Gasoline price asymmetries in the Euro Zone

Polemis, Michail and Fotis, Panagiotis

12 August 2011

Gasoline price asymmetries in the euro zone[+]

Michael L. Polemis[*]
Hellenic Competition Commission and University of Piraeus

Panagiotis N. Fotis,
Hellenic Competition Commission and University of Central Greece

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 weekly observations covering the period from January 2000 to February 2011 for eleven euro zone countries (Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Netherlands, Portugal and Spain). The results favor the common perception that retail and wholesale gasoline prices respond asymmetrically to cost increases and decreases.

JEL classification L11· C51 · C33

Keywords: Generalized method of moments · panel data · asymmetries· euro zone · error correction models.

[*] **Corresponding author**: 60 Papanikoli Street, Halandri, 152 32, Greece. Tel.: +30 210 6090182; E-mail: mpolemis@epant.gr.

+ The views expressed in this paper are those of the authors.

I. Introduction

Within the last years there is a plethora of studies on the existence of price asymmetry in the gasoline market with controversial results. The majority of these studies apply time series cointegration techniques to discover the existence of price asymmetries (Galeotti, *et al.*, 2003; Grosso and Manera, 2007; Asplund, *et al.*, 2000).

This paper has two objectives. Firstly, we explore whether asymmetric pricing can be identified in the eleven euro zone countries by utilizing ECM on the weekly price changes. Despite its crucial importance due to the recent oil price hikes, this analysis has not yet been done for the euro zone area. Secondly, we employ sophisticated econometric techniques such as GMM and cointegrated panel data analysis. This article is organized as follows. Section II provides a detailed description of the empirical model and the methodology employed. Section III reports our results and Section IV concludes the article.

II. Methodology

Consider the dynamic model with invariant individual term α_i, (Arellano and Bond, 1991),

$$y_{i,t} = \beta y_{i,t-1} + \alpha_i + \varepsilon_{i,t} \quad \textbf{(1)}$$

Fist differences eliminate the invariant individual term α_i and the model becomes

$$y_{i,t} - y_{i,t-1} = \beta\left(y_{i,t-1} - y_{i,t-2}\right) + \varepsilon_{i,t} - \varepsilon_{i,t-1} \quad \textbf{(2)}$$

Since an OLS estimator is biased under the presence of autocorrelation (Wooldridge, 2002) a GMM estimator with instruments Π[1], which are not correlated with the error term and satisfy specific orthogonality conditions[2], is

[1] The over-identifying restrictions may be tested via the commonly employed J statistic (Hansen, 1982). The J statistic is distributed as χ^2 (p-k), where k is the number of estimated coefficients and p is the instrument rank. A rejection of the null hypothesis implies that the instruments are not satisfying the orthogonality conditions.

$$\arg\min_{\mu} \phi' W \phi = \hat{\mu}_{GMM} = \frac{dy'_{-1} \Pi V_N^{-1} \Pi' dy}{dy'_{-1} \Pi V_N^{-1} \Pi' dy_{-1}} \quad \textbf{(3)}$$

where W is the inverse of the covariance matrix V_N^{-1} of the ϕ_i, N is the number of cross sectional observations and $\Pi = \left(\Pi'_1 ......\Pi'_N\right)$a $N(T-2)xm$ matrix[3].

If we extend the dynamic model with additional independent variables (Hansen, 1982),

$$y_{i,t} = \beta y_{i,t-1} \gamma x'_{i,t} + \alpha_i + \varepsilon_{i,t} \quad \textbf{(4)}$$

the GMM estimator becomes

$$\hat{v}_{GMM} = \left[\left(D\overline{X}\right)' \Pi V_N^{-1} \Pi' (D\overline{X})\right]^{-1} \left(D\overline{X}\right)' \Pi V_N^{-1} \Pi' dy \quad \textbf{(5)}$$

where $D\overline{X}$ is a matrix which is composed of $(T-2)NxK$ elements of $d\overline{x}_{i,t}$. In this case the instrumental matrix Π is equal to $\Pi_i = diag\left(dy_{i,1}....dy_{i,s}, dx'_{i,1}.....dx'_{i,s+1}\right)$, $i = 1.....N, \mathrm{s} = 1....\mathrm{T-2}$.

The asymmetry in the transmission of changes in input prices to output prices can be accommodated within a dynamic model (see Equation 4). In order to allow for possible price and exchange rate asymmetries we construct the following ECM specifications:

$$\varDelta SPG_t = \sum_{i=0}^{k} a_i^+ \Delta \mathrm{CRP}_{\mathrm{t-i}} + \sum_{i=0}^{l} a_i^- \Delta \mathrm{CRN}_{\mathrm{t-i}} + \sum_{i=0}^{m} b_i^+ \Delta EXRP_{\mathrm{t-i}} + \sum_{i=0}^{n} b_i^- \Delta EXRN_{\mathrm{t-i}} + \sum_{i=1}^{p} c_i \Delta SPG_{\mathrm{t-i}} + \lambda^+ ECM_{t-1} + \lambda^- ECM_{t-1} + \varepsilon_t \quad \textbf{(6)}$$

$$\varDelta NRPG_t = \sum_{i=0}^{k} a_i^+ \Delta SPGP_{\mathrm{t-i}} + \sum_{i=0}^{l} a_i^- \Delta SPGN_{\mathrm{t-i}} + \sum_{i=1}^{p} b_i \Delta NRPG_{\mathrm{t-i}} + \lambda^+ ECM_{t-1} + \lambda^- ECM_{t-1} + \varepsilon_t \quad \textbf{(7)}$$

---

[2]

$$E\begin{bmatrix} d_{u,3}.......d_{y,i} & & \\ . & . & . \\ d_{u,i,T} & d_{y,i,T-2} & \end{bmatrix}_{mx1} = E\left(Z' u_i\right) = E(\phi_i) = 0, u_i = \alpha_i + \varepsilon_{i,t}$$

$$Z_i = diag\left\lfloor d_{y,i,1}......d_{y,s} \right\rfloor_{T-2xm}, \mathrm{s} = 1....T-2$$

$du_{i,t} = \left[du_{i,3}......du_{i,T}\right]'$ and T the periods of cross section observations.

[3] Estimation of μ_{GMM} is based on the empirical moments $\phi = E(\phi_i) = \left(\frac{1}{N}\right)\sum_{i=1}^{N} \Pi'_i du_i = \frac{1}{N} \Pi' du.$

where Δ is the first difference operator. NRPG measured in Euro/litre, denotes the net price of gasoline (excluding taxes) while SPG is the Rotterdam gasoline spot price measured in USD/litre[4]. CR is the Brent spot price for Europe measured in USD/litre. In the above ECMs, changes in the input prices and fluctuations in the exchange rate are split into positive and negative changes, respectively. In this way, short-run asymmetry is captured by similarly decomposing price and exchange rate into $\Delta x_t^{+} = x_t - x_{t-1} > 0$ and $\Delta x_t^{-} = x_t - x_{t-1} < 0$ for x= SPG, EXR. 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. Finally, the exchange rate between the national currencies and the US dollar is obtained from the European Central Bank and the Federal USA Bank.

III. Empirical Results

Applying the relevant tests (Table 1), 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].

Table 1: Panel unit root test results

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*	1857.800*	1859.040*	-1.786
Δ(NRPG)	-69.952*	-62.989*	1521.370*	1775.270*	-2.865
Δ(SPG)	-43.360*	-34.968*	972.988*	1886.120*	-3.082
Δ(CR)	-84.224*	-67.890*	1873.170*	1873.050*	-2.675

Notes: * and ** imply statistical significance at the 1% and 5% levels, respectively. Under the null hypothesis Hadri test assumes the absence of a unit root whereas the other unit root tests assume a unit root. The lag lengths were selected by using Schwarz criterion with an individual intercept as an exogenous regressor.

[4] Due to lack of data we use from 4.4.2008 onwards, the New York spot prices of gasoline as a good proxy for the European spot gasoline prices.

[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).

Table 2 presents the panel cointegration tests. It is clear that the null hypothesis of no cointegration is rejected at 1% level according to the employed cointegration 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.

Table 2: Panel cointegration tests

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* (v-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* (v-Statistic) -7.111* (rho-Statistic) -4.812* (PP-Statistic) -8.136* (ADF-Statistic)

Notes: * and ** imply statistical significance at the 1% and 5% levels, respectively. Null hypothesis implies absence of cointegration, while r denotes the number of cointegrating equations with no deterministic trend.

To implement GMM we have used as instruments the exogenous variables of the models lagged *L* and lead *LD* periods. In the wholesale segment (Equation 1) by setting *L= LD=7* the model gave acceptable results as reported below. In the retail segment (Equation 2) we set *L* = 5.

From the empirical results (asymptotic P-values are in parentheses), we see that all the coefficients have the anticipated signs (Equation 8). Negative crude oil variations are generally larger than their positive counterparts. Moreover, positive and negative changes of the error correction term affect significantly the level of adjustment to long-run equilibrium (-0.34 and -0.25 respectively).

$$\begin{aligned}\Delta SPG = &\underset{(0.00)}{-0.23\Delta SPG_{t-1}} \underset{(0.00)}{-0.34\Delta SPG_{t-2}} \underset{(0.00)}{-0.13\Delta SPG_{t-3}} \underset{(0.00)}{-0.12\Delta SPG_{t-4}} \underset{(0.00)}{+0.24\Delta CRP} \underset{(0.00)}{+0.76\Delta CRN} \underset{(0.00)}{+0.24\,\Delta EXRP} \\ &\underset{(0.00)}{+0.16\Delta EXRN} \underset{(0.00)}{-0.34ECM^{+}_{t-1}} \underset{(0.00)}{-0.25ECM^{-}_{t-1}} \qquad \textbf{(8)}\end{aligned}$$

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%. The reported J-statistic is 12.9 and the p-value is 0.12, implying that the null hypothesis cannot be rejected. The instrument rank is 18 greater than the number of estimated coefficients (10). Hence the instrumental variables are valid.

From the retail ECM (Equation 9), we see that positive short-run price effect is larger (in absolute terms) than its negative counterpart. 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. Furthermore, the coefficients on the error correction term (positive and negative) are significantly negative.

$$\underset{}{\Delta NRPG} = \underset{(0.00)}{0.27\Delta SPG^{+}} \underset{(0.04)}{-0.07\Delta SPG^{-}} \underset{(0.00)}{+0.53\Delta NRPG_{t-1}} \underset{(0.00)}{-0.11\Delta NRPG_{t-2}} \underset{(0.00)}{-0.43ECM^{+}_{t-1}} \underset{(0.03)}{-0.30ECM^{-}_{t-1}} \qquad \textbf{(9)}$$

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.

By using the relevant Wald tests (Table 3), 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_o: $\lambda^+ = \lambda^-$) cannot be rejected suggesting the existence of symmetric adjustment speeds in the long-run.

Table 3: F-tests of asymmetric responses

Segment	***$\lambda^+ = \lambda^-$ (Symmetric adjustment speeds)***	***$\alpha^+ = \alpha^-$ (price asymmetry)***	***$b^+ = b^-$ (exchange rate asymmetry)***	***$\alpha^+ = \alpha^- = \beta^+ = \beta^- = 0$ (short-run asymmetry)***
Wholesale level	20153.4* [0.00]	942068.4* [0.00]	572.2* [0.00]	11309434* [0.00]
Retail level	0.83 [0.36]	15.66* [0.00]	-	-

* and ** imply statistical significance at the 1% and 5% levels, respectively. The numbers in square brackets are the asymptotic P- values.

IV. Conclusions

The relevant empirical study uses an updated weekly dataset 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 sophisticated 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. The empirical results favor the common perception that wholesale and retail gasoline prices respond asymmetrically to cost increases and decreases. Except for the possible exercise of market power by the refineries operating in an oligopolistic way, asymmetries in the gasoline market are likely to be the outcome of other market parameters (i.e regulatory barriers, legal framework, etc).

References

Arellano, M. and Bond, S. (1991) Some tests of specication for panel data: Monte carlo evidence and an application to employment equations, *Review of Economic Studies*, **58**, 277-297.

Asplund M, Erikson R, Friberg R (2000) Price adjustments by a retail gasoline chain. Scandinavian Journal of Economics, **102**: 101-121.

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

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

Hadri, K. (2000) Testing for stationarity in heterogeneous panel data, *Econometric Journal*, **3**, 148–61.

Hansen, L. 1982. Large sample properties of generalized method of moments estimators. *Econometrica* **50**, (3): 1029-1054.

Im, K. S., Pesaran, M. H. and Shin, Y. (2003) Testing for unit roots in heterogeneous panels, *Journal of Econometrics*, **115**, 53–74.

Johansen, S., (1992) Cointegration in partial systems and the efficiency of single-equation analysis. *Journal of Econometrics,* **52,** 389– 402.

Kao, C. (1999) Spurious regression and residual-based tests for cointegration in panel data, *Journal of Econometrics*, **90**, 1–44.

Levin, A., Lin, C. F. and Chu, C. (2002) Unit root tests in panel data: asymptotic and finite-sample properties, *Journal of Econometrics*, **108**, 1–24.

Maddala, G. S. and Wu, S. (1999) A comparative study of unit root tests with panel data and a new simple test, *Oxford Bulletin of Economics and Statistics*, **61**, 631–52.

Pedroni, P. (1999) Critical Values for cointegration tests in heterogeneous panels with multiple regressors, *Oxford Bulletin of Economics and Statistics*, **61**, 653–70.

Wooldridge, J. 2002. Econometric analysis of cross section and panel data. The M.I.T Press.