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Do gasoline prices respond asymmetrically in the euro zone area? Evidence from cointegrated panel data analysis

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HIGHLIGHTS

- We model gasoline price volatility for 11 euro zone countries from 2000 to 2011.
- We examine price adjustment speed in various euro zone gasoline markets.
- The results indicate price asymmetry in downstream gasoline segment.
- We state possible reasons (gasoline market characteristics) for price asymmetry.

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ABSTRACT

This paper uses the generalized method of moments (GMM) estimation to a panel data error correction model (ECM) in order to measure the asymmetries in the transmission of shocks to input prices and exchange rate onto the wholesale and retail gasoline price, respectively. For this purpose, we use an updated data set of 6369 weekly observations (January 2000 to February 2011) for 11 euro zone countries. The results indicate the existence of asymmetric responses in the retail and wholesale segment due to possible reasons (oligopolistic structure of the refining industry, existence of consumers search costs, regulatory and legal barriers).

1. Introduction

Market structure and market dynamics in oil industry across the globe are highly complicated and diversified in many aspects. To mention but a few, these are the existing differences in oil reserves, different levels of oil markets development, different political and regulatory environments, and different responses to growth challenges (Fafaliou and Polemis, 2011). Hence, to avoid generalization pitfalls and gain better policy insights, the existing oil literature often examines this industry's issues by distinguishing two broad sub-markets' categories. These are namely the upstream and the downstream oil market segment. The upstream segment comprises all the activities that have to be done to extract oil from earth whereas the downstream segment relates to activities necessary to get oil from producers to final consumers. In particular, the oil downstream includes the transportation of oil to refineries, the refinement of crude oil into final products, the transportation of these products to storage terminals, and the trading of the products produced by the wholesalers and retailers.

The oil industry in the European Union (EU) continues to be dominated by large, integrated and often multinational companies that are active in all stages of oil production (extraction, processing/refinement and retail). They can be distinguished into multinational majors (ExxonMobil, Royal Dutch Shell and BP) and minimajors—multinational companies that limit their activities to few Member States (TexacoChevron or TotalFinaElf). Other competitors, predominantly active at the national level, include Eni (Italy), Statoil, Orlen, OMV (Austria) and MOL (Hungary). The average size of companies differs between the different stages of the production process.

It is worth mentioning that in the EU retail market segment, there is a consolidation in the number of sites, which leads to

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increased average throughput and reductions in the number of sites per capita (PÖYRY, 2009). Furthermore, there is an increasing emergence of supermarkets/hypermarkets selling road fuel at their sites in some markets (most notably in the UK and France).

Due to the retail price volatility, consumers have become more reluctant to the oil companies' price setting behaviour. In other words, they tend to believe that the oil companies adjust the retail gasoline price more quickly to cost increases than to cost decreases. The phenomenon whereby prices tend to adjust differently depending on their direction is known as price asymmetry (Bettendorf et al., 2003).

This paper has two objectives. First, we explore whether asymmetric pricing can be identified in the 11 euro zone countries (Austria, Belgium, Finland, Greece, France, Germany, Ireland, Italy, Netherlands, Portugal and Spain) by utilizing ECM on the weekly price changes in order to assess current and future potential. For that purpose, we employ econometric techniques such as GMM and cointegrated panel data analysis. Second, we provide possible explanations for the existence of asymmetric responses in the retail and wholesale segment in euro zone area, which are expected to help government officials to formulate better policies in order to promote in a more effective way the functioning of the oil segments.

This paper differs from other relevant work in the field in a sense that it is the first approach focused at a comparative examination of the two downstream sub-markets of 11 euro zone countries, by using the generalized method of moments (GMM) estimation to a panel data error correction model (ECM). It is widely acknowledged that panel data can effectively deal with individual heterogeneity and can give more informative data, more variability, and less colinearity (Klevmarken, 1989; Solon, 1989; Hsiao, 1986). Besides panel data are better able to identify and measure effects that are simply not detectable in pure cross sections or pure time series data and allow us to construct and test more complicated behavioral models than purely cross-section or time-series. Lastly, with panel data, both the random and fixed effects specifications constitute improvements over the simple linear OLS model, which does not adequately account for differences in the characteristics of cross-sectional units (Baltagi, 1995; Boylaud and Nicoletti, 2001).

The remaining of the paper is organized in the following way. Section 2 explores the gasoline price volatility in the euro zone area from 2000 to 2011 and Section 3 reviews the literature. Section 4 provides the data and empirical model and the methodology employed and Section 5 reports the empirical results. Section 6 depicts the results of price asymmetry tests and Section 7 explores possible reasons of asymmetric price adjustment and draws some policy implications in favour of free competition in less competitive gasoline markets.

2. Crude oil and gasoline price volatility in the euro zone

In most European countries oil industry is still heavily regulated due to fears of problems that may arise particularly in case of an oil crisis. Globalized oil markets are not homogenous and the characteristics and competition differ even among the various sub-markets of the same oil industry (Fafaliou and Polemis, 2011). The different level of competition between the countries or between the oil segments within the same country (i.e., refining, wholesale and retail segment) may trigger price volatility, which in turns may lead to gasoline price asymmetry.

Oil prices were characterized by high volatility within the last years (see Fig. 1). The net retail gasoline prices (net pump prices) in the euro zone area have shown a tremendous increase during the last two years (38.5%), reaching the level of 1.313 Euro/litre on average (February 2011).

Spot prices and net retail prices are highly correlated and follow each other closely (see Fig. 1). More specifically, during the period running from January 2000 until February 2011 net retail price of unleaded gasoline was strongly fluctuated (560 times). 325 adjustments were upward and 235 adjustments were downward covering the 58% and 42% of the total price fluctuations, respectively. Examining the distribution of the size of the adjustments we see that they were quite small in the period 2000–2007 whereas became more volatile from 2008 onwards. The price of crude oil has followed a similar pattern. More specifically, within the same period, the price of crude oil has fluctuated 503 times; 301 (60%) adjustments were upward and 202 (40%) adjustments were downward.

3. Literature review

Within the last years there is a plethora of studies on the existence of price asymmetry in the gasoline market with controversial results[2]. In these studies there is a wide variation in the following fields (Polemis, 2012): (a) the country under examination, (b) the time frequency of the period of the data set, (c) the stage of the transmission mechanism (retail or wholesale), (d) the econometric model employed in the empirical investigation and finally, (e) the main causes of price asymmetry.

One of the most influential and contributing paper on the topic is the one of Borenstein et al. (1997). In this theoretical and empirical approach, the authors by using econometric time-series analysis argue that gasoline price asymmetry in the retail level may be triggered by the existence of tacit collusion in the market. More specifically, the authors use semimonthly prices from 1986 (March) to 1992 (to the end) in order to estimate with two stage least squares a lag adjustment model of price asymmetry. The empirical results indicate that retail prices respond more quickly to increases than to decreases in crude oil prices. Among the possible sources of this asymmetry are production/inventory adjustment lags and market power of some sellers.

Since then, there is an increasing interest from economists in assessing and explaining the main causes of gasoline price asymmetric movements and lot of research has been conducted in this field with controversial results. In order to assess the issue of asymmetric gasoline pricing, a small number of studies use daily data (Asplund et al., 2000), while other studies and the one of the authors use weekly price data (see for example Bettendorf et al., 2003; Radchenko, 2005a; Kuper and Poghosyan, 2008; Clerides, 2010). There is no doubt that a disaggregated data set (i.e., daily basis data) is preferable since it captures better the variation of the variables used in the analysis. However, for the scope of this paper we prefer to use a panel consisting of weekly price data since more disaggregated data set was limited for the 11 euro zone countries.

It is worth mentioning that the majority of the related studies focus on prices asymmetries and few of them allow for other asymmetries. However, the paper by Galeotti et al. (2003) re-examines the issue of asymmetries in the retail market of gasoline by allowing possibly asymmetric role of the exchange rate. In their stimulating paper the issue of asymmetric pricing on specific European countries (Germany, France, UK, Italy, and Spain) is examined by using an error-correction model and

[2] For an analytical review of the literature of price asymmetry see also Frey and Manera (2007) and Fotis and Polemis (2012).

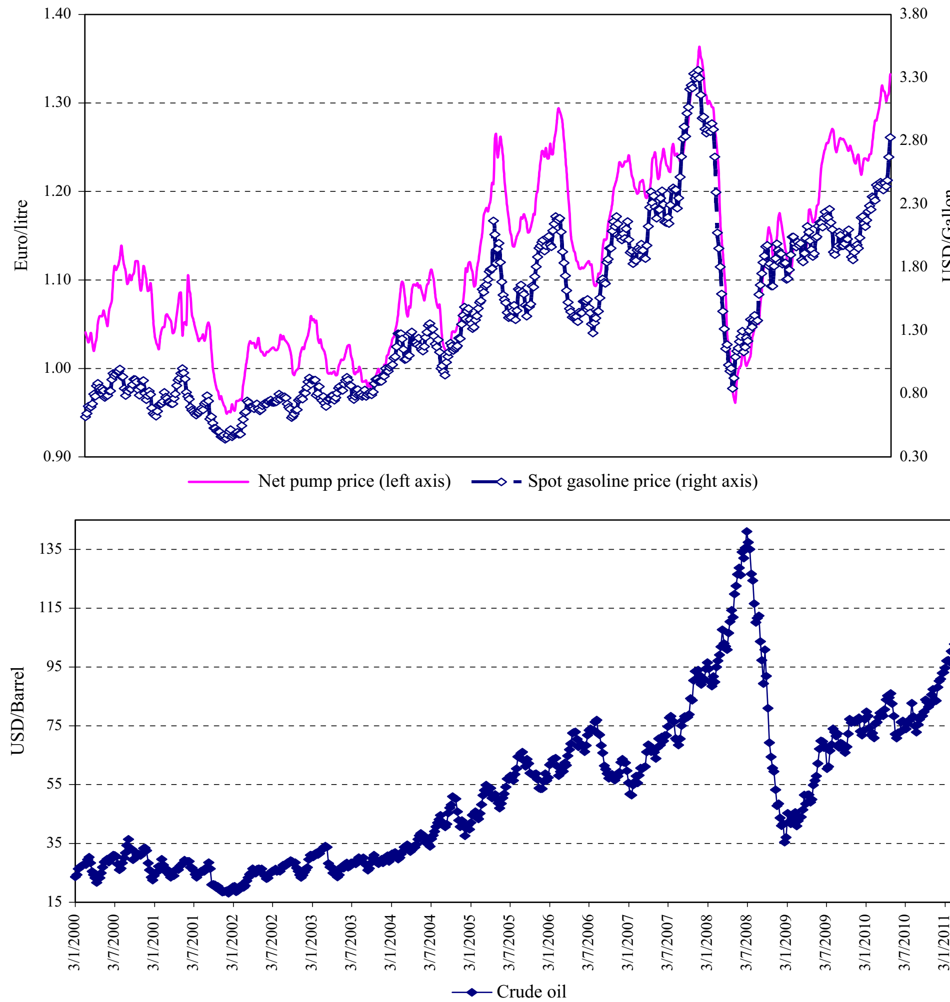

Fig. 1. Oil price evolution in the euro zone area for the period 2000–2011 (*Weekly average*)[a].
[a]Euro zone area includes Austria, Belgium, Finland, Greece, France, Germany, Ireland, Italy, Netherlands, Portugal and Spain.
Source: Adapted from Fotis and Polemis (2012).

bootstrapping techniques in order to overcome the low-power problem of conventional testing procedures. In contrast to several previous findings, the results generally point to widespread differences in both adjustment speeds and short-run responses on prices and exchange rate when input prices are volatile.

Most of these studies apply cointegration techniques in order to explore the existence and the causes of price asymmetries. More specifically, Polemis (2012), uses the error-correction model in the Greek gasoline market and argue that retail gasoline prices respond asymmetrically to cost increases and decreases both in the long and the short-run. However, at the wholesale segment, there is a symmetric response of the spot prices of gasoline towards the adjustment to the short-run responses of the exchange rate. Furthermore, Bermingham and O' Brien (2010) empirically test whether Irish and United Kingdom (UK) petrol and diesel markets are characterised by asymmetric pricing behaviour. The econometric assessment uses threshold autoregressive models (TAR) and a dataset of monthly refined oil and retail prices covering the period 1997 to mid-2009. Their study concluded that for both the Irish and UK liquid fuel markets at national levels, there is no evidence to support the hypothesis that retail prices rise faster than they fall in response to changes in oil prices (price asymmetry).

Lastly, in another econometric study (Clerides, 2010), the author by using the error-correction methodology in certain European Union countries, claim that there is a significant variation in the adjustment mechanism across countries. Fluctuations in the international price of oil are transported to local prices with some delay but evidence of asymmetric adjustment is fairly weak. Statistically significant evidence of asymmetric responses is only found in a small number of countries, while in some countries there is even (weak) evidence of asymmetry in the reverse direction: prices drop faster than they rise.

In most of the aforementioned studies the main research questions that were addressed by the authors were related with the existence of gasoline (price) asymmetry but not with its causes. The empirical literature on the possible causes of gasoline price asymmetries provides mixed results. Especially, Kuper (2012) explores the role of gasoline price history, marketing and storage cost of inventories on price asymmetry. The author examines daily gasoline prices from 1986 to 2005 and finds that until 1999, the possible source of asymmetry is due to reactions to previous gasoline price changes. After 1999, marketing and storage costs of inventories affect gasoline price asymmetrically.

Radchenko and Shapiro (2011) identify anticipated, unanticipated shocks and inventories as possible sources of gasoline price asymmetry. Their study covers the geographical area of U.S. from March 1991 to March 2010. The authors use weekly data analysis and they conclude that fluctuations in both unanticipated shocks in crude oil price and gasoline inventories cause asymmetric effects on gasoline prices. Anticipated shocks lead to significantly stronger and faster responses than unanticipated shocks.

Lewis (2011) argues that if search costs are high in a particular market then margins will be higher because fewer people will chose to search for a given price level. More specifically, according to the search cost theory, each filling station has a local monopoly which is limited by consumer search. When wholesale prices rise, the owner of each station acts to increase profit margins and quickly passes the increase to final consumers. On the contrary, when prices fall each station temporarily boosts its profit margins by slowly passing the decrease on to consumers. Only after the consumers engage in a costly and time-consuming search to find the lowest prices are the filling stations operators forced to lower prices to a competitive level. According to this theory, volatile crude oil prices create a signal-extraction problem for consumers; it encourages consumers to search less thus making gasoline filling station operators less competitive (Radchenko, 2005b).

Lastly, Angelopoulou and Gibson (2010) have stressed the significant effect of tax changes and market power on diesel prices across different regions in Greece. The authors argue that even though their empirical results provide little evidence on price asymmetries, different tax regimes and market power across different geographical regions in Greece indicates that diesel market in Greece does not behave competitively.

4. Data and methodology

4.1. Data

In this section, we present the econometric methodology. As mentioned earlier, the econometric estimation was based on pooled time-series cross-section weekly data for 11 countries covering the period from January 2000 to February 2011. Especially, the sample spans the period from January 2000 to February 2011 using an updated weekly dataset which consists of 6369 observations ($n=11$ and $T=579$) to carry out a thorough investigation of gasoline market in the euro zone area. It is clear that the sample provides long range results, which differs significantly from short term results because they both need different modeling (time series analysis vs. panel data analysis). To be more specific, the reason for using a panel data set in order to investigate possible cointegrating vectors instead of time series analysis is that residual based cointegration tests are known to have low power and are subject to normalization problems. Since economic time series are typically short, it is desirable to exploit panel data in order to draw sharper inferences (Christopoulos and Tsionas, 2003a).

4.2. Dynamic panel data specification

In order to allow for the dynamic aspects in our models we try to investigate our main research questions by using dynamic panel data techniques such as dynamic generalised method of moments (DGMM) estimators attributed to Arellano and Bond (1991) and cointegration analysis. It is worth mentioning that among the GMM, the estimators by Arellano and Bond (1991) are the most widely applied in empirical analysis (Gutierrez, 2003).

Consider the multiple linear regression model for individual $i=1,\ldots,N$ who is observed at several time periods $t=1,\ldots,T$

$$Y_{it}=\alpha+\beta_{it}X'_{it}+\gamma_i+\varepsilon_{it} \tag{1}$$

where $i=1, 2,\ldots,N$ and $t=1, 2,\ldots,T$. The N cross sectional countries are observed over T time periods. α is the intercept in the panel model, while γ_i is an individual specific effect which can be fixed or random, respectively. Y_{it} represents the dependent variable and X_{it} is a k-vector of explanatory (control) variables. Finally, ε_{it} are the disturbance terms. The vector β may be divided into sets of common, period specific and cross-section specific regressor coefficients, allowing for b coefficients to differ across periods or cross sections. In terms of Eq. (1), GMM panel estimators may reflect moments of the following type:

$$f(b)=\sum_{i-1}^{M} Z'_j\varepsilon_j(b) \tag{2}$$

where Z_i represents a $T_i\times p$ matrix of instruments for cross-section i, and

$$\varepsilon_\imath(b)=\left[\Upsilon_\imath-\gamma(X_{\imath\tau},\beta)\right] \tag{3}$$

while GMM panel estimators minimize the quadratic form:

$$g(X_{it}b)=\Phi\left(X_{it}b'\right)H\Phi(X_{it}b) \tag{4}$$

where H is a $p\times p$ weighting matrix. After estimating the coefficient vector $\hat{b}$, the coefficient covariance matrix is calculated as follows:

$$Var(\hat{b})=\left(\Phi'H\Phi\right)^{-1}\left(\Phi'H\Lambda\Phi\right)\left(\Phi'H\Phi\right)^{-1} \tag{5}$$

where Λ is an estimator of $E\left[Z'_i\varepsilon_i(b)\varepsilon_i\left(b'\right)Z_i\right]=Z'_iE\left[(\varepsilon_i(b))\varepsilon_i\left(b'\right)\right]Z_i$. In case where $g(X_{it},\ b)=X'_{it}b$, the coefficient estimator $\widehat{b}$ and its variance estimator may be specified as:

$$\hat{b}=M\left(M'_{ZX}HM_{ZX}\right)^{-1}\left(M'_{ZX}HM_{ZX}\right) \tag{6}$$

and

$$Var(\hat{b})=\left(M'_{ZX}HM_{ZX}\right)^{-1}\left(M'_{ZX}H\Lambda HM_{ZX}\right)\left(M'_{ZX}HM_{ZX}\right)^{-1} \tag{7}$$

where $M'_{ZX}HM_{ZX}=(\Sigma_{\imath-1}^{M}Z'_iX_i)$ and $M_{ZY}=(\Sigma_{\imath-1}^{M}Z'_iY_i)$. GMM estimation procedure is based upon three main steps: (i) determining the instruments Z, (ii) computing the weighting matrix H, and (iii) specifying an estimator for Λ. Efficient GMM estimators can be computed by employing dynamic panel data techniques (Vamvoukas, 2012). To introduce dynamic panel data, consider the following specification:

$$\Upsilon_{it}=\sum_{j-1}^{\rho}\pi_j\Upsilon_{it-j}+X'_{it}\beta+\varepsilon_{it} \tag{8}$$

By taking first differences in Eq. (8), we get:

$$\Delta\Upsilon_{it}=\sum_{j-1}^{\rho}\pi_j\Delta\Upsilon_{it-j}+\Delta X'_{it}\beta+\Delta\varepsilon_{it} \tag{9}$$

The individual effect γ_i has been eliminated by first differencing. Specification (9) depicts a dynamic panel model which may be estimated employing GMM methods. In GMM model the period specific instruments are related to lagged values of the dependent and predetermined variables. In the estimation procedure, along with the group of strictly exogenous variables, various instruments for each period will be used for pursuing to produce efficient GMM coefficients. Given that the disturbances are not autocorrelated, the weighting matrix H is defined as:

$$H=\left(M^{-1}\sum_{i=1}^{M}Z'_iJZ_i\right)^{-1} \tag{10}$$

where Z_i includes a group of strictly exogenous and predetermined instruments and the matrix J is employed in the two-step Arellano–Bond estimator (Arellano and Bond, 1991).

Following the specification of Polemis (2012), Kuper (2012), Hofstretter and Tovar (2010), Bettendorf et al. (2003), Kaufmann and Laskowski (2005) and Galeotti et al. (2003), various unrestricted error-correction models are used to link the relevant variables. In order to investigate the adjustment path in the relevant gasoline markets (wholesale and retail segment), we followed the error correction mechanism attributed to Engle and Granger (1987). This is a two-stage procedure in which the first step corresponds to two multiequational long-run models (wholesale and retail specification, respectively) applying GMM and the second stage corresponds to the estimation of two error-correction models (short-run models) including the long-run relations estimated in the previous step. The basic statistical assumption underlying this approach is that the variables are stationary with the first two moments of the underlying data generation process not depending on time. In fact, many time series are not well characterized as being stationary processes and so the first step is to examine the stationarity of the variables. In other words, we have to check for the presence of unit roots. If variables are non-stationary I(1) processes, then there may exist a linear combination which may well be stationary I(0) processes. If this is the case then the variables are cointegrated. Using an error correction model (ECM), short-and long-run effects can be captured by estimating the short and long-run elasticities, respectively. Therefore, the long-run equation relationships are the following:

$$SPG_t = \beta_0 + \beta_1 CR_t + \beta a_2 EXR_t + \varepsilon_t \tag{11}$$

$$NRPG_t = \beta_0 + \beta_1 SPG_t + \varepsilon_t \tag{12}$$

The above equations represent the long-run relationships in the wholesale (Eq. (11)) and retail market, respectively (Eq. (12)). The interpretation of the relevant variables comes as follows: *NRPG* measured in Euro/litre, denotes the net price of gasoline (excluding taxes and duties) while *SPG* is the Amsterdam–Rotterdam–Antwerp (ARA) 50 ppm conventional gasoline regular spot price Free on Board (FOB) quotation measured in USD/gallon[3]. *CR* is the Brent spot price for Europe measured in USD/barrel and EXR_t is the exchange rate between U.S dollar and national currencies, while finally ε_t stands for the error term. The reason for using *EXR* in the wholesale model is related with the fact that exchange rate may be a relevant source of asymmetry in non-US countries. More specifically, as stated by Galeotti et al. (2003), since crude oil is paid for in dollars whereas gasoline sells for different sums of national currencies, the exchange rate plays a significant, possibly asymmetric role.

In order to allow for possible price and exchange rate asymmetries we construct the following ECM specifications in the wholesale (Eq. (13)) and retail (Eq. (14)) market:

$$\begin{aligned}\Delta SPG_t = a_0 \sum_{i=0}^{k} a_i^{+} \Delta CRP_{t-i} + \sum_{i=0}^{l} a_i^{-} \Delta CRN_{t-i} + \sum_{i=0}^{m} b_i^{+} \Delta EXRP_{t-i} \\ + \sum_{i=0}^{n} b_i^{-} \Delta EXRN_{t-i} + \sum_{i=1}^{p} c_i \Delta SPG_{t-i} + \lambda^{+} ECMP_{t-1} \\ + \lambda^{-} ECMN_{t-1} + \varepsilon_t\end{aligned} \tag{13}$$

$$\begin{aligned}\Delta NRPG_t = a_0 + \sum_{i=0}^{k} a_i^{+} \Delta SPGP_{t-i} + \sum_{i=0}^{l} a_i^{-} \Delta SPGN_{t-i} + \sum_{i=1}^{p} b_i \Delta NRPG_{t-i} \\ + \lambda^{+} ECMP_{t-1} + \lambda^{-} ECMN_{t-1} + \varepsilon_t\end{aligned} \tag{14}$$

The Greek letter Δ is the first difference operator. In the above asymmetric ECMs, changes in the input prices (crude oil and spot prices) and fluctuations in the exchange rate are split into positive and negative changes, respectively. In other words as suggested by Galeotti et al. (2003), short-run asymmetry is captured by similarly decomposing price and exchange rate changes into $\Delta x_t^{+} = x_t - x_{t-1} > 0$ and $\Delta x_t^{-} = x_t - x_{t-1} < 0$for $x = CR$, SPG, EXR. Hence $\Delta CRP = \Delta CR$ if $\Delta CR > 0$ and 0 otherwise. $\Delta SPGP = \Delta SPG$ if $\Delta SPG > 0$ and 0 otherwise and $\Delta EXRP = \Delta EXR$ if $\Delta EXR > 0$ and 0 otherwise. The opposite holds for ΔCRN, $\Delta SPGN$ and $\Delta EXRN$. Finally *ECMP* and *ECMN* denote the one-period lagged deviation from the long-run equilibrium and account for asymmetry in the adjustment process. Similarly $ECMP = \varepsilon_t > 0$ and 0 otherwise and $ECMN = \varepsilon_t < 0$ and 0 otherwise. The orders *k, l, m, n* represent the number of lagged terms for decreases and increases in the explanatory variables, respectively, and are chosen by using the Akaike information criterion so as to make ε_t white noise.

All variables are in their natural logarithms. Energy prices are taken from the USA Department of Energy and are deflated by using the Harmonised Consumer Price Index (2005=100) provided by Eurostat. However, pre-tax gasoline retail prices are obtained from the Oil Bulletin[4]. Finally, the exchange rate between the national currencies and the US dollar is obtained from the European Central Bank and the Federal USA Bank.

5. Empirical results

In this section we present our empirical findings from the estimation of the ECMs starting from the long-run (cointegrated) equations followed by the short-run estimations. The models were estimated incorporating corrections for autocorrelated errors within cross-sectional units. In order to handle for cross-section fixed effects we used differenced data in the estimation procedure (Arellano and Bond 1991). The estimation was conducted using EVIEWS 7 software.

5.1. Stationarity and cointegration of the variables

Given the relatively short span of the cross section element ($n=11$), all the commonly used unit root tests (Augmented Dickey–Fuller, Phillips–Perron and KPSS tests) separately to each country may have low power, (Christopoulos and Tsionas, 2003a). Thus our results for the stationarity properties of the data could be seriously misguided. An increase in the power of individual unit root tests can be achieved by pooling individual time series and performing panel unit root tests (Banerjee, 1999).

To test for the existence of a unit root in a panel data setting, we have used various econometric tests (Im, Pesaran and Shin *W*-test, Fisher type tests, Levin, Lin and Chu-*t* test, and Hadri test). In all the above tests except for Handri test, the null hypothesis is that of a unit root (Table 1). The *W*-test is based on the application of the ADF test to panel data, and allows heterogeneity in both the constant and slope terms of the ADF regression (Christopoulos and Tsionas, 2003b). The ADF and PP tests are distributed as χ^2 with degrees of freedom twice the number of cross-section units (2N), under the null hypothesis. This test has the advantage over the *W*-test that its value does not depend on different lag lengths in the individual ADF regressions

[3] Due to lack of data we use from 4.4.2008 onwards, the New York (NY) spot prices of gasoline as a good proxy for the ARA 50 ppm conventional gasoline regular spot prices. It is worth mentioning, that that the correlation coefficient between the NY gasoline spot price quotation and ARA 50 ppm gasoline quotation is nearly unity (0.993). The said result is further reimbursed by unit root and cointegration analysis. As it concerns the unit root tests we have applied a series of diagnostic tests (Augmented Dickey–Fuller, Phillips-Perron, Elliot, Rothenberg and Stock Point Optimal tests) both in levels and first differences of the variables and we observe that the null-hypothesis of a unit root cannot be rejected at 5% critical value for both variables. In other words, both series are non-stationary in levels and stationary in first differences (I–1). In terms of cointegration analysis (Johansen, 1992 technique) the maximum-likelihood eigenvalue statistics indicate that the null hypothesis (no cointegration) is rejected at 5% level (for 1 cointegrating eigenvalue the trace & Max–Eigen Statistics are 0.55 and their associated *p*-values are 0.48). Therefore, the estimated likelihood ratio tests depict that there is one cointegration vector for each model.

[4] The bulletin reports weekly the average Monday's pump price without taxes and duties in each member state of the European Union.

Table 1
Panel unit root test results[b].

Variable	Levin, Lin and Chu *t*-test	Im, Pesaran and Shin *W*-test	ADF-Fisher chi-square	PP-Fisher chi-square	Hadri *z*-statistic
Levels					
EXR	−0.194	1.550	7.634	7.220	54.275*
NRPG	0.533	−3.553**	46.230*	44.529*	25.935*
SPG	−0.180	−0.502	17.115	14.468	54.174*
CR	0.801	1.741	7.033	6.607	55.634*
First differences					
Δ(*EXR*)	−80.351*	−66.762*	1,857.800*	1,859.040*	−1.786
Δ(*NRPG*)	−69.952*	−62.989*	1,521.370*	1,775.270*	−2.865
Δ(*SPG*)	−43.360*	−34.968*	972.988*	1,886.120*	−3.082
Δ(*CR*)	−84.224*	−67.890 *	1,873.170*	1,873.050*	−2.675

[b] Under the null hypothesis Hadri test assumes the absence of a unit root whereas the other unit root tests assume a unit root (Hadri, 2000). The lag lengths were selected by using Schwarz criterion with an individual intercept as an exogenous regressor.
* Significant at 1%, respectively.
** Significant at 5%, respectively.

(Christopoulos and Tsionas, 2003a). Moreover, Baltagi and Kao (2000) report that Fisher type tests such as ADF and PP are superior to the aforementioned one in terms of size-adjusted power.

Applying the relevant tests, we observe that the null-hypothesis of a unit root cannot be rejected at 5% critical value for all of the relevant variables. In other words they are integrated of order one including a deterministic component (intercept)[5].

Panel co-integration tests are used in order to draw sharper inferences since time spans of economic time series are typically short. However, when dealing with panel data the question of homogeneity arises. In order to investigate the existence of one or more cointegrated vectors we apply several tests. First, we use Pedroni's (1999) panel version of the *ADF* statistic. Second, we use Kao test (Kao, 1999) based on Engle–Granger methodology and finally we apply a Johansen test in the context of panel unit roots, which we apply to estimated residuals from long run relations (Table 2). It is worth mentioning that Pedroni's panel ADF test allows for heterogeneity in all parameters, so on *a priori* grounds we would be willing to place more emphasis on results of Pedroni's test (Christopoulos and Tsionas, 2003b).

Table 2 indicates that the null hypothesis of no cointegration is rejected at 1% level according to the employed co-integration tests. More specifically, by employing the Fisher test, (Johansen, 1992; Maddala and Wu, 1999), it is evident that there is one cointegrating vector at the 5% level for each market segment.

5.2. Long-run estimations

In this subsection, we take up estimation of the long run coefficients (asymptotic *p*-values are in parentheses)[6] given that we have established cointegration. That is, given that Eqs. (11) and (12) represent structural and not spurious long-run relations, we proceed to estimate the parameters. In order to draw significant inferences, we will need to pool the data, and use estimating techniques appropriate for panel data.

$$\begin{array}{lll} SPG = 0.77CR & +0.10EXR+\varepsilon & \\ \quad (0.00) & \quad (0.00) & \\ \quad (0.000) & \quad (0.002) & \\ \quad (15356.92) & \quad (4.72) & \end{array} \tag{15}$$

$$\begin{array}{l} NRPG = 0.46SPG+\varepsilon \\ (0.04) \\ (0.22) \\ (2.05) \end{array} \tag{16}$$

The first, second and third parenthesis under the estimated coefficients depict the *p*-values, the standard errors and the *t*—statistics of the estimated coefficients respectively. In the wholesale specification (Eq. (15)), the estimated coefficient on *CR* is significantly different from zero at the 1% significance level. The magnitude of the relevant coefficient means that a 1% increase of crude oil will lead to an increase of spot price of gasoline by 0.77%. In other words in the long run, a change in the crude oil price is nearly fully passed to the wholesale price of gasoline. Fluctuations in the exchange rate do play a significant role in the wholesale price formation. The relevant coefficient is positively related to the spot price of gasoline and its magnitude equals 0.10. In other words, an increase in the exchange rate (devaluation of national currencies against the US dollar) will tend to increase the spot price of gasoline whereas the reverse holds in case of a decrease in the exchange rate (revaluation of national currencies against the US dollar).

According to the economic theory, exchange rate does influence oil prices and this could possible raise a collinearity problem in Eq. (15)[7] That is, the more correlated the *X* variables are with each other, the bigger the standard errors become, and the less likely it is that a coefficient will be statistically significant. However, as we can see from the equation in question the reported standard errors of the estimated coefficients of exchange rate and crude oil are two low and the said coefficients are statistical different from zero at the 1% significance level.

Furthermore, the Variance Inflation Factor (VIF) analysis indicates that the variance of exchange rate or/and crude oil has not been inflated due to collinearity with the other regressor of Eq. (15)[8]. Besides, the proportion of regressors' variance is independent of the variance of the other regressors' variance in the said equation.[9]

In the retail segment (Eq. (16)) it is evident that the coefficient of price effect (*SPG*) is statistically significant and has the anticipated sign. More specifically, the price effect on the net retail price of gasoline is positive and substantial in magnitude, with the relevant coefficient bellow unity (0, 46). This means that a 10% increase (decrease) of gasoline spot price will lead to an increase (decrease) of the net retail gasoline price by 4.6%. Similarly to the wholesale specification, a change in the gasoline spot price is not fully passed through to the net retail price. The relatively smaller pass-through price mechanism compared to the wholesale segment is due to the fact that as we are moving down the oil supply chain, the price of upstream oil becomes a smaller portion of the cost of the price of oil in the next stage (wholesale). Therefore a change in the upstream oil price would generate a smaller price increase downstream. Notice also that the coefficient on *SPG* represents the combined effect of a change in the

[5] According to the three of the unit root tests this is decisively not the case for *NRPG*. However, Levin, Lin and Chu *t*-test denotes implicitly that *NRPG* is I(1).

[6] Only the statistically significant parameters are reported in the analysis.

[7] Multicollinearity may exist due to many reasons, such as, the inclusion of the same independent variable, the presence of too many dummy variables, when an independent variable is computed by other(s) independent variable(s). See, inter alia, Yousefi and Wirjanto (2004); Huang and Tseng (2010).

[8] The VIF coefficient is 0.62E-04 indicating that collinearity is not present.

[9] The estimated proportion is 123.26. See Hamilton (1994), Furthermore, the correlation coefficient between the exchange rate and crude oil is 0.10 (in absolute value), while the pair wise Granger Causality tests depict that the regressors in Eq. (15) do not Granger cause each other. The relevant *F* statistics indicate that the coefficients of the lagged values of either exchange rate or crude oil are not statistical significant (the *p*-values of the *F* statistics for the two ways of causation are 0.15 and 0.35).

Table 2
Panel cointegration tests[c].

Segment	Fisher (combined Johansen)	Kao (Engle–Granger based)	Pedroni (Engle–Granger based)
Wholesale	*Trace statistic* 191.8* [r=0] 25.35 [r > =1] *Maximum eigenvalues* 217.5* [r=0] 34.75** [r > =1] 5.306 [r > =2]	−19.556*	14.054* (ν-Statistic) −19.743* (*rho*-Statistic) −10.525* (*PP*-Statistic) 15.588* (*ADF*-Statistic)
Retail	*Trace statistic* 111.9* [r=0] 25.03 [r > =1] *Maximum eigenvalues* 114.0* [r=0] 25.03 [r > =1]	−7.775*	6.415* (ν-Statistic) −7.111* (*rho*-Statistic) −4.812* (*PP*-Statistic) −8.136* (*ADF*-Statistic)

[c] Null hypothesis implies absence of cointegration, while r denotes the number of cointegrating equations with no deterministic trend.
* Significant at1%, respectively.
** Significant at 5%, respectively.

world price of gasoline on the cost and on the mark up (Bettendorf et al., 2003).

5.3. Short-run estimations

To implement GMM we have used as instruments the exogenous variables of the models lagged L and lead LD periods. In the wholesale segment (Eq. (15)) by setting L=LD=7 the model gave acceptable results as reported below. In the retail segment (Eq. (16)) we set L=5.

From the empirical results[10] (asymptotic p-values are in parentheses), we see that in the wholesale ECM, all the coefficients have the anticipated signs (Eq. (17)).

$$\begin{array}{rrrrr}
\Delta SPG = -0.23\Delta SPG_{t-1} & -0.34\Delta SPG_{t-2} & -0.13\Delta SPG_{t-3} & -0.12\Delta SPG_{t-4} & +0.24\Delta CRP \\
(0.00) & (0.00) & (0.00) & (0.00) & (0.00) \\
(0.00) & (0.00) & (0.00) & (0.00) & (0.00) \\
(-2198.07) & (-3193.40) & (-7628.07) & (-2307.19) & (406.99) \\
+0.76\Delta CRN & +0.24\Delta EXRP & +0.16\Delta EXRN & -0.34ECM^{+}_{t-1} & -0.25ECM^{-}_{t-1} \\
(0.00) & (0.00) & (0.00) & (0.00) & (0.00) \\
(0.00) & (0.00) & (0.00) & (0.00) & (0.00) \\
(10650.09) & (110.97) & (164.49) & (-556.27) & (-765.84)
\end{array} \quad (17)$$

Spot prices register a well determined response to variations in the euro dollar exchange rate. Our point estimate suggests that a 10% increase (or devaluation) in the euro/dollar exchange rate, rendering imported crude oil more expensive in terms of euro, raises spot prices by approximately 2.5%.

Since the number of instruments is larger than that of explanatory variables, the J-test, known as the Sargan test of over-indentifying restrictions, is used to test the null of $E(\varepsilon_{it}/Z_{it})=0$. The J-test follows asymptotically the Chi-square distribution. The P-value of J-test (22.96) is 0.10 showing that the null hypothesis is not rejected leading to the conclusion that the instrumental variables are exogenous, and thus, appropriately chosen in all of the specifications.

From the retail ECM (Eq. (18)), we see that positive short-run price effect is larger (in absolute terms) than its negative counterpart[11] . This means that retail gasoline prices seem to react more to price increases and to negative gaps to the equilibrium than to price decreases and positive disequilibrium.

$$\begin{array}{rrrr}
\Delta NRPG = -0.27\Delta SPG^{+} & -0.07\Delta SPG^{-} & +0.53\Delta NRPG_{t-1} & -0.11\Delta NRPG_{t-2} \\
(0.00) & (0.04) & (0.00) & (0.00) \\
(0.07) & (0.03) & (0.09) & (0.04) \\
(-3.94) & (-2.06) & (5.70) & (-2.65) \\
-0.43ECM^{+}_{t-1} & -0.30ECM^{-}_{t-1} & & \\
(0.00) & (0.03) & & \\
(0.07) & (0.14) & & \\
(-6.10) & (-2.18) & &
\end{array} \quad (18)$$

Furthermore, the coefficients on the error correction term (positive and negative) are significantly negative. The instrument rank is greater than the number of estimated coefficients (P=10), while the reported J-statistic is 7.40 (p-value=0.11) implying that the instrument list satisfies the orthogonality conditions.

6. Testing for asymmetric responses

Having estimated the short-run responses of the output prices to input variations we focus on the gasoline price asymmetry hypothesis both in the wholesale and retail level.

Table 3 reports the calculated Wald tests testing the asymmetry hypothesis in the wholesale and retail level. More specifically, rejection of the null hypothesis H_0:$\lambda^{+}=\lambda^{-}$ implies asymmetric long-run adjustment, whereas short-run asymmetries (price and exchange rate) arise when at least one of the hypotheses H_0:$\alpha^{+}=\alpha^{-}$ or $b^{+}=b^{-}$, is rejected. By using the relevant Wald tests, we see that the hypothesis of symmetric adjustment speeds can be rejected at the wholesale and retail level as well. However, when we test for asymmetries in the retail segment, the null hypothesis (H_0:$\lambda^{+}=\lambda^{-}$) cannot be rejected suggesting the existence of symmetric adjustment speeds in the long-run.

7. Concluding remarks

The relevant empirical study uses an updated weekly dataset of 6369 observations to carry out a thorough investigation of asymmetric gasoline price responses within the euro zone area. In the specific study, we used panel data analysis and econometric techniques (GMM) in order to estimate two asymmetric ECMs at each market segment. This technique allows us to distinguish between asymmetries arising from short-lived deviations in input prices and asymmetries concerning the speed at which the gasoline price reverts to its long-run (equilibrium) level.

[10] The first, second and third parenthesis under the estimated coefficients depict the p-values, the standard errors and the t-statistics of the estimated coefficients.
[11] The first, second and third parenthesis under the estimated coefficients depict the p-values, the standard errors and the t-statistics of the estimated coefficients.

Table 3
F-tests of asymmetric responses[d].

Segmentt	$\lambda^+=\lambda^-$ (Symmetric adjustment speeds)	$\alpha^+=\alpha^-$ (price asymmetry)	$b^+=b^-$ (exchange rate asymmetry)	$\alpha^+=\alpha^-=\beta^+=\beta^-=0$ (short-run asymmetry)
Wholesale level	20.4* [0.00]	94.4* [0.00]	572.2* [0.00]	113.4* [0.00]
Retail level	0.83 [0.36]	15.66* [0.00]	–	–

[d] Null hypothesis implies absence of cointegration, while *r* denotes the number of cointegrating equations with no deterministic trend.
* Significant at 1%. The numbers in square brackets are the asymptotic *p*-values.

An in-depth analysis of the oil industry aiming at qualitative aspects of competition in euro zone area is expected to help government officials formulate better policies (that is policies which promote in a more effective way the functioning of the wholesale and retail oil segments). This paper differs from other similar work in a sense that it is the first approach focused at a comparative examination of the two downstream sub-markets of 11 euro zone countries

The empirical results favor the common perception that wholesale and retail gasoline prices respond asymmetrically to cost increases and decreases. In particular, ın the wholesale specification in the long run, a change in the crude oil price is nearly fully passed to the wholesale price of gasoline. Fluctuations in the exchange rate do also play a significant role in the wholesale price formation. In the retail segment it is evident that a change in the gasoline spot price is not fully passed through to the net retail price. The relatively smaller pass-through price mechanism compared to the wholesale segment is due to the fact that as we are moving down the oil supply chain, the price of upstream oil becomes a smaller portion of the cost of the price of oil in the next stage (wholesale).

In the short run, the wholesale ECM indicates that negative crude oil variations are generally larger than their positive counterparts, while positive and negative changes of the error correction term affect significantly the level of adjustment to long-run equilibrium. Besides, spot prices register a well determined response to variations in the euro dollar exchange rate. 10% devaluation in the euro/dollar exchange rate raises spot prices by approximately 2.5%.

Additionally, the empirical results from the retail ECM depict that retail gasoline prices seem to react more to price increases and to negative gaps to the equilibrium than to price decreases and positive disequilibrium.

Lastly, the gasoline price asymmetry hypothesis both in the wholesale and retail level indicates that symmetric adjustment speeds can be rejected at the wholesale and retail level as well. However, in the retail segment, the Wald test suggests the existence of symmetric adjustment speeds in the long-run.

In order to eliminate price asymmetries in the euro zone area, government officials should pursue policies to enhance the level of competition in the relevant markets. One suitable policy to protect consumers from welfare loses concerns the implementation of regulatory and behavioural measures as well. To be more specific, the strengthening of the role of the wholesalers and the elimination of certain barriers to entry in the oil market could provide a suitable mechanism to enhance the level of petroleum imports in the euro zone area.

Another suitable policy in order to prevent the market players from the imposition of exploitative practices (i.e., price fixing, abuse of dominant position) that hinder the level of competition is linked with a thorough investigation of mergers by the national competition authorities. Mergers in the oil sector that increase market concentration without creating economies of scale or scope may lead to anticompetitive effects and increase the market power of the incumbents. In such cases where competition is hampered, the government should develop a closely monitoring of the market in order to prevent the marketers from concerted practices.

In less deregulated countries (i.e., Greece, Portugal, Spain), the government could enhance the level of competition by a further opening of the market to new entrants such as hypermarkets or big stores and by removing certain legal or technical barriers for the establishment of new filling stations. The industry structure in other European countries (United Kingdom, France and Germany) consisted of vertically integrated companies and significant market players (hypermarkets) in the retail chain of the industry could constitute a useful paradigm to the government officials and policy makers.

Given the above contributions, our analysis could be further expanded in order to tackle a number of constraints which may be addressed in future work. Most specifically, an analysis using more disaggregated retail price data (data from branded and unbranded petrol stations) or with data on wholesale prices may reach different conclusions. Such a consideration will capture better the competitive dynamism of the oil sector and lead our research to further outcomes concerning consumer policy. However, an investigation of this matter would be very useful, but is not possible with the existing available data.

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