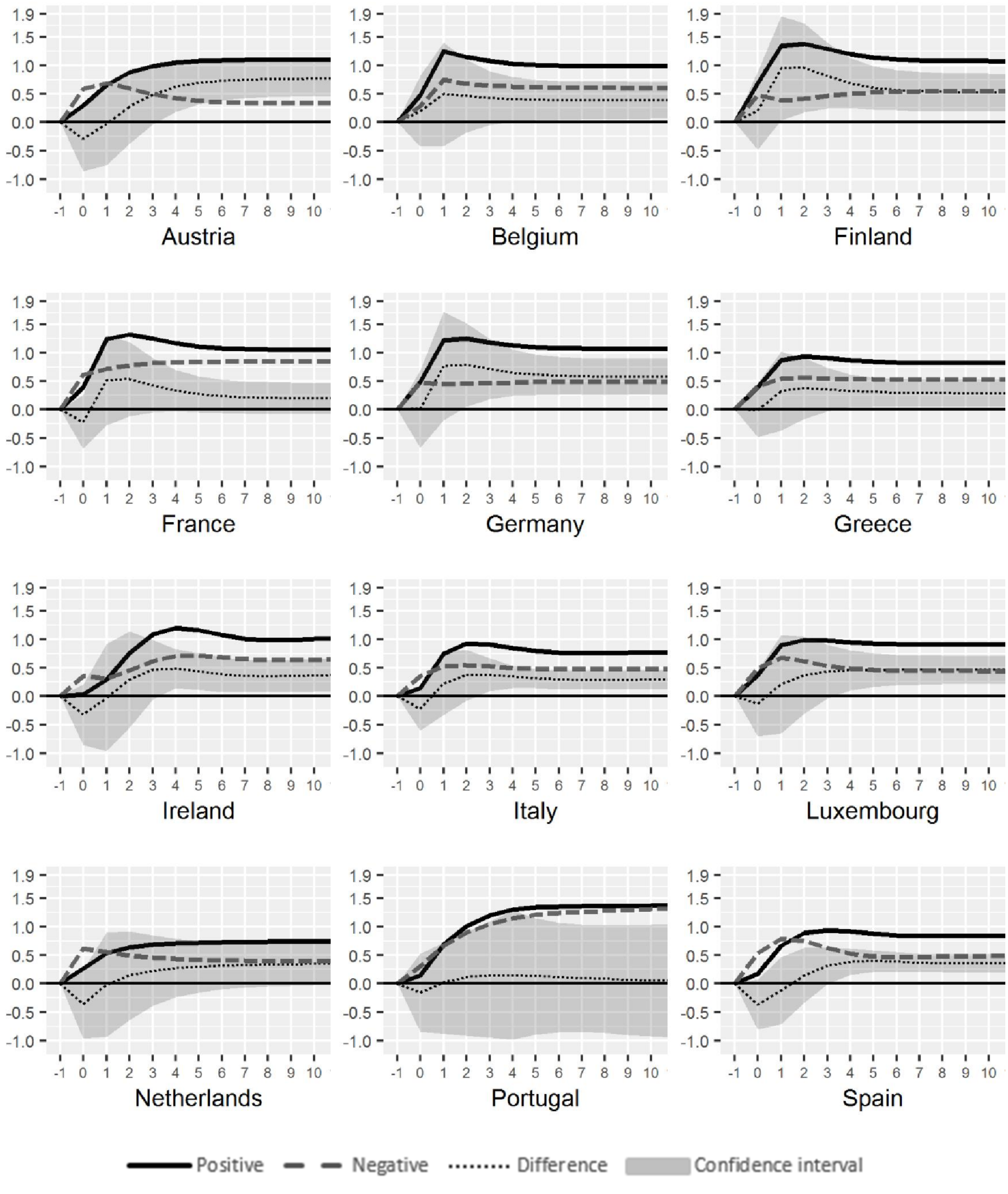

Figure 2 – Dynamic multipliers for a 1% exchange rate shock, calculated with stochastic simulation (1000 repetitions).

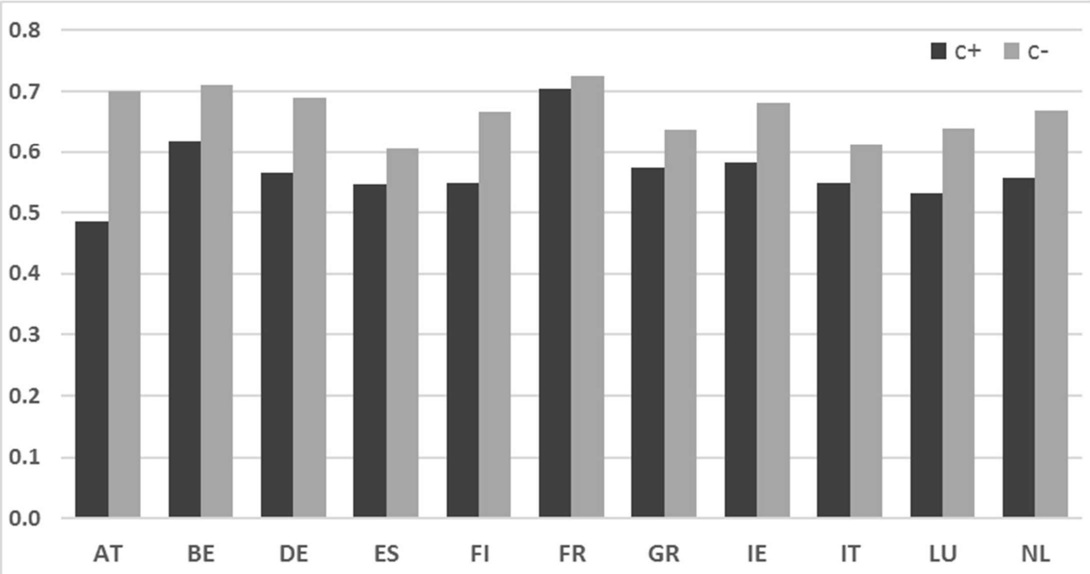

Figure 3 – Long term effect on gasoline prices of a 1% positive (c+) and negative (c−) variations of crude oil price.

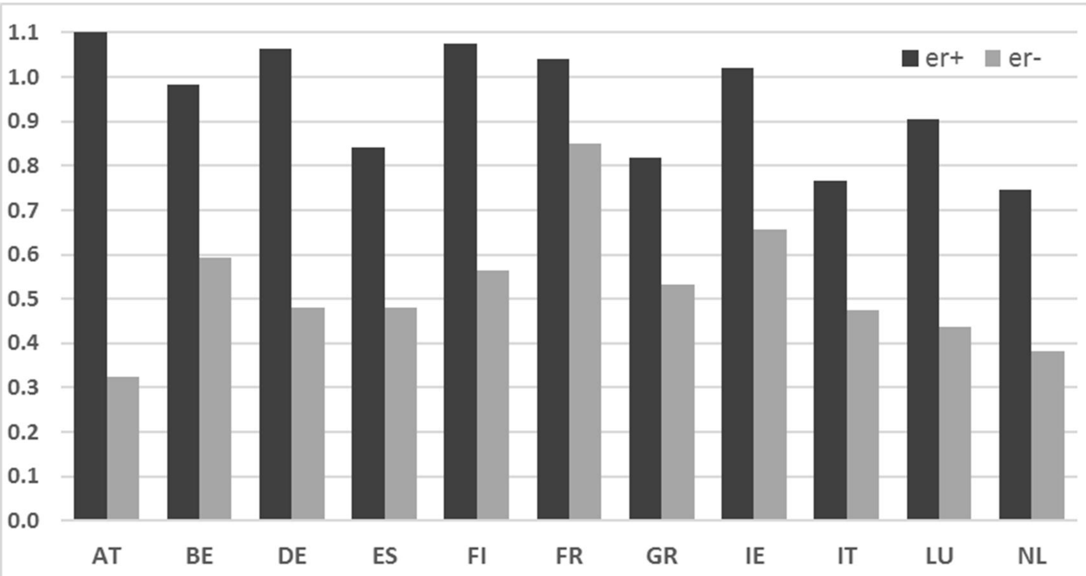

Figure 4 – Long term effect on gasoline prices of a 1% positive (er+) and negative (er−) variations of the exchange rate.

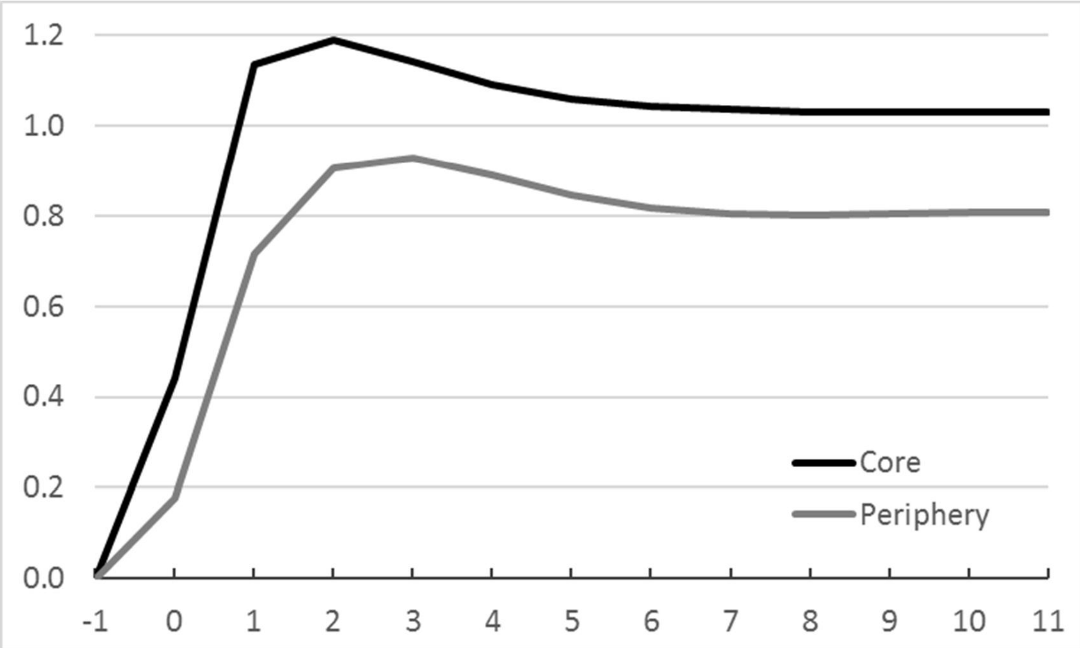

Figure 5 – Dynamic multipliers for a 1% exchange rate depreciation in "Core" and "Periphery" Euro area countries (group averages weighted with the gasoline consumption shares within each group), calculated with stochastic simulation (1000 repetitions).

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**

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); long-run elasticities are not reported by Galeotti et al. (2003) and Grasso and Manera (2007); 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 – Symmetric cointegration estimates. Dependent variable: Δ(r).

	AT	*BE*	*DE*	*ES*	*FI*	*FR*	*GR*	*IE*	*IT*	*LU*	*NL*	*PT*
constant	**-0.05**	**-0.06**	-0.04	-0.03	**-0.04**	**-0.04**	**-0.06**	**-0.09**	**-0.06**	**-0.05**	**-0.07**	-0.01
r_{-1}	**-0.22**	**-0.34**	**-0.22**	**-0.15**	**-0.25**	**-0.22**	**-0.31**	**-0.43**	**-0.28**	**-0.26**	**-0.28**	**-0.18**
er_{-1}	**0.21**	**0.30**	**0.22**	**0.12**	**0.23**	**0.22**	**0.23**	**0.41**	**0.20**	**0.21**	**0.18**	**0.24**
c_{-1}	**0.15**	**0.25**	**0.17**	**0.11**	**0.18**	**0.18**	**0.21**	**0.29**	**0.18**	**0.18**	**0.17**	**0.13**
$\Delta(r_{-1})$	**0.16**	-0.04	0.06	**0.16**	**0.14**	0.10	0.13	**0.41**	**0.18**	0.11	0.12	**0.36**
$\Delta(er)$	**0.47**	**0.40**	**0.51**	**0.35**	**0.61**	**0.49**	**0.41**	0.23	**0.26**	**0.45**	**0.47**	**0.23**
$\Delta(er_{-1})$	0.08	**0.52**	0.23	**0.24**	0.12	**0.30**	0.17	-0.25	**0.23**	0.23	0.01	0.18
$\Delta(c)$	**0.51**	**0.62**	**0.71**	**0.51**	**0.48**	**0.60**	**0.57**	**0.16**	**0.46**	**0.57**	**0.61**	**0.37**
$\Delta(c_{-1})$	**0.10**	0.15	0.07	0.07	0.09	0.11	0.06	-0.06	**0.08**	0.06	-0.03	0.05
β_1	**0.68**	**0.74**	**0.80**	**0.74**	**0.71**	**0.82**	**0.67**	**0.68**	**0.65**	**0.68**	**0.61**	**0.74**
β_2	**0.97**	**0.88**	**0.98**	**0.82**	**0.93**	**0.99**	**0.75**	**0.94**	**0.70**	**0.79**	**0.65**	**1.35**
Adj. R^2	0.72	0.66	0.68	0.75	0.54	0.74	0.75	0.56	0.77	0.68	0.69	0.54
t_BDM	**-6.65**	**-5.95**	**-4.55**	-3.07	**-5.19**	**-4.74**	**-4.88**	**-7.51**	**-6.00**	**-4.36**	**-5.30**	-2.47
F_PSS	**22.86**	**15.80**	**9.79**	4.51	**9.56**	**9.73**	**9.44**	**21.25**	**15.11**	**7.45**	**11.04**	3.96
Trace	1	1	1	0	1	1	1	1	1	1	2	1
Max-Eig.	1	1	1	0	1	1	1	1	1	1	1	1
*SC(4) **	0.772	0.451	0.558	0.028	0.024	0.081	0.049	0.974	0.258	0.168	0.387	0.311
*SC(12) **	0.011	0.117	0.010	0.003	0.000	0.071	0.014	0.993	0.023	0.007	0.057	0.768
*HET **	0.003	0.005	0.449	0.056	0.009	0.135	0.074	0.000	0.371	0.022	0.450	0.025
*NOR **	0.653	0.939	0.741	0.007	0.000	0.304	0.037	0.000	0.008	0.215	0.045	0.000
*FF **	0.082	0.004	0.407	0.022	0.483	0.500	0.813	0.296	0.908	0.714	0.414	0.846

*Notes: * indicates the p-value of the associated statistic; coefficients in bold are significant at 5%; β_1 and β_2 are the long-run elasticities to crude oil price and USD exchange rate respectively; 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; 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 3 – Dynamic nonlinear estimation. Dependent variable: Δ(r).

	AT	BE	DE	ES	FI	FR	GR	IE	IT	LU	NL	PT
const.	**-0.54**	**-0.81**	**-0.74**	**-0.73**	**-0.61**	**-0.71**	**-0.67**	**-0.84**	**-0.70**	**-0.65**	**-0.50**	**-0.34**
r_{-1}	**-0.31**	**-0.42**	**-0.37**	**-0.40**	**-0.33**	**-0.33**	**-0.38**	**-0.47**	**-0.41**	**-0.37**	**-0.30**	**-0.18**
er^{+}_{-1}	**0.34**	**0.42**	**0.40**	**0.34**	**0.36**	**0.35**	**0.31**	**0.48**	**0.31**	**0.33**	**0.22**	**0.25**
er^{-}_{-1}	**0.10**	**0.25**	**0.18**	**0.19**	**0.19**	**0.28**	**0.20**	**0.31**	**0.19**	**0.16**	**0.11**	**0.22**
c^{+}_{-1}	**0.15**	**0.26**	**0.21**	**0.22**	**0.18**	**0.23**	**0.22**	**0.28**	**0.22**	**0.20**	**0.17**	**0.13**
c^{-}_{-1}	**0.22**	**0.30**	**0.26**	**0.24**	**0.22**	**0.24**	**0.24**	**0.32**	**0.25**	**0.24**	**0.20**	**0.13**
$\Delta(r_{-1})$	**0.22**	0.01	0.13	**0.31**	**0.20**	**0.18**	**0.18**	**0.44**	**0.28**	**0.16**	**0.15**	**0.36**
$\Delta(er^{+})$	0.29	**0.48**	**0.49**	0.18	**0.69**	**0.40**	**0.39**	0.03	0.14	**0.38**	0.26	0.15
$\Delta(er^{-})$	**0.59**	0.28	0.48	**0.55**	0.48	**0.62**	**0.41**	0.36	**0.36**	**0.50**	**0.62**	0.31
$\Delta(er^{+}_{-1})$	0.05	**0.54**	0.45	0.17	**0.39**	**0.56**	0.24	-0.22	**0.32**	0.26	0.09	0.26
$\Delta(er^{-}_{-1})$	0.05	0.34	-0.10	0.11	-0.19	-0.08	0.03	-0.32	0.02	0.13	-0.08	0.08
$\Delta(c^{+})$	**0.51**	**0.74**	**0.71**	**0.46**	**0.43**	**0.65**	**0.56**	0.11	**0.46**	**0.56**	**0.60**	**0.40**
$\Delta(c^{-})$	**0.50**	**0.53**	**0.69**	**0.53**	**0.51**	**0.58**	**0.56**	**0.19**	**0.46**	**0.58**	**0.62**	**0.35**
$\Delta(c^{+}_{-1})$	0.10	0.13	0.11	0.06	0.13	0.12	0.11	-0.04	0.09	0.11	0.01	0.10
$\Delta(c^{-}_{-1})$	0.05	**0.17**	0.07	-0.05	0.03	0.07	0.01	-0.12	0.02	0.02	-0.10	0.02
Adj. R^2	0.732	0.681	0.725	0.801	0.560	0.767	0.766	0.570	0.797	0.704	0.687	0.524
t_BDM	**-6.735**	**-7.402**	**-7.182**	**-8.868**	**-7.603**	**-6.639**	**-6.080**	**-8.381**	**-7.820**	**-6.763**	**-4.944**	-2.551
F_PSS	**15.26**	**16.06**	**14.78**	**19.68**	**12.93**	**10.95**	**8.950**	**15.48**	**14.10**	**11.59**	**6.592**	2.690
*SC(4) **	0.326	0.189	0.973	0.062	0.001	0.837	0.343	0.917	0.689	0.558	0.212	0.444
SC(12)	0.084	0.071	0.028	0.006	0.000	0.221	0.044	0.985	0.184	0.027	0.067	0.810
*HET **	0.036	0.001	0.112	0.121	0.188	0.000	0.022	0.000	0.031	0.000	0.003	0.174
*NOR **	0.651	0.232	0.046	0.011	0.000	0.021	0.003	0.000	0.217	0.000	0.033	0.000
*FF **	0.016	0.326	0.225	0.030	0.614	0.692	0.140	0.193	0.048	0.476	0.436	0.245
Long-run elasticities												
c^{+}	**0.49**	**0.62**	**0.56**	**0.55**	**0.55**	**0.70**	**0.57**	**0.58**	**0.55**	**0.53**	**0.56**	**0.71**
c^{-}	**0.70**	**0.71**	**0.69**	**0.61**	**0.67**	**0.72**	**0.64**	**0.68**	**0.61**	**0.64**	**0.67**	**0.75**
er^{+}	**1.10**	**0.98**	**1.06**	**0.84**	**1.08**	**1.04**	**0.82**	**1.02**	**0.77**	**0.90**	**0.75**	**1.42**
er^{-}	**0.33**	**0.59**	**0.48**	**0.48**	**0.56**	**0.85**	**0.53**	**0.66**	**0.47**	**0.44**	**0.38**	**1.26**
Long-run asymmetry *												
c	0.000	0.100	0.027	0.031	0.034	0.660	0.117	0.043	0.024	0.013	0.072	0.749
er	0.000	0.042	0.002	0.000	0.007	0.225	0.035	0.028	0.002	0.001	0.068	0.696
Short-run asymmetry *												
c	0.657	0.196	0.713	0.651	0.887	0.252	0.317	1.000	0.366	0.478	0.530	0.470
er	0.488	0.515	0.366	0.408	0.172	0.401	0.665	0.678	0.816	0.978	0.738	0.982

*Notes: * indicates the p-value of the associated statistic; 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 (see footnote 16); 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. (4); 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.*