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The taxation effect on gasoline price asymmetry nexus: Evidence from both sides of the Atlantic[☆]

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HIGHLIGHTS

- We examine the possible causes of gasoline price asymmetry across the globe.
- We investigate the effect of taxation on the retail gasoline price adjustments.
- There is a symmetric gasoline price response in the EU wholesale level.
- Less competitive gasoline markets exhibit price asymmetry.
- The oligopolistic structure of the gasoline markets inflates price asymmetry.

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ABSTRACT

This paper explores the degree of competition in various gasoline markets and infers possible causes of price asymmetry across the globe. For this purpose we use the Dynamic Ordinary Least Square method in order to estimate price asymmetry in twelve European countries and the United States for a sample of weekly observations which spans the period from June 1996 to August 2011. The results indicate the common perception that less competitive gasoline markets exhibit price asymmetry, while highly competitive gasoline markets follow a symmetric price adjustment path. Finally, the inclusion of taxes (VAT and excise tax) into retail gasoline prices, supports the existence of price asymmetry in many European countries.

1. Introduction

Oil prices are often characterized by high volatility (Bettendorf et al., 2003; Kapetanios and Tzavalis, 2010; Fafaliou and Polemis, 2012; Fotis and Polemis, 2012; Polemis and Fotis, 2013). Due to this 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).

Many researchers and academic scholars have investigated the existence of price asymmetry in the oil markets (Frey and Manera, 2007).[1] Borenstein et al. (1997) use semimonthly prices from 1986 to 1992 in order to estimate a lag adjustment model of price asymmetry by employing two stage least squares. The empirical results indicate that retail prices respond more quickly to increases than to decreases in crude oil prices. The authors have proposed tacit collusion[2], consumer search costs, changes in inventories and crude oil price volatility as possible sources of price asymmetries.

Radchenko and Shapiro (2011) employ Engle–Granger methodology in order to examine asymmetric price movements in the

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[1] Related empirical studies are, *inter alia*, those of Douglas (2010), Deck and Wilson (2008, 2008),Galeotti et al. (2003), Eckert (2002, 2003), Johansen (1992), Salas (2002), Godby et al. (2000), Asplund et al. (2000), Peltzman (2000), Balke et al. (1998), Reilly and Witt (1998), Duffy-Deno (1996), Shin (1994), Kirchgässner and Kübler (1992), Bacon (1991), Manning (1991), Karrenbrock (1991), Lanza (1991), Norman and Shin (1991).

[2] See also Borenstein (1991), Borenstein and Shepard (1992, 1996),Slade (1992), Garcia (2010), Valadkhani (2009), Verlinda (2008) and Eckert and West (2005).

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U.S. retail gasoline market for the period 1991 to 2010. They claim that gasoline inventories and unanticipated shocks are possible causes of price asymmetry. On the other hand, Kuper (2012) indicates gasoline price history and gasoline storage costs as possible sources that affect price adjustment process. Lewis (2011) and Hofstretter and Tovar (2010) indicate search cost and government suggested retail price (*common knowledge reference price)*, as possible explanations of price asymmetry. The latter study explores the retail gasoline market of ten cities in Colombia and postulate that gasoline prices rise more slowly when costs are higher.

Clerides (2010) uses data from several European Union (EU) countries and reports significant variation in the adjustment mechanism across countries. Bermingham and O'Brien (2010) conclude that for both the Irish and UK fuel markets, there is no evidence to support the hypothesis that retail prices rise faster than they fall in response to changes in oil prices. Contin-Pilard et al. (2009) have popularized the *Political Economy Hypothesis*.

Deltas (2008) has shown that retail gasoline prices respond faster to wholesale price increases than to equivalent wholesale price decreases. The author elaborates monthly data from 1988 to 2002 and concludes that markets with high average retail-wholesale margins experience an asymmetric response. Yang and Ye (2008) have indicated that search and learning cost fluctuations may play a crucial role in the price adjustment process. Tappata (2009) has also indicated consumer search behavior and learning cost fluctuations as possible explanations of price asymmetry

The contribution of this paper is three-fold. First, it moves beyond the existing literature in that it uses a particularly extended data set covering thirteen developed countries. Second, it overcomes the problem of sample bias in the OLS estimates of the existing literature which in turn affects the tests for symmetry restriction. This is performed by the application of the Dynamic Ordinary Least Square (DOLS) method developed by Stock and Watson (1993). Third and most importantly, the paper investigates the effect of taxation (Value added and excise tax) on the retail gasoline price adjustments in which little attention has been paid by the previous studies.

The remainder of this paper is organized in the following way. Section 2 provides the data and the methodology employed, while Section 3 depicts the empirical results. Lastly, Section 4 concludes the paper, together with some policy implications.

2. Material and methods

2.1. Theoretical background

Following the specification of, *inter alia* Borenstein et al. (1997), Galeotti et al. (2003), Bettendorf et al. (2003), Hofstretter and Tovar (2010), Kuper (2012) and Polemis (2012), various unrestricted error-correction models (ECMs) are used to link the relevant variables. In order to investigate the adjustment path in the different relevant gasoline markets, we estimate two distinct asymmetric ECMs accounting for the wholesale and retail oil market segments, respectively. It is worth mentioning that in alignment with other similar studies, we focus on the examination of price and exchange rate asymmetries appeared only in the short-run. By taking into account the previous considerations, the basic (long-run) relationships are the following:

$$\mathrm{SPG}_{r,t} = \beta_0 + \beta_1 \mathrm{CR}_{r,t} + \beta_2 \mathrm{EXR}_{c,t} + \varepsilon_t \quad (1)$$

$$\mathrm{NRPG}_{c,t} = \beta_0 + \beta_1 \mathrm{SPG}_{r,t} + \varepsilon_t \quad (2)$$

The interpretation of the relevant variables comes as follows: NRPG measured in Euro/litre for EZ-11, pounds/litre for the UK and USD/gallon for the US, denotes the net price of gasoline (excluding taxes and duties), SPG is the Rotterdam gasoline spot price measured in USD/gallon. CR is the Brent spot price for Europe measured in USD/barrel[3], EXR is the exchange rate between U.S dollar and national currencies (euro for EZ-11 countries and pound for the UK, respectively) and 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.

The sample data contains weekly time series observations covering the period from June 1996 to August 2011. It is noteworthy that the specific sample period covers the years before and after the adoption of the euro (1999 for the 10 sample countries and 2001 for Greece). Moreover, our sample includes eleven European countries together with two major global oil "players" (such as the United Kingdom and the US).

Eqs. (1) and (2) represent the long-run relationships in the wholesale (Eq. (1))[4] and retail market, respectively (Eq. (2)).[5] The aforementioned equations as well as the ECMs are estimated by using DOLS. This method gives an asymptotically efficient estimator which eliminates the feedback in the cointegrating system (Stock and Watson, 1993, 2003). It involves augmenting the cointegrating regression with lags and leads so that the resulting cointegrating equation error term is orthogonal. Moreover, DOLS increases the efficiency and reduces the small sample bias relative to the OLS estimator, while DOLS generates asymptotically efficient estimates of the regression coefficients for variables that are cointegrated (Kaufmann and Laskowski, 2005). The main reason for using this method, is that although the OLS estimate of the cointegrating vector is superconsistent, it will contain a small-sample bias and the limiting distribution is non normal with a nonzero mean (Stock, 1987). A bias in the estimate for the cointegrating vector will thus affect the cointegrating residual. It is worth mentioning that most of the empirical studies devoted to this topic use the OLS method instead of DOLS (i.e. Clerides, 2010; Faber, 2009; Valadkhani, 2009; Kuper and Poghosyan, 2008). Therefore the relevant bias in the OLS estimates of the cointegrating relations affects the tests of the symmetry restriction. In order to overcome this problem we apply the DOLS method in our empirical models.

The asymmetry in the transmission of changes in input prices to output prices can be accommodated within a dynamic model. In order to allow for possible price and exchange rate asymmetries we built and estimate the ECM specifications in the wholesale (Eq. (3)) and retail market (Eq. (4)) following the existing literature (Galeotti et al. 2003; Grasso and Manera, 2007; Contin et al. 2001; Polemis, 2012):

$$\Delta \mathrm{SPG}_{r,t} = a_0 + \sum_{i=0}^{k} a_i^+ \Delta \mathrm{CRP}_{r,_{t-i}} + \sum_{i=0}^{l} a_i^- \Delta \mathrm{CRN}_{r,_{t-i}} + \sum_{i=0}^{m} b_i^+ \Delta \mathrm{EXRP}_{c,_{t-i}} + \sum_{i=0}^{n} b_i^- \Delta \mathrm{EXRN}_{c,_{t-i}} + \sum_{i=1}^{p} c_i \Delta \mathrm{SPG}_{r,_{t-i}} + \lambda^+ \mathrm{ECMP}_{t-1} + \lambda^- \mathrm{ECMN}_{t-1} + \varepsilon_t \quad (3)$$

[3] However, for the US, we used the weekly WTI spot price as traded on the New York Mercantile Exchange (NYMEX) for delivery at Cushing, Oklahoma.

[4] The subscripts r and c denote the geographic region {i=Europe, US} and the sample country respectively {n=Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, The Netherlands, Portugal, Spain and UK}.

[5] The effect of taxation on output prices in the retail oil market segment is presented in Section 3.3.

$$\Delta \mathrm{NRPG}_{c,t} = a_0 + \sum_{i=0}^{k} a_i^+ \Delta \mathrm{SPGP}_{r,t-i} + \sum_{i=0}^{l} a_i^- \Delta \mathrm{SPGN}_{r,_{t-i}} + \sum_{i=1}^{p} b_i \Delta \mathrm{NRPG}_{c,_{t-i}} \lambda^+ \mathrm{ECMP}_{t-1} + \lambda^- \mathrm{ECMN}_{t-1} + \varepsilon_t \quad (4)$$

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 ΔCRP=ΔCR if ΔCR>0 and 0 otherwise. ΔSPGP=ΔSPG if ΔSPG>0 and 0 otherwise and ΔEXRP=ΔEXR if ΔEXR>0 and 0 otherwise. The opposite holds for ΔCRN, ΔSPGN and ΔEXRN. Finally ECMP and ECMN denote the one-period lagged deviation from the long-run equilibrium (Eqs. (1) and (2)) 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.

The sample spans the period from July 1996 to August 2011 using an updated weekly dataset of 792 observations to carry out a thorough investigation of gasoline market in certain European countries and the US.[6] All variables are in their natural logarithms. Energy prices for crude oil and spot price of gasoline are taken from the US Energy Information Administration and are deflated by the Harmonised Consumer Price Index (HCPI) provided by the Eurostat. However, retail pre-tax gasoline prices measured in real terms (deflated by the HCPI) are obtained directly from the European Oil Bulletin.[7] Finally, data on the exchange rate between the national currencies and the US dollar are obtained from the European Central Bank and the Federal US Bank.[8]

2.2. Stationarity and cointegration

Unit root inference is an important step in the analysis of data. If time series are integrated of order one (I-1), cointegration is necessary to establish that we are estimating structural and not spurious equations (Christopoulos and Tsionas, 2003). For the investigation of the order of integration we have applied a series of diagnostic tests (Augmented Dickey–Fuller, Phillips–Perron, Elliot, Rothenberg and Stock Point Optimal tests).[9] both in levels and first differences of the variables. Applying the relevant tests, we observe that the null-hypothesis of a unit root cannot be rejected at 5% critical value for all the relevant variables. In other words, all the series are non-stationary in levels and stationary in first differences (I-1).

The next step is to examine if there is a cointegrated relationship between the non-stationary variables of the models. The reason for using cointegration techniques is that nonstationary time series result to spurious regressions and hence do not allow statistical interpretation of the estimations. For this purpose, we applied the Augmented Dickey–Fuller (ADF) and the cointegrated Durbin Watson tests (CRDW) and found that at the 5% level of significance the disturbance term on each (long-run) equation is stationary and integrated I(0). This means that according to the Granger representation theorem there is one cointegrating vector which corresponds to long-run equilibrium between the nonstationary variables of each model (Engle and Granger, 1987).[10]

3. Results

3.1. Long-run estimations

In this subsection, we take up estimation of the long run coefficients given that we have established cointegration. It is worth mentioning that all of the long-run equations (Eqs. (5)–(14)) are obtained by the DOLS framework.

In the wholesale specification, the estimated coefficients on crude oil (CR) are significantly different from zero at the 1% significance level for all the countries involved. The magnitude of the relevant coefficient, which is significantly high exceeding 0.92, does not reveal a severe variation between the sample countries indicating that the crude oil is an important cost marker. In other words, in the long run, a change in the crude oil price is fully passed to the wholesale price of gasoline.

On the other hand, fluctuations in the exchange rate (EXR) do not play significant role in the wholesale price formation since the relevant coefficients for all of the sample countries are not statistically significant.[11] Eqs. (5)–(9) present the estimated coefficients in the wholesale market oil segment for the selected European countries (Germany, UK, Spain and Greece) and the US. [12]

$$\underset{}{\mathrm{SPG}_{\mathrm{Germany}}} = \underset{(0.08)}{-3.33^*} + \underset{(0.02)}{0.93^*}\mathrm{CR} + \underset{(0.06)}{0.006}\mathrm{EXR} + \varepsilon \quad (5)$$

$$\mathrm{SPG}_{\mathrm{UK}} = \underset{(0.05)}{-3.34^*} + \underset{(0.01)}{0.92^*}\mathrm{CR} - \underset{(0.07)}{0.087}\mathrm{EXR} + \varepsilon \quad (6)$$

$$\mathrm{SPG}_{\mathrm{USA}} = \underset{(0.05)}{-3.38^*} + \underset{(0.01)}{0.95^*}\mathrm{CR} + \varepsilon \quad (7)$$

$$\mathrm{SPG}_{\mathrm{Spain}} = \underset{(0.33)}{-3.37^*} + \underset{(0.02)}{0.93^*}\mathrm{CR} + \underset{(0.06)}{0.007}\mathrm{EXR} + \varepsilon \quad (8)$$

$$\mathrm{SPG}_{\mathrm{Greece}} = \underset{(0.33)}{-3.34^*} + \underset{(0.01)}{0.93^*}\mathrm{CR} + \underset{(0.05)}{0.01}\mathrm{EXR} + \varepsilon \quad (9)$$

[6] The sample for the US spans the period from December 1997 to June 2011 (n=709).

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

[8] Taking into account the fixed exchange rate for the EZ-11 countries and that of Euro/dollar provided by the European Central Bank we calculate the exchange rate national currency/dollar on each week for the period January 2002 onwards by using the following formulation: national currency/dollar=fixed exchange rate euro/dollar.

[9] See Said and Dickey (1984), Dickey and Fuller (1981), Phillips and Perron (1988) and Elliott et al. (1996).

[10] The empirical results from the unit root/cointegration tests for wholesale and retail market oil segments are available from the authors upon request.

[11] The Variance Inflation Factor (VIF) analysis indicate that the variance of exchange rate or/and crude oil has not been inflated due to collinearity with the other regressor in Eqs. (5)–(9). The same result holds for the other sample European countries and is reimbursed by the correlation coefficient and the pair wise Granger Causality tests as well.

[12] In this paper we present selected long-run estimations for some of the European countries (Germany, Spain and Greece) UK and US as well. The empirical results of the econometric estimations for all of the EU countries are available from the authors upon request.

* denotes significance at 0.01; the numbers in parenthesis are the standard errors.

In the retail segment it is evident that the spot price estimated coefficients of the variable SPG are statistically significant and have the anticipated signs. Eqs. (10)–(14) present the estimated coefficients in the retail market oil segment for the selected European countries (Germany, UK, Spain and Greece) and the US, respectively.

$$\underset{}{NRPG_{Germany}} = \underset{(0.007)}{-1.103^{*}} + \underset{(0.01)}{0.515^{*}}SPG + \varepsilon \tag{10}$$

$$NRPG_{UK} = \underset{(0.01)}{-1.424^{*}} + \underset{(0.02)}{0.873^{*}}SPG + \varepsilon \tag{11}$$

$$NRPG_{USA} = \underset{(0.004)}{0.427^{*}} + \underset{(0.008)}{0.707^{*}}SPG + \varepsilon \tag{12}$$

$$NRPG_{Spain} = \underset{(0.007)}{-1.000^{*}} + \underset{(0.01)}{0.431^{*}}SPG + \varepsilon \tag{13}$$

$$NRPG_{Greece} = \underset{(0.009)}{-0.980^{*}} + \underset{(0.02)}{0.431^{*}}EXR + \varepsilon \tag{14}$$

* denotes significance at 0.01; the numbers in parenthesis are the standard errors.

More specifically, the price effect on the net retail price of gasoline is positive and substantial in magnitude, with the relevant coefficients bellow unity. It is worth mentioning that the relevant magnitude of the spot price coefficients shows significant variation between the sample countries. More specifically, in countries such as Austria, Ireland, Portugal, The Netherlands, Greece (see Eq. (14)), Italy and Spain (see Eq. (13)) the estimated coefficient is bellow 0.5, indicating 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 upstream oil price becomes a smaller portion of the cost of the price of oil in the next stage (Polemis, 2012). Therefore a change in the upstream oil price would generate a smaller price increase downstream. On the other hand, in countries like the UK(see Eq. (11)) and the US (see Eq. (12)), the long-run response of net gasoline price to spot price variations is bigger in its magnitude estimated to 0.873 and 0.707, respectively.

3.2. Short-run estimations

The empirical results from the estimation of the two ECM's (wholesale and retail market oil segments) are reported in Table A1 in the Appendix. Each coefficient of the explanatory variables of the equations in question denotes the short-run response to the spot and retail prices. In order to select the appropriate number of lags in the ECMs, we try to minimise the Akaike Information Criterion (AIC).

In the wholesale segment, we obtain larger negative coefficients, in absolute value, than their positive counterparts for all the sample countries. This finding, which is also evident in other empirical studies (Polemis, 2012; Grasso and Manera, 2007; Contin et al., 2001), reflects the consumers' perception of the actual effects of oil price variations on gasoline price changes at least in the short-run. This means that the effects of upstream price decreases are larger than those of price increases.

Moreover, on average over the estimation period, spot prices of gasoline do not register a significant response to increases (or devaluations) in the euro dollar exchange rate. In other words, in the wholesale level, positive and negative changes of the exchange rate appear to be insignificant. This evidence suggests that refineries are generally reluctant to transfer to consumers those price increases or reductions originated from movements in exchange rates.

The coefficients of the variables $ECMP_{t-1}$ and $ECMN_{t-1}$ indicate asymmetric adjustment speeds. In other words the positive and negative ECM coefficients are associated with adjustment to the long-run equilibrium level of price from above and from bellow. From the empirical results, we see that the positive coefficients are generally larger (in their absolute terms) than the negative ones for all the sample countries indicating a positive long-run asymmetry, which is not in alignment with the Wald test results (see Table 1).

However, the magnitude of the relevant error-correction terms varies significantly between the selected countries. In countries such as the UK and Ireland, the negative error-correction term is slightly above unity, and appears to be significantly smaller in the US and Greece. The same conclusion can be reached regarding the positive error-correction term. To sum up, the variation in the magnitude of the adjustment speeds primarily between the US and the European countries (e.g. UK, Ireland, France and Austria) reveals important differences in the oil industry structure regarding the level of competition in the wholesale segment.

Finally, the estimated autoregressive coefficients, which enter the model when the lag-length is equal to one (ΔSPG_{t-1}) are statistically significant and have the anticipated positive signs for the sample countries. The opposite holds when the lag length is set to two (ΔSPG_{t-2}).

We now stress our attention into the examination of point estimates in the retail level specification. From the empirical results, we see that positive short-run spot price effect is larger than its negative counterpart in a number of countries (Austria, Finland, France, Germany, Italy, Spain, UK and US), while the reverse holds for the rest of the sample countries (Belgium, Greece, The Netherlands and Portugal). 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. From the magnitude of the relevant estimates, we see that a 10% short-run increase in spot price of gasoline (wholesale price) will increase the net retail price of gasoline within the range from 1.57% (UK) to 5.05% (Germany) respectively. This outcome is intuitively valid, since crude oil, refining costs and profit account for roughly 30–40% of retail costs, while taxes (excise taxes and VAT) and wholesale margin account for another 70–60% on average.

Regarding the speed of adjustment to the long-run equilibrium, we infer that in most countries the positive coefficients are generally larger (in their absolute terms) than the negative ones thus indicating a positive long-run asymmetry in the retail segment for selected countries (Austria, France, Greece, Ireland, The Netherlands, Spain and US). However, in countries such as the UK, and Ireland, the negative error-correction term is larger than the positive one. Finally, the estimated autoregressive coefficient when the lag-length is equal to one (ΔNRP_{t-1}) is statistically significant with the anticipated positive sign for all the sample countries but for Ireland.

Table 1 depicts the calculated Wald and *F*-statistics testing the asymmetry hypothesis in the two oil market segments. 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 long-run symmetric adjustment speeds cannot be rejected at the

Table 1
Computed Wald and *F*-tests of asymmetric responses.
Source: Authors' elaboration.

Country	$\lambda^+=\lambda^-$(symmetric adjustment speeds)	$\alpha^+=\alpha^-$ (price asymmetry)	$b^+=b^-$ (exchange rate asymmetry)	$\alpha^+=\alpha^-=\beta^+=\beta^-=0$ (short-run asymmetry)
Wholesale segment: SPG=*f*(CR, EXR)				
Austria	1.17 (0.24)	0.52 (0.60)	0.08 (0.93)	137.88* (0.00)
Belgium	1.17 (0.24)	0.55 (0.58)	0.11 (0.91)	141.07* (0.00)
Finland	1.16 (0.24)	0.52 (0.60)	0.23 (0.81)	137.04* (0.00)
France	1.17 (0.24)	0.53 (0.60)	0.13 (0.89)	138.46* (0.00)
Germany	1.20 (0.23)	0.52 (0.60)	0.23 (0.81)	142.87* (0.00)
Greece	1.29 (0.19)	0.32 (0.75)	0.08 (0.93)	167.37* (0.00)
Ireland	1.23 (0.22)	0.49 (0.63)	0.58 (0.56)	145.77* (0.00)
Italy	1.17 (0.24)	0.51 (0.61)	0.20 (0.84)	138.85* (0.00)
The Netherlands	1.16 (0.25)	0.54 (0.59)	0.58 (0.56)	137.91* (0.00)
Portugal	1.18 (0.24)	0.53 (0.60)	0.23 (0.82)	139.57* (0.00)
Spain	1.17 (0.24)	0.51 (0,.61)	0.24 (0.81)	137.08* (0.00)
United Kingdom	1.22 (0.22)	0.52 (0.60)	0.28 (0.78)	154.55* (0.00)
United States	4.30*(0.00)	0.45 (0.65)	–	–
Retail segment: NRPG=*f*(SPG)				
Austria	1.95** (0.05)	2.14* (0.03)	–	–
Belgium	2.58* (0.01)	0.30 (0.76)	–	–
Finland	0.84 (0.40)	0.10 (0.92)	–	–
France	3.05* (0.00)	5.52* (0.00)	–	–
Germany	0.56 (0.58)	1.29 (0.20)	–	–
Greece	0.55 (0.57)	1.29 (0.20)	–	–
Ireland	1.70 (0.09)	0.48 (0.63)	–	–
Italy	0.24 (0.81)	0.30 (0.77)	–	–
The Netherlands	3.70* (0.00)	0.41 (0.68)	–	–
Portugal	0.27 (0.79)	0.61 (0.54)	–	–
Spain	1.96** (0.05)	2.76* (0.01)	–	–
United Kingdom	0.14 (0.88)	1.78 (0.06)	–	–
United States	4.20* (0.00)	5.61* (0.00)	–	–

**,* denote significance at 0.05 and 0.01, respectively. The numbers in parenthesis are the asymptotic *p*-values.

wholesale level for all the European countries. However, the same finding does not apply to the US. When we test for short-run asymmetries (price and exchange rate) we reach the same outcome for all the European countries and the US since the null hypothesis (H_0: $\alpha^+=\alpha^-$ and H_0: $b^+=b^-$, respectively) cannot be rejected suggesting the existence of symmetric adjustment speeds in the short-run.

When we simultaneously test the equality of all short-run parameters of the same lags in the wholesale level, the null hypothesis (equality hypothesis) is rejected for all of the sample countries. However, we must be cautious when we perform the joint equality test, since there is a tendency to over-reject the null hypothesis of symmetry due to the low power of standard *F* statistics (Galeotti et al., 2003).

From the combined results of the Wald-tests, we conclude that in the wholesale level there is symmetric gasoline price response in the European countries, which is evident both in the short and the long run. This conclusion is in alignment with other empirical studies (Godby et al., 2000; Galeotti et al., 2003; Chen et al., 2005; Contin et al., 2001; Polemis, 2012) and runs contrary to the common perception regarding the price asymmetries that emerge in the gasoline market. Similar results can be found when testing for exchange rate asymmetry in the wholesale level. However, in the US, the hypothesis of the symmetric adjustment speeds appears to be valid only in the short-run. This finding concurs with other similar studies (see for example Kaufmann and Laskowski, 2005; Bachmeier and Griffin, 2003).

When we investigate the issue of asymmetry in the retail segment, some important remarks emerge. First, there is a wide variation in the existence of asymmetric price responses within the European sample countries. It is worth mentioning that in countries with a high degree of competition such as Germany and the UK[13], whose oil industry structure is characterized by the existence of significant market players (hypermarkets, big groceries stores, etc.) in the retail chain, the null hypothesis (symmetry) cannot be rejected at least in the long-run.

The absence of (long-run) asymmetry in the retail segment of the market is consisted with previous studies focusing on the UK (OFT, 1998; Bermingham and O'Brien, 2010). On the other hand, the long-run symmetry hypothesis is rejected in a number of countries (Austria, Belgium, France, The Netherlands, Spain and the US). This finding could be traced in other studies as well (see for example Polemis, 2012; Faber, 2009; Bettendorf et al. 2003). As it concerns the short-run perspective, the existence of price asymmetry seems to be valid in Austria, France, Spain and the US.

3.3. Taxation effect

In order to investigate the effect of taxation (VAT and excise tax) on the possible asymmetrical movements of price in the retail segment, we estimated the following ECMs by using three different dependent variables representing final price without VAT (FPRV), final price without excise tax (FPREX) and final price with both taxes (FPR) in Eqs. (15)–(17), respectively.

$$\Delta \mathrm{FPRV}_{c,t} = a_0 + \sum_{i=1}^{p} b_i \mathrm{FPRV}_{c,_{t-i}} + \sum_{i=0}^{m} c_i^+ \Delta \mathrm{SPGP}_{r,_{t-i}} + \sum_{i=0}^{n} c_i^- \Delta \mathrm{SPGN}_{r,_{t-i}} + \lambda^+ \mathrm{ECMP}_{t-1} + \lambda^- \mathrm{ECMN}_{t-1} + \varepsilon_t \quad (15)$$

[13] In the UK, the supermarkets and the hypermarkets have grown continuously and significantly over the last years, whereas their volumes have grown at the expense of the traditional road site filling stations (OFT, 1998).

$$\Delta \mathrm{FPREX}_{c,t} = a_0 + \sum_{i=1}^{p} b_i \mathrm{FPREX}_{c,t-i} + \sum_{i=0}^{m} c_i^{+} \Delta \mathrm{SPGP}_{r,t-i} + \sum_{i=0}^{n} c_i^{-} \Delta \mathrm{SPGN}_{r,t-i} + \lambda^{+} \mathrm{ECMP}_{t-1} + \lambda^{-} \mathrm{ECMN}_{t-1} + \varepsilon_t \quad (16)$$

$$\Delta \mathrm{FPR}_{c,t} = a_0 + \sum_{t=1}^{p} b_i \mathrm{FPR}_{c,t-i} + \sum_{i=0}^{m} c_i^{+} \Delta \mathrm{SPGP}_{r,t-i} \sum_{i=0}^{n} c_i^{-} \Delta \mathrm{SPN}_{r,t-i} + \lambda^{+} \mathrm{ECMP}_{t-1} + \lambda^{-} \mathrm{ECMN}_{t-1} + \varepsilon_t \quad (17)$$

Eqs. (15) and (16) are estimated based on net oil prices (excluding levies and taxes), while Eq. (17) includes the taxation effect. From the empirical results (see Table A2 of the Appendix), we argue that positive short-run spot price effect is larger than its negative counterpart in a number of countries (Austria, Belgium, Finland, France, Germany, Italy, UK), while the reverse holds for the rest of the sample countries (Greece, The Netherlands, Portugal and Spain). From the magnitude of the relevant estimates, we see that a 10% short-run increase in the spot price of gasoline (wholesale price) will increase the final gasoline (pump) price without VAT within the range from 0.1% (Portugal) to 2.04% (Germany), respectively.

Regarding the speed of adjustment to the long-run equilibrium, we see that in most cases the positive coefficients are generally larger (in their absolute terms) than the negative ones thus indicating a positive long-run asymmetry in the retail segment for selected countries (France, Germany, Ireland, Italy, The Netherlands and Portugal). Similar conclusions hold in the rest specifications.

Table 2 depicts the calculated Wald and F-statistics testing the asymmetry hypothesis in the retail segment for the three dependent variables (Eqs. (15)–(17)). More specifically, from the relevant table it is evident that in the retail segment only for some European sample countries (Austria, Finland, Germany, Italy and Spain) the null hypothesis of long-run symmetry cannot be rejected (see column 1). This means that in the long-run the final price of gasoline (with taxes and duties) follows closely the variation of spot price of gasoline (SPG) indicating a smoothly symmetric path.

On the other hand, in the rest of the sample countries (i.e. Benelux countries, Greece, Ireland, United Kingdom and Portugal) this outcome is not valid since the null hypothesis is rejected implying that in the long-run variations increases (decreases) in the spot price of gasoline are not followed by similar variations in the final retail price. The existence of short-run price asymmetry is evident only in Austria, France, Germany and Spain, whereas in the rest of the sample countries the null hypothesis of price symmetry cannot be rejected at any level of statistical significance.

Regarding the net gasoline prices (excluding VAT and excise tax) it is evident that the null hypothesis of long-run symmetry is rejected in the majority of the sample countries (Belgium, France, Ireland, Spain, United Kingdom, Greece, and The Netherlands). In other words, in the long-run the net retail price of gasoline does not follow the variation of spot price of gasoline (SPG) indicating a rather asymmetric adjustment path. This could be attributed *inter alia* to the existing level of competition in the relevant geographic markets, in tandem with the regulatory restrictions that act as an obstacle to the effective market opening. However, the existence of short-run price asymmetry is evident only in France, Germany, Spain, Austria and Greece.

All in all, this paper clearly shows that on average the retail gasoline prices react more to price increases and to negative deviations from their equilibrium path, compared to price decreases and positive deviations from this path. It is highlighted that the above results also hold before VAT or other taxes (excise taxes).

Table 2
Computed Wald tests of asymmetric responses (taxation effect).
Source: Authors' elaboration.

Country	$\lambda^{+}=\lambda^{-}$ (symmetric adjustment speeds)	$\alpha^{+}=\alpha^{-}$ (price asymmetry)
FPR=*f*(SPG)		
Austria	0.16 (0.87)	7.87** (0.00)
Belgium	2.34** (0.02)	0.40 (0.69)
Finland	0.13 (0.90)	0.28 (0.78)
France	2.47* (0.01)	5.01** (0.00)
Germany	1.00 (0.32)	2.61** (0.01)
Greece	1.22** (0.22)	1.68 (0.09)
Ireland	1.98** (0.05)	0.79 (0.43)
Italy	0.65 (0.52)	1.05 (0.30)
The Netherlands	3.50* (0.00)	0.19 (0.85)
Portugal	2.47* (0.01)	0.59 (0.55)
Spain	1.58 (0.11)	2.60* (0.01)
United Kingdom	4.94* (0.00)	0.31 (0.76)
FPRV=*f*(SPG)		
Austria	0.00 (0.99)	1.54 (0.12)
Belgium	2.01** (0.05)	0.55 (0.58)
Finland	0.12 (0.90)	0.48 (0.63)
France	1.92** (0.05)	4.58* (0.00)
Germany	1.05 (0.29)	2.60 * (0.01)
Greece	0.08 (0.93)	1.76 (0.08)
Ireland	2.08** (0.04)	0.77 (0.44)
Italy	0.66 (0.51)	0.90 (0.37)
The Netherlands	3.43* (0.00)	0.23 (0.82)
Portugal	1.58 (0.12)	0.63 (0.53)
Spain	1.94** (0.05)	2.99* (0.00)
United Kingdom	4.55* (0.00)	0.35 (0.73)
FPREX=*f*(SPG)		
Austria	1.34 (0.18)	3.13* (0.00)
Belgium	3.15* (0.00)	0.32 (0.75)
Finland	0.75 (0.46)	0.55 (0.58)
France	3.53* (0.00)	5.80* (0.00)
Germany	0.07 (0.97)	2.44* (0.01)
Greece	3.36* (0.00)	2.34** (0.02)
Ireland	2.32** (0.02)	0.84 (0.40)
Italy	0.50 (0.62)	1.10 (0.27)
The Netherlands	3.26* (0.00)	0.28 (0.78)
Portugal	0.12 (0.90)	0.66 (0.51)
Spain	1.22 (0.22)	2.69* (0.01)
United Kingdom	4.58* (0.00)	0.68 (0.50)

**,* denote significance at 0.05 and 0.01, respectively. The numbers in parenthesis are the asymptotic p-values.

4. Conclusions and policy implications

This paper attempts to cast light on the existence of the adjustment gasoline price mechanism within the eleven Eurozone countries, the UK and the US. For this reason, we distinguish the process of transmission of oil price shocks to gasoline prices into two parts corresponding to a first wholesale level and then to a second distribution stage (retail level). By doing so, we are able to assess possible asymmetries at either one or both levels. In the specific study, we use two asymmetric ECMs at each market segment in order to distinguish between asymmetries arising from short-lived deviations in input prices (crude oil and spot prices) and asymmetries concerning the speed at which the gasoline price reverts to its long-run (equilibrium) level. Moreover, we allow for an explicit role of the fuel taxation and a detailed competition analysis to which rather scant attention has been paid by the earlier studies.

In order to obtain asymptotically efficient estimators, we use DOLS instead of OLS in contrast to the vast majority of the previous studies. The empirical findings indicate that in the European countries there is a symmetric response of the output prices of gasoline without taxes (excise and value added tax) in the wholesale level both in the short and the long run, respectively. This conclusion is in alignment with other empirical studies and runs contrary to the common perception regarding the price asymmetries that emerge in the gasoline market. Similar results can be found when testing for exchange rate asymmetry in the wholesale level.

In the retail segment the long-run price effect on the net retail price of gasoline is positive and substantial in magnitude, with the relevant coefficients bellow unity. In the majority of European countries, such as Austria, Ireland, Portugal, The Netherlands, Greece, Italy and Spain, a change in the gasoline spot price is not fully passed through to the net retail price indicating a relatively smaller pass-through price mechanism compared to the wholesale segment. In the short-run the retail gasoline prices seem to react more to price increases and to negative gaps to the equilibrium than to price decreases and positive disequilibrium.

If we try to compare the two-level analysis, some interesting remarks emerge. First, the magnitude of short-run coefficients is in the most sample countries larger in the wholesale than in the retail level. Second, the adjustment towards the equilibrium level is more gradual in the retail level, revealing the structural differences between the wholesale and retail segment of the gasoline industry. Furthermore, the retailers tend to react more to price increases than price decreases compared to the wholesalers, indicating a different adjustment path to the long-run equilibrium level of price. Lastly, from the relevant magnitude of the price coefficients in the wholesale and the retail equations, we assume that retailers do not immediately transfer onto final prices (pump prices) all the adjustments in the wholesale prices.

Regarding the effect of taxation on gasoline price responses, we infer that in some countries positive short-run spot price effect is larger than its negative counterpart, while the speed of adjustment justifies the existence of price asymmetry in many European countries (i.e. Belgium, France, The Netherlands, Ireland, UK, Greece, Spain and Portugal).

In many European sample countries the oligopolistic structure of the local gasoline markets along with crude oil volatility enables the price asymmetric adjustment path in the retail oil segment (Greece, Spain, The Netherlands, Portugal and Ireland). However, in European countries such as France Austria, Belgium, crude oil volatility seems to be the only cause of asymmetric price adjustment in the retail segment.

The gasoline price adjustment in the United Kingdom reveals that local gasoline market is characterized by a high degree of competition. This implies that long-run variations in the spot price of gasoline are not followed by similar variations in the final retail price. In the short-run, the positive spot price effect is larger than its negative counterpart. This result may be attributed to the gradually increase of fuel taxes within the last twenty years.

Lastly, in the US, the hypothesis of the symmetric adjustment speeds appears to be valid only in the short-run. The long-run asymmetric adjustment speed along with crude oil volatility may reflect the relevant high profit margins of the US companies in tandem with their positive financial performance within the last decade.

Appendix

Table A1
Estimation results of the ECMs without the effect of taxation.
Source: Authors' elaboration.

Variables	Austria	Belgium	Finland	France	Germany	Greece	Ireland	Italy	The Netherlands	Portugal	Spain	UK	US
c	−0.001 (−0.000)	−0.001 (−0.002)	−0.001 (−0.001)	−0.001 (−0.001)	−0.001 (−0.002)	−0.001 (0.001)	−0.001 (0.002)	−0.001 (−0.000)	−0.000 (0.002**)	−0.001 (0.001)	−0.001 (−0.000)	−0.001 (−0.000)	−0.005 (0.000)
ΔSPG_{t-1}	1.202*	1.173*	1.120*	1.187*	1.123*	0.891*	1.245*	1.176*	1.181*	1.161*	1.201*	1.244*	–
ΔSPG_{t-2}	−0.052**	−0.050**	−0.052**	−0.051**	−0.048**	0.680*	−0.053**	−0.051**	−0.052**	−0.050**	−0.052**	−0.052**	–
$\Delta NRPG_{t-1}$	(0.500*)	(0.411*)	–	(0.448*)	(0.384*)	(0.553*)	(0.088)	(0.512*)	(0.100*)	(0.298*)	(0.529*)	(0.467*)	(0.437*)
$ECMP_{t-1}$	−1.087* (−0.693*)	−1.059* (−0.601*)	−1.059* (−0.121*)	−1.074* (−0.406*)	−1.012* (−0.660*)	−0.791* (−0.648*)	−1.137* (−0.106***)	−1.062* (−0.356*)	−1.067* (−0.571*)	−1.048* (−0.165*)	−1.088* (−0.547*)	−1.136* (−0.213*)	−0.244* (−0.119**)
$ECMN_{t-1}$	−0.974* (−0.519*)	−0.946* (−0.781*)	−0.946* (−0.185*)	−0.961* (−0.208*)	−0.896* (−0.668*)	−0.670* (−0.602*)	−1.018* (−0.041)	−0.949* (−0.377*)	−0.955* (−0.256*)	−0.934* (−0.187*)	−0.975* (−0.357*)	−1.017* (−0.223*)	−0.213* (−0.064***)
ΔCRP_t	0.651*	0.654*	0.650*	0.652*	0.657*	0.680*	0.662*	0.653*	0.650*	0.652*	0.651*	0.668*	0.745*
ΔCRP_{t-1}	−0.729*	−0.711*	−0.729*	−0.719*	−0.672*	−0.508*	−0.761*	−0.710*	−0.712*	−0.699*	−0.728*	−0.760*	0.158*
ΔCRN_t	0.690*	0.695*	0.689*	0.691*	0.695*	0.704*	0.699*	0.691*	0.690*	0.691*	0.689*	0.707*	0.801*
ΔCRN_{t-1}	−0.930*	−0.907*	−0.933*	−0.920*	−0.873*	−0.705*	−0.967*	−0.912*	−0.916*	−0.900*	−0.931*	−0.969*	–
$\Delta SPGP_t$	(0.424)*	(0.452*)	(0.444*)	(0.410*)	(0.505*)	(0.315*)	(−0.011)	(0.226*)	(0.463*)	(0.059)	(0.296*)	(0.157*)	(0.370*)
$\Delta SPGN_t$	(0.334)*	(0.471*)	(0.436*)	(0.233*)	(0.366*)	(0.363*)	(−0.025)	(0.218*)	(0.481*)	(0.099**)	(0.207*)	(0.066*)	(0.216*)
$\Delta SPGP_{t-1}$	–	–	(−0.103*)	–	–	–	–	–	–	–	–	–	(−0.083*)
$\Delta SPGN_{t-1}$	–	–	–	–	–	–	–	–	–	–	–	–	(0.083*)
$\Delta EXRP_t$	−0.129	−0.081	−0.167	−0.129	−0.152	−0.139	0.030	−0.150	−0.191	−0.159	−0.155	0.063	–
$\Delta EXRN_t$	−0.105	−0.106	−0.099	−0.091	−0.086	−0.118	−0.031	−0.092	−0.026	−0.092	−0.088	−0.021	–
Diagnostics													
Adjusted R^2	0.545 (0.423)	0.550 (0.277)	0.550 (0.190)	0.550 (0.633)	0.550 (0.303)	0.541 (0.489)	0.543 (0.005)	0.544 (0.524)	0.545 (0.525)	0.544 (0.588)	0.545 (0.506)	0.544 (0.256)	0.466 (0.700)
Durbin–Watson	1.995 (2.022)	1.997 (2.175)	1.996 (2.031)	1.995 (2.154)	1.997 (2.133)	1.981 (2.215)	2.002 (2.008	1.995 (2.093)	1.996 (2.054)	1.995 (2.082)	1.996 (2.043)	1.998 (2.205)	2.041 (2.210)

C denotes the constant term. ** and * denote significance at 0.05 and 0.01. respectively.

Table A2
Estimation results of the ECMs with the effect of taxation.
Source: Authors' elaboration.

Variables	Austria	Belgium	Finland	France	Germany	Greece	Ireland	Italy	The Netherlands	Portugal	Spain	UK
$\Delta FPRV_{c,t} = a_0 + \sum_{i=1}^{p} b_i FPRV_{c,t-i} + \sum_{i=0}^{m} c_i^+ \Delta SPGP_{r,t-i} + \sum_{i=0}^{n} c_i^- \Delta SPGN_{r,t-i} + \lambda^+ ECMP_{t-1} + \lambda^- ECMN_{t-1} + \varepsilon_t$												
c	−0.000 (0.21)	−0.001(0.06)	−0.000 (0.59)	−0.000** (0.01)	−0.000 (0.28)	0.001 (0.12)	0.002** (0.04)	−0.000 (0.67)	0.001 (0.07)	0.001 (0.07)	0.000 (0.46)	−0.002**(0.03)
$\Delta FPRV_{t-1}$	0.508* (0.00)	0.429* (0.00)	−0.089 (0.13)	0.473* (0.00)	0.295* (0.00)	0.584* (0.00)	0.031 (0.63)	0.540* (0.00)	0.114* (0.00)	0.527* (0.00)	0.863* (0.00)	0.200** (0.05)
$ECMP_{t-1}$	−0.639* (0.00)	−0.603* (0.00)	−0.125 (0.10)	−0.550* (0.00)	−0.606* (0.00)	−0.565* (0.00)	−0.125 (0.06)	−0.494* (0.00)	−0.605* (0.00)	−0.260* (0.00)	−0.804* (0.00)	0.147 (0.16)
$ECMN_{t-1}$	−0.639* (0.00)	−0.787* (0.00)	−0.136 (0.08)	−0.362* (0.00)	−0.510* (0.00)	−0.557* (0.00)	0.068 (0.35)	−0.424* (0.00)	−0.267* (0.00)	−0.149* (0.00)	−1.024* (0.00)	−0.245**(0.02)
$\Delta SPGP_t$	0.186* (0.00)	0.167* (0.00)	0.178* (0.00)	0.196* (0.00)	0.204* (0.00)	0.105* (0.00)	−0.021 (0.39)	0.126* (0.00)	0.170* (0.00)	0.010 (0.36)	0.065* (0.00)	0.072** (0.03)
$\Delta SPGN_t$	0.154* (0.00)	0.151* (0.00)	0.162* (0.00)	0.127* (0.00)	0.136* (0.00)	0.162* (0.00)	0.009 (0.68)	0.111* (0.00)	0.175* (0.00)	0.022** (0.04)	0.135* (0.00)	0.056 (0.08)
Adjusted R^2	0.452 [2.015]	0.252 [2.170]	0.216 [2.041]	0.598 [2.087]	0.326 [2.171]	0.433 [2.119]	0.598 [2.087]	0.462 [2.023]	0.477 [2.035]	0.229 [2.180]	0.330 [1.965]	0.072 [1.994]
$\Delta FPREX_{c,t} = a_0 + \sum_{i=1}^{p} b_i FPREX_{c,t-i} + \sum_{i=0}^{m} c_i^+ \Delta SPGP_{r,t-i} + \sum_{i=0}^{n} c_i^- \Delta SPGN_{r,t-i} + \lambda^+ ECMP_{t-1} + \lambda^- ECMN_{t-1} + \varepsilon_t$												
c	−0.000 (0.43)	−0.003** (0.05)	−0.000 (0.97)	−0.001** (0.04)	−0.003 (0.09)	−0.000 (0.95)	−0.004** (0.04)	−0.001 (0.23)	0.002 (0.06)	0.000 (0.65)	0.001 (0.24)	−0.005**(0.01)
$\Delta FPREX_{t-1}$	0.524*(0.00)	0.429* (0.00)	0.003 (0.97)	0.520* (0.00)	0.425* (0.00)	0.564* (0.00)	0.079 (0.23)	0.563* (0.00)	0.107* (0.006)	0.455* (0.00)	0.676* (0.00)	0.274* (0.00)
$ECMP_{t-1}$	−0.729*(0.00)	−0.569* (0.00)	−0.156 (0.07)	−0.702* (0.00)	−0.693* (0.00)	−0.527* (0.00)	−0.164** (0.02)	−0.478* (0.00)	−0.581* (0.00)	−0.282* (0.00)	−0.548* (0.00)	0.081 (0.27)
$ECMN_{t-1}$	−0.590*(0.00)	−0.842* (0.00)	−0.222** (0.02)	−0.314* (0.00)	−0.699* (0.00)	−0.878* (0.00)	0.059 (0.44)	−0.531* (0.00)	−0.257* (0.00)	−0.294* (0.00)	−0.688* (0.00)	−0.281* (0.00)
$\Delta SPGP_t$	0.390*(0.00)	0.386* (0.00)	0.407* (0.00)	0.451* (0.00)	0.429* (0.00)	0.312* (0.00)	−0.031 (0.52)	0.241* (0.00)	0.396* (0.00)	0.065 (0.07)	0.090* (0.00)	0.128** (0.02)
$\Delta SPGN_t$	0.303*(0.00)	0.406* (0.00)	−0.083 (0.16)	0.251* (0.00)	0.279* (0.00)	0.465* (0.00)	0.033 (0.44)	0.207* (0.00)	0.408* (0.00)	0.102* (0.00)	0.191* (0.00)	0.072 (0.17)
Adjusted R^2	0.445 [2.018]	0.269 [2.177]	0.221 [2.053]	0.650 [2.004]	0.366 [2.126]	0.391 [2.080]	0.005 [2.012]	0.500 [2.039]	0.509 [2.042]	0.129 [2.117]	0.354 [2.108]	0.091 [2.004]
$\Delta FPR_{c,t} = a_0 + \sum_{i=1}^{p} b_i FPR_{c,t-i} + \sum_{i=0}^{m} c_i^+ \Delta SPGP_{r,t-i} + \sum_{i=0}^{n} c_i^- \Delta SPGN_{r,t-i} + \lambda^+ ECMP_{t-1} + \lambda^- ECMN_{t-1} + \varepsilon_t$												
ΔFPR_{t-1}	0.511* (0.00)	0.429* (0.00)	−0.086 (0.17)	0.485* (0.00)	0.316* (0.00)	0.597* (0.00)	0.039 (0.55)	0.548* (0.00)	0.117* (0.00)	0.490* (0.00)	0.840* (0.00)	0.180** (0.04)
ΔFPR_{t-2}	–	–	−0.073** (0.03)	–	–	–	–	–	–	–	–	–
$\Delta SPGP_t$	0.208* (0.00)	0.190* (0.00)	0.200* (0.00)	0.221* (0.00)	0.223* (0.00)	0.157* (0.00)	−0.022 (0.41)	0.139* (0.00)	0.191* (0.00)	0.014 (0.29)	0.077* (0.00)	0.071** (0.02)
$\Delta SPGN_t$	0.169* (0.00)	0.177* (0.00)	0.188* (0.00)	0.138* (0.00)	0.149* (0.00)	0.215* (0.00)	0.012 (0.63)	0.120* (0.00)	0.195* (0.00)	0.027** (0.03)	0.136* (0.00)	0.058** (0.05)
$\Delta SPGN_{t-1}$	–	–	–	–	–	–	–	–	–	–	–	0.068* (0.00)
$ECMP_{t-1}$	−0.648* (0.00)	−0.594* (0.00)	−0.140 (0.09)	−0.598* (0.00)	−0.629* (0.00)	−0.587* (0.00)	−0.126 (0.06)	−0.504* (0.00)	−0.619* (0.00)	−0.260* (0.00)	−0.741* (0.00)	0.181** (0.05)
$ECMN_{t-1}$	−0.630* (0.00)	−0.805* (0.00)	−0.152** (0.07)	−0.353* (0.00)	−0.536* (0.00)	−0.733* (0.00)	0.060 (0.42)	−0.435* (0.00)	−0.276* (0.00)	−0.028 (0.68)	−0.924* (0.00)	−0.233* (0.01)
c	−0.001 (0.21)	−0.002** (0.05)	−0.000 (0.72)	−0.001** (0.02)	−0.001 (0.27)	0.001 (0.49)	0.002** (0.05)	0.000 (0.60)	0.001 (0.07)	0.001** (0.03)	0.000 (0.50)	−0.001 (0.23)
Adjusted R^2	0.459 [2.018]	0.258 [2.176]	0.220 [2.014]	0.614 [2.071]	0.332 [2.164]	0.421 [2.073]	0.001 [2.005]	0.480 [2.015]	0.487 [2.038]	0.204 [2.128]	0.387 [1.975]	0.102 [1.977]

Notes: The numbers in square brackets refer to the Durbin Watson statistic. The numbers in the parenthesis refer to *p*-values. ** and * denote significance at 0.05 and 0.01, respectively.

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