On the Examination of Competition in the Petroleum Industry: A Pooled Panel Threshold Analysis

Chen, Chaoyi and Polemis, Michael and Stengos, Thanasis

University of Guelph, University of Piraeus, University of Guelph

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Chaoyi Chen[a], Michael Polemis[b,c*] and Thanasis Stengos[a]

Abstract

This paper contributes to the literature since it tries to link the Exchange Rate Pass-Through (ERPT) with the "*rockets and feathers*" hypothesis using a panel of EU-28 countries. Allowing for the existence of an endogenous threshold variable our empirical findings indicate that the threshold model is better suited to this analysis than the baseline linear adjustment model. This is the case since the latter restricts the threshold to be centered around zero and the dynamic response to cumulative shocks cannot be properly identified. The empirical findings reveal that the threshold variable expressed by the trade-weighted dollar exchange rate index is statistically significant only in the sample above the threshold (high regime). This means that for the net EU exporting countries, fluctuations in the real effective exchange rate of the US against its major EU trading partners does affect the level of pre-tax retail gasoline prices with the relevant elasticity exceeding unity (complete ERPT). Moreover, all the statistical tests reject the null hypothesis that there is no significant threshold and thus an asymmetric adjustment gasoline mechanism prevails.

Key words: Asymmetric gasoline adjustment; ERPT; Threshold analysis; Exchange rate; Non-linear effects
JEL classifications: F41; C14; C23; C24

[a] University of Guelph, Department of Economics, University of Guelph, Guelph, Ontario, N1G 2W1

[b*] University of Piraeus, Department of Economics Piraeus, Greece, E-mail address: mpolemis@unipi.gr (corresponding author).

[c] Hellenic Competition Commission, Athens, Greece

1. Introduction

ERPT, namely the change in import prices resulting from an exchange rate shock, is an important topic in Economics that has received significant attention from the researchers within the last twenty years (see for example Camba and Goldberg, 2005; Gopinath et al, 2010; Ceglowski, 2010; Devereux and Yetman, 2010; Aguerre et al, 2012; Auer and Schonle, 2016).

From an international economics perspective, a key question is to what extent the exchange-rate fluctuations are passed-through to the prices of imported goods (Fabra and Reguant, 2014). Exchange rate fluctuations between dollar and other currencies play a crucial role in determining the transmission pricing mechanism in commodity markets including oil industry as well (Galeotti et al, 2003). As a consequence, the estimation of sensitivity (elasticity) of local-currency import prices (i.e gasoline prices) to changes in local-currency price of foreign currency known as ERPT is of paramount importance for controlling the transmission of inflation between countries, testing the law of one price and the existence of Purchasing Power Parity (Goldberg and Knetter, 1997; Camba and Goldberg, 2005; Krugman, 1986; Helpman and Krugman, 1987).

Within the last years there is a plethora of studies in the Industrial Organization (IO) literature investigating the existence of gasoline price asymmetry with controversial results. Most of these studies apply cointegration techniques by utilizing an asymmetric (vector) error-correction model (Borenstein et al., 1997; Eckert, 2002; Galeotti et al., 2003; Deltas, 2008; Polemis, 2012; Wlazlowski et al, 2012; Greenwood-Nimmo and Shin, 2013; Bumbass et al., 2015; Kristoufek and Lunackova, 2015; Blair et. al, 2017; Eleftheriou et al, 2018), while others rely on non-parametric methods (Godby et al, 2000; Mann, 2016; Polemis and Tsionas, 2016; 2017; Bagnai et. al, 2018) in order to uncover the existence of price asymmetries. The asymmetric price adjustment mechanism has also been examined on a theoretical ground as well. Theories of asymmetric price adjustment identify possible causes of asymmetry in a number of reasons such as *inter alia* tacit collusion (Radchenko, 2005), inventory capacity and hoarding (Borenstein and Shepard, 1996), and consumer search (Johnson, 2002).

Despite the rich body of literature, existing studies fail to explain the role of exchange rate fluctuations in determining the causes of the asymmetric gasoline adjustment path (commonly known as "*rockets and feathers*" hypothesis). [1] In particular, past studies have been methodologically restrictive in the sense that the retail gasoline short-run responses, given an input (crude) cost shock, were attributed to crude oil fluctuations. However, "*these studies would therefore be biased these studies would therefore be misspecified if mark-up rules were actually described by an alternative relationship, as would be the case if, for example, price asymmetries were instead triggered by a minimum absolute increase in crude cost*" (Godby et al, 2000). Specifically, the authors argue that this is a possibility, not that it is the usual case and try to estimate a TAR to investigate this possibility, but do not find any evidence of asymmetric pricing in the Canadian market.

Using several possible exchange rate-retail price relationships, we attempt to determine whether an asymmetric pricing pattern in the weekly data for 28 EU countries can be explained by the ERPT mechanism. This approach traces the effects of the exchange rate on the coefficient of each regressor (marginal response) over the sample. In this case, the trade-weighted dollar exchange rate index acts as a threshold variable

[1] This means that prices increase rapidly in response to cost increases (like a rocket) but fall only slowly in response to cost decreases (like a feather).

in order to capture the marginal effect of a given variable as an unknown function of an observable covariate, introducing heterogeneity. Subsequently, the EU-28 countries will be sorted according to their level of international competitiveness toward the US economy placing them into net exporters (high regime countries) and net importers (low regime countries) respectively. This happens since a rise of the exchange rate index tends to increase the value of the US imports and lower the value of the exports. Therefore, EU countries increase their exports to the US compared to their imports (net exporters). The opposite mechanism is triggered when the relevant index decreases.

The contribution of this paper is three-fold. First, it goes beyond the existing literature in that it uses a particularly long panel of EU-28 countries at a weekly basis. Second, in contrast to the existing empirical studies which assume that the variables are not correlated across the panel dimension (cross sectional independence) we perform appropriate cointegration techniques in order to deal with this issue. The latter may arise due to common unobserved effects generated by changes in the European legislation (i.e taxation, currency regulatory restrictions, import quotas, etc). Third and foremost, it is the first study to our knowledge that tries to examine the impact of the ERPT on asymmetric gasoline price adjustment. Moreover, the application of the dynamic panel GMM threshold model developed by Seo and Shin (2016) constitutes an additional novelty of this paper. Previous studies assume the threshold to be zero. However, it is possible that this might not be the case for the European gasoline market as a whole. It may be possible that the threshold lies at some positive value or it may be that the asymmetric behaviour is not triggered until a certain change in input price is felt in some fixed time period (Godby et al, 2000). Using the GMM threshold model allows us to test for possible asymmetric gasoline pricing mechanism triggered by exchange rate fluctuations.

In this study, we employ a pooled panel threshold model within an error correction framework and allowing for the presence of an endogenous threshold variable to investigate the following research questions: Is there evidence of short-run gasoline asymmetric pricing in the EU-28 as a whole over the sample period? Does the ERPT mechanism constitute a possible cause of gasoline asymmetric adjustment? Are there any non-linear effects in "*rockets and feathers*" hypothesis? Asymmetric pricing is tested for in both net and final retail unleaded EU gasoline markets. The empirical findings confirm the superiority of the threshold model compared to the baseline linear specifications, while attributing the asymmetric gasoline adjustment mechanism to ERPT.

The rest of the paper is organized as follows. Section 2 provides a comprehensive survey to the ERPT literature. Section 3 describes the data while Section 4 presents the empirical models (baseline and threshold model) estimated in this paper and discusses econometric issues. Section 5 reports the estimation results, Section 6 concludes the paper.

2. Literature review

The literature on ERPT starts with the seminal paper of Kreinin (1977) who uses an experimental approach to estimating the degree of ERPT in six OECD countries (US, Japan, Canada, Germany, Belgium and Italy). He finds an incomplete ERPT for all the sample countries except for Italy (100%). This is attributed to factors such as the different level of market power prevailing in each country or the ability of the importing country to influence the world price due to its relatively large size.

However, the majority of the empirical studies regarding ERPT use linear econometric models (i.e log linear, error correction models, VAR, etc) dealing with stationarity and cointegration properties where the dependent (exogenous) variable is the import price regressed on several control/predetermined variables such as exporter's cost, competing prices, income (GDP), and nominal exchange rate between the importing and the exporting country (see for example Woo, 1984; Hooper and Mann 1989). The coefficient of the estimated nominal exchange rate variable denotes the elasticity of domestic/importing prices to variations in the exchange rate referred to as the pass-through coefficient.[2] All of these studies consent that the ERPT in the US is incomplete ranging from 50-60%, where the rest (50-40%) of the exchange rate change is offset by changes in the markup (Goldberg and Knetter, 1997). One possible explanation for such asymmetric pass-through is that firms adjust their markups to accommodate the local market environment (Krugman, 1986; Helpman and Krugman, 1987). The study of Feenstra, (1989), sheds some light on the explanation of the incomplete ERPT by linking the latter to the presence of imperfect competition.[3] Feenstra uses a log-linear model and quarterly data over the period 1974:1 to 1987:1 for the U.S. imports of Japanese cars, compact trucks and heavy motorcycles to find that there is a symmetric response of import prices to changes in the bilateral exchange rate and an import tariff.

A number of past studies also investigate the extent of ERPT using disaggregated industry level data. More specifically, Dornbusch (1987) uses two-digit industry level data to link the incomplete ERPT with micro-economic factors (i.e market concentration, product homogeneity, market shares). Yang (1997), uses monthly data for the 87 (three and four-digit SIC) manufacturing sectors over the period from 1980:12 to 1991:12 in order to estimate the speed of ERPT in the US industry sector.

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[2] If the estimated elasticity γ is less than unity then the ERPT is incomplete, otherwise is full or complete (γ=1).
[3] The study of Engel (2002) provides a complete review of the possible ERPT explanations.

He adopts a two-stage procedure, in which the ERPT elasticities are estimated through a typical log linear model expressed in first differences and these estimates are regressed against several independent variables (costs, market power, market concentration, etc). His findings suggest that ERPT is asymmetric and varies across industries. The degree of pass-through is positively (negatively) correlated to product differentiation, (elasticity of marginal cost). Subsequent work by Taylor (2000) argues that the responsiveness of ERPT depends positively on the level of inflation in a sense that low ERPT in low inflation countries comes as a result of the low inflation environment.

Other studies such as Schröder and Hüfner (2002), Choudhri et al. (2002), Choudhri and Hakura (2002), Hahn (2003), Bailliu and Fujii (2004), Gagnon and Ihrig (2004), Choudhri et al. (2005), Faruqee (2006), and Campa and Goldberg (2006a and b) have tried to explore the impact of ERPT on import prices and core inflation in the euro zone area or a number of European Monetary Union (EMU) countries by applying standard econometric techniques (log linear models, ECMs and VARs) with controversial results about the rate and the causes of the adjustment.

In an interesting study, Campa and Minquez (2006), investigate the ERPT into the import prices of twelve EMU countries originating outside the eurozone area. They use monthly time series data over the period 1989:1 to 2001:3 for thirteen different product categories for each country. They argue that in the short-run, ERPT is incomplete since the estimated pass-rate coefficients (elasticities) are in their vast majority less than one ($\gamma<1$). However, the same conclusion does not hold in the long run where it is reported a symmetric ERPT. McCarthy (2006) also examines the speed of ERPT on producer and consumer prices for nine selected industrialized countries. He estimates a parsimonious VAR model including variables such as oil price inflation, output gap, nominal exchange rate, import price inflation, consumer and producer price inflation, short-term interest rate and money growth. His results confirm the aforementioned literature suggesting an incomplete ERPT due to market distortions (lack of effective competition).

Subsequent work by Gopinath et al (2010) investigates the ERPT by developing a dynamic currency choice model. They use monthly time series (at a country level) and panel data (at industry level) on the US import prices for dollar and non-dollar goods over the period 1994-2005 to find that there is a large difference in the pass-through between the two pricing categories. The econometric methodology is based on (fixed effects) OLS estimators employing standard pass-through regression models appeared in first differences. These findings have also been corroborated by the studies of Bhattacharya et al (2008), Ceglowski (2010), Devereux and Yetman (2010) and Aguerre et al (2012).

The impact of market structure on the ERPT nexus is more evident in the recent study of Auer and Schonle (2016). The authors use annual firm-level data on standard ERPT regression analysis over the period 1994-2005 for the thirty four largest trading partners of the US. They argue that market share affects the rate at which firms react to changing competitor prices.

Earlier work by Al-Abri and Goodwin (2009) and Aleem and Lahiani (2014) stands apart from those discussed above in that it uses non linear econometric methodology. Al-Abri and Goodwin (2009) use a threshold cointegration model (TAR) in order to reveal the determinants of the ERPT in sixteen OECD countries and five categories of imported goods (Food and agricultural products, energy, raw materials, manufacturing, and non-manufacturing). The authors use quarterly data spanning the period 1975:1 to 2002:2 to support that in their non-linear model the import prices respond faster and by a larger degree to nominal exchange rate fluctuations than in the standard log linear models. On the other hand, Aleem and Lahiani (2014) rely on the flexible threshold vector autoregression model (TVAR) to investigate the degree of ERPT rate in Mexico by utilizing monthly seasonally adjusted data from 1994:1 to 2009:11. They find that domestic prices react strongly to a positive one unit exchange rate shock only above the threshold level of the rate of inflation.

Although the issue of ERPT into domestic prices is well documented in the literature, there are few studies focusing on products that are relatively homogeneous and priced in an international market known as "*commodities*" (i.e petroleum prices, agriculture products, precious metals, etc).

Yanagisawa (2012) uses weekly data for the Japan over the period January 2012 to February 2013 and ECM techniques in order to investigate the ERPT into domestic oil price. He decomposes the pass through structure of gasoline price into two distinct features comprising of the dollar and the exchange rate factor. It is worth mentioning that this study considers the issue of the "*numeraire*" currency (dollar) for the ERPT into commodity pricing. He finds an incomplete but rather symmetric of the pass-through rate of the dollar factor, a premise also supported by the empirical literature. The opposite result is confirmed when the pass-through of the exchange rate factor is taken into account.

Finally, Akçelik and Ogünç (2016) examine the degree of ERPT to domestic fuel prices at different oil market segments in Turkey over the period 2004-2014. They use monthly data and VAR methodology to depict that the ERPT to domestic gasoline prices is considerably fast and just one third of a change in crude oil prices is reflected to the gasoline prices. This is attributed to the significant share of taxation on retail prices. On the other hand, they argue that the impact of oil prices on transport services takes a longer time compared to other domestic prices, suggesting that a 10% change in the international crude oil prices is associated with a 0.42% change in consumer inflation at the end of one year.

All in all the majority of the above ERPT papers treat the exchange rate as a cost shifter. They have no distinction between the change in the price of the product and change in the exchange rate. The reason is that the product typically does not have an international price denominated in a specific currency.

3. Data and variables

We use a large unbalanced panel dataset of weekly observations spanning the period from January 1994 to January 2015. The primary sample includes all 28 European Union countries, but the coverage for each country varies, largely because of differences in accession dates into the EU. All variables are in their natural logarithms expressed in real terms and deflated by the Harmonised Consumer Price index provided by Eurostat. Input cost price (i.e Brent crude oil price) measured in dollars per barrel is taken from the USA Department of Energy (EIA).[4] It is worth mentioning that, the coverage period for the tax-inclusive gasoline price (price at the pump) is more limited than the coverage period for the pre-tax (net) retail gasoline price.

Pre-tax gasoline retail prices expressed in local currencies are obtained from the Weekly Oil Bulletin.[5] It is worth mentioning that pre-tax prices are used to avoid the possibility that countries with heterogeneous excise tax levels (e.g Italy and Estonia) experience very different percentage responses to one percent change in the underlying marginal cost, solely because the fixed amount of the excise tax moves up the origin of

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[4] https://www.eia.gov/dnav/pet/pet_pri_spt_s1_d.htm.
[5] http://ec.europa.eu/energy/en/data-analysis/weekly-oil-bulletin.

the retail price. However, we will also estimate the final specifications with post-tax retail prices (final prices) to check for the robustness of our findings.

The exchange rate effect is quantified by two indicators: a) The Dollar trade-weighted exchange rate index (1997=100) which is drawn directly from the Federal Reserve Bank of St. Louis, and b) The nominal effective Euro trade-weighted exchange rate index obtained by the European Central Bank. The first term is the change in the trade-weighted value of the dollar (or the consumption weighted dollar exchange rate), and the second term is the change in the number of units of local currency to the dollar.

Specifically the Dollar trade-weighted exchange rate index (commonly known as "*broad*" index) is the weighted average of the foreign exchange value of the U.S. dollar against the currencies of a broad group of major U.S. trading partners (FRED, 2017)[6]. This index, which will act as the endogenous threshold variable in our model, is used to determine the U.S. dollar purchasing value, and to summarize the effects of dollar appreciation and depreciation against foreign currencies. When the value of the dollar increases, imports to the U.S. become less expensive while exports to other countries become more expensive. In other words, if the index rises (decreases), ceteris paribus, the purchasing power of the US dollar also rises (decreases) which will reduce (increase) the cost of imports but will undermine (enhance) the competitiveness of the US exports.[7] Alternatively, if this index rises (decreases), the value of the EU (and of

[6] This index includes the Euro Area, Canada, Japan, Mexico, China, United Kingdom, Taiwan, Korea, Singapore, Hong Kong, Malaysia, Brazil, Switzerland, Thailand, Philippines, Australia, Indonesia, India, Israel, Saudi Arabia, Russia, Sweden, Argentina, Venezuela, Chile and Colombia.

[7] Trade-weighted dollar index places importance (weight) to currencies most widely used in international trade, over comparing the value of the U.S. dollar to all foreign currencies. Since the currencies are weighted differently, changes in each currency will have a unique effect on the trade-weighted dollar and their corresponding indexes.

the other foreign countries as well) exports (imports) to the US also rises (decreases) constituting the EU countries as net exporters (importers).

One could argue that ranking or splitting countries according to their exports/imports to and from the USA seems arbitrary. The reason is that many EU countries are not really dependent on the USA, mainly the smaller ones that are much more dependent on exports within the EU (Germany, Greece, Portugal, etc). However, the broad index was introduced by the U.S. Federal Reserve Board in 1998 in response to the implementation of the euro (which replaced many of the foreign currencies that were previously used in the earlier index) and to more accurately reflect current U.S. trade patterns. The Federal Reserve selected 26 currencies to use in the broad index, anticipating the adoption of the euro by eleven countries of the European Union (EU). It is noteworthy that when the broad index was introduced, U.S. trade with the 26 represented economies accounted for over 90% of the total U.S. imports and exports (FRED, 2017).

The second exchange rate factor can be represented by the inclusion of the nominal effective Euro trade-weighted exchange rate index. The latter denotes a geometric weighted average of the bilateral exchange rates of the euro against the currencies of a selection of trading partners. More specifically, this indicator is computed against a group of 42 partner countries (EER-42), accounting for roughly 90% of total euro area manufacturing trade in 1999-2001. It is worth mentioning that a fixed weighting scheme is employed in these computations. According to the ECB, the scheme is based on manufacturing trade and takes into account so-called third-market effects, (i.e. competition faced by euro area products in a partner country from products of a third country). This index was first constructed in 1999 and the first update of the weights took place in 2004. Moreover, the overall trade weights underpinning the EER-

42 index are updated every five years. Similarly with the other exchange rate index, the interpretation of this indicator is straightforward. In particular, if the index rises (decreases), ceteris paribus, the euro appreciates (depreciates) against its major trading countries resulting in a reduction (increase) of the exports (imports).

Based on the above considerations, we argue that the ERPT specifications differ from the "*standard*" specifications provided by the IO literature (see among others Galeotti et al, 2003; Deltas, 2008; Polemis and Tsionas, 2017) in the following ways. First, all prices are in logs and coefficient estimates denote elasticities since there is no other meaningful way to jointly estimate the model involving series from different countries in different units. Second, the retail prices are in local currency, and not in euros. Pre-tax prices are used to avoid the possibility that countries with very different (fixed amount) excise tax levels experience very different percentage responses to one percent change in the underlying marginal cost, solely because the fixed amount of the excise tax moves up the origin of the retail price. Third, the input price is the "*real*" price of crude oil (i.e., the price deflated by the US dollar price index). The deflator that we used is the trade-weighted value, but we have also used the consumption-weighted values as a robustness check. Fourth, we have included two exchange rate terms that will be treated in exactly the same way as we treat input prices, i.e., we will have the lags, and in the asymmetric model we will distinguish between positive and negative changes. Note that the two exchange rate terms will be treated in exactly the same way as we treat input prices (i.e., we have the lags, and in the asymmetric model we distinguish between positive and negative changes). They may also be in the co-integration vector, but an alternative is to have the co-integration vector be in a common currency (e.g., euros, under the premise that in the long run pass-through is equal to one).[8] Specifically, the first term is the change in the trade-weighted value of the dollar (or the consumption weighted dollar exchange rate), and the second term is the change in the number of units of local currency to the dollar. These changes will be differences in the log values of the corresponding variables. [9] Finally, our approach allows for an endogenous treatment of all the regressors and the threshold variable at the same time, contrary to the threshold autoregressive model of Godby et al (2000).

Table 1, provides a complete description of the variables (expressed in natural logarithms) included in this study. As it is evident over the sample period, net retail gasoline prices (not including taxes) averaged 6 dollars per gallon while final gasoline prices were approximately 70 cents higher (6.7). As it is expected the retail gasoline prices and crude oil fluctuations follow a similar pattern. Specifically, gasoline prices have been rising slightly over the examined period, with a drift of 0.08 cents per week. Regarding the short run price fluctuations it is important to note that the standard deviation of net retail prices (expressed in Euros) is smaller than that of crude oil (Brent) and spot gasoline price (New York) suggesting the existence of a "*dampening*" effect in the gasoline market (Polemis and Tsionas, 2017; Deltas, 2008). In other words, retail gasoline prices are relatively sticky and do not fully transmit short run fluctuations in the input prices.

Table 1: Descriptive statistics

[8] In such a case, the basic equation becomes $\Delta\ln(R_{j,t}^{lc}) = a_j + b_{0,j}\Delta\ln(C_t^r) + b_{1,j}\Delta\ln(C_{t-1}^r) + b_{0,j}^{W\$}\Delta\ln(X_t^{W\$}) + b_{0,j}^{W\$}\Delta\ln(X_{t-1}^{W\$}) + b_{0,j}^{lc/\$}\Delta\ln(X_t^{lc/\$}) + b_{0,j}^{lc/\$}\Delta\ln(X_{t-1}^{lc/\$}) + c_{1,j}\Delta\ln(R_{j,t-1}^{lc}) + d_j\left[\ln(R_{j,t-1}^{lc}) - k_j - m_j^r\ln(C_{t-1}^r) - m_j^{W\$}\ln(X_{t-1}^{W\$}) - m_j^{lc/\$}\ln\left(X_{t-1}^{lc/\$}\right)\right] + \varepsilon_{j,t}$

[9] We have also estimated the two separate models, using just one exchange rate index in each model but the results were not satisfactory.

Variables	Observations	Mean	Standard Deviation	Min	Max
ln(GasNetPrice)	22,645	6.038	0.416	4.908	6.758
ln(GasNetPrLC)	22,645	7.345	1.838	4.536	13.60
ln(Brent)	31,813	3.704	0.746	2.245	4.949
ln(BrentR)	30,218	-0.927	0.778	-2.488	0.386
ln(DolrTWXin)	30,218	4.681	0.0906	4.489	4.869
ln(LCtoUSD)	22,622	1.091	1.978	-1.241	7.746

Notes: GasNetPrice, is the net retail price of gasoline, GasNetPrLC, is the net retail price of gasoline in local currency, Brent is the Brent crude oil price, BrentR is the Brent crude oil price in trade-weighted real dollars, DolrTWXin is the trade-weighted dollar exchange rate index, LCtoUSD denotes the units of local currency to USD dollar. All variables are expressed in natural logarithms.

4. Econometric framework

In this section, we describe the baseline linear one step error correction model (symmetric and asymmetric) that will be contrasted with the pooled panel GMM threshold model developed by Seo and Shin (2016) that accounts for the inclusion of endogenous regressors. In order to check for the validity of the threshold model we first used three alternative specifications: a) The Threshold Error Correction Model (TR), which follows the methodology of Hansen (1999; 2000) in an error correction framework, b) The Structural Threshold Error Correction Model (STR), described in Kourtellos *et al* (2016) and c) The Semiparametric Structural Threshold Error Correction Model (SMSTR), developed by Kourtellos *et al* (2017).

4.1. The Baseline Linear Model

The base model follows the estimation approach in Deltas (2008). We estimate first symmetric and asymmetric error correction models (ECMs) at the country level. The basic symmetric error correction model is of the following form:

$$\Delta\ln(R_{j,t}^{lc}) = a_j + \sum_{l=0}^{L} b_{l,j}\Delta\ln(C_{t-l}^{r}) + \sum_{l=1}^{L} c_{l,j}\,\Delta\ln(R_{t-l}^{lc}) + \sum_{l=0}^{L} d_{l,j}\Delta\ln(X_{t-l}^{W\$}) +$$

$$\sum_{l=0}^{L} e_{l,j}\Delta\ln(X_{t-l}^{lc/\$}) + z_j\left[\ln(R_{j,t-1}) - k_j - m_j\ln(C_{t-1})\right] + \varepsilon_{j,t} \quad (1)$$

where $R_{j,t}^{lc}$ is the retail price of gasoline in country j and week t in local currency, C_t^r is the price of crude oil (common to every country) in trade-weighted real dollars (the price in dollars divided by the trade-weighted dollar index), $X_t^{W\$}$ is the trade-weighted dollar exchange rate index, $X_t^{lc/\$}$ is the exchange rate of local currency units per dollar, $R_{j,t}$ is the retail price of gasoline in country j and week t in Euros, $C_{j,t}$ is the price of crude oil (common for every country) in dollars. The dependent variable $\Delta\ln(R_{j,t}^{lc})$ denotes the change in the log retail price in local currency from week t-1 to week t in country j and similarly for other difference terms. Note that in our models, all prices are in natural logarithms and coefficient estimates denote elasticities since there is no other meaningful way to jointly estimate the models involving series from different countries in different units. [10]

When estimating this regression in one step, the error correction term is multiplied out yielding the linear regression of the form:

$$\Delta\ln(R_{j,t}^{lc}) = a_j - k_j z_j + \sum_{l=0}^{L} b_{l,j}\Delta\ln(C_{t-l}^{r}) + \sum_{l=1}^{L} c_{l,j}\,\Delta\ln(R_{t-l}^{lc}) + \sum_{l=0}^{L} d_{l,j}\Delta\ln(X_{t-l}^{W\$}) +$$

$$\sum_{l=0}^{L} e_{l,j}\Delta\ln(X_{t-l}^{lc/\$}) + z_j\ln(R_{j,t-1}) - z_j m_j\ln(C_{t-1}) + \varepsilon_{j,t} \quad (2)$$

It is worth mentioning that the regression constant is a composite term each component of which is not separately identified in the one-step regression. However,

[10] We also used the US dollar price index with the consumption-weighted values being a robustness check. However, the empirical results did not pose any significant differences.

this is not important for assessing the price dynamics or for performing simulations of the retail price response to upstream price changes.[11]

4.2. The Threshold Model

We use the novel pooled panel GMM threshold method of Sheo and Shin (2016). More specifically, they study a dynamic threshold panel data model, which allows both regressors and threshold effect to be endogenous. Seo and Shin (2016) propose first-difference GMM (FD-GMM) and two-step least squares estimators and derive their limiting behaviors based on Hansen's asymptotic framework (Kourtellos et al, 2017). In order to check for the presence of a threshold effect, they rely on bootstrap-based testing procedure.

One could also resort alternatively to a semiparametric specification using local smoothers or splines/series to capture possible turning points. However such methods involve bandwidth choices and they do not lend themselves to estimating sharp turning points/thresholds as it is the case in the threshold model that we adopt in a fully interactive way (Polemis and Stengos, 2017; Kourtelos et al, 2016). Moreover, one important advantage of this methodology is that it avoids the ad hoc, subjective pre-selection of threshold values which has been a major critique of previous studies (Christie, 2014). In contrast to a simple case where the sample is split according to a known pre-assigned threshold value, the method that we use first tests for the presence of such a threshold and then estimates it (see for example Hansen, 2000; Caner and

[11] The basic symmetric ECM (see Equation 1) can also be estimated in two steps. In order to check the validity of the results, we also ran the other way and found similar results. Due to space competition the results are available upon request.

Hansen, 2004 and Kourtellos *et al,* 2016). In principle, one can test for additional sample splits, something that we did and we were able to detect.

Based on the above, Equation (1) can be cast in terms of threshold regression model that can be expressed as follows:

$$\Delta ln(R_t^{lc}) = a_i + \beta_1^T \Omega_t + v_t + \varepsilon_t, X_t^{W\$} \leq \gamma \tag{3}$$

$$\Delta ln(R_t^{lc}) = a_i + \beta_2^T \Omega_t + v_t + \varepsilon_t, X_t^{W\$} > \gamma \tag{4}$$

where we suppress the country index j and only use time as subscript. $X_t^{W\$}$ is the threshold variable, γ is the threshold level and Ω_t is a $d_x \times 1$ vector expressed in first differences containing all the regressors of the model in a compact form, including also all the lags (C_t, R_{t-1}^{lc}, $X_t^{W\$}$ and $X_t^{lc/\$}$), while β_1 and β_2 are regime specific coefficients. Moreover, a_i is the country fixed effect that control for differences across the cross-section element (i.e taxation level, demand and supply characteristics, gasoline market structure, etc), capturing individual heterogeneity. We also include the relevant year (time) fixed effect (v_t) which captures the co movement of the series due to external shocks (Polemis and Stengos, 2017). Finally ε_t denotes the idiosyncratic i.i.d error term.

For concreteness, the above two equations can be integrated into one as follows:

$$\Delta ln(R_t^{lc}) = a_i + \beta_2^T \Omega_t + v_t + \delta^T \Omega_t I(q_t \leq \gamma) + \varepsilon_t \tag{5}$$

where $\delta = \beta_1 - \beta_2$, q_t represents the scalar endogenous threshold variable ($X_t^{W\$}$) that splits the sample into two different groups (low and high regime). *I* (.) is the indication function denoting the regime defined by the threshold variable and the threshold level γ (sample split value). The indication function takes the value one when the condition in the parenthesis is satisfied and zero otherwise.[12]

We estimate Equation (5) using the novel GMM method of Seo and Shin (2016) as fully described in Asimakopoulos and Karavias (2015). The latter which uses Arellano and Bond (1991) type instruments is more advanced than other threshold methods such as Hansen (1999) and Kremer et al., (2013). This is attributed to the fact that it allows for endogeneity in both the regressors and the threshold variable (Sheo and Shin, 2016). The potential endogeneity problem is associated with exchange rate fluctuations in asymmetric gasoline pricing mechanism. While there remains debate in the literature whether fluctuations in the exchange rate drives asymmetric gasoline pricing mechanism or gasoline price asymmetry drives exchange rate volatility, the fact is that the potential for endogeneity exists. As a consequence this model fully incorporates this issue by allowing the exchange rate factor variable (trade-weighted dollar exchange rate index) to be endogenously determined.

5. Results and discussion

This section presents the results of the threshold models along with the benchmark linear specifications (symmetric and asymmetric). In addition, we offer a comparative discussion between the threshold effects and the static panel fixed effects linear specification benchmark models, while we firstly check for the existence of cross-section dependency and stationarity properties of our sample variables by using "*second generation*" tests for unit roots.

[12] The choice of lag length p = 2 is chosen by Akaike's selection Information Criterion (AIC).

5.1 Testing for cross-section dependence

One of the additional complications that arise when dealing with panel data compared to the pure time-series case, is the possibility that the variables or the random disturbances are correlated across the panel dimension. The early literature on unit root and cointegration tests adopted the assumption of no cross-sectional dependence. However, it is common for macro-level data to violate this assumption which will result in low power and size distortions of tests that assume cross-section independence (Polemis and Stengos, 2017). We use the cross-section dependence test proposed by Pesaran (2004). The test is based on the estimation of the linear panel model of the form:

$$y_{it} = \alpha_i + \beta_i' x_{it} + u_{it}, \quad i = 1,..N; T = 1,..T \tag{13}$$

where T and N are the time and panel dimensions respectively, α_i the provincial-specific intercept, and x_{it} a $kx1$ vector of regressors, and u_{it} the random disturbance term. The null hypothesis in both tests assumes the existence of cross-section correlation: $Cov(u_{it}, u_{jt}) = 0$ for all t and for all $i \neq j$. This is tested against the alternative hypothesis that $Cov(u_{it}, u_{jt}) \neq 0$ for at least one pair of i and j. The Pesaran (2004) test is a type of Lagrange-Multiplier test that is based on the errors obtained from estimating Equation 13 by the OLS method. If the relevant test strongly rejects the null hypothesis of cross-section independence for all the models then we proceed to test for unit roots using tests that are robust to cross-section dependence (the so-called "s*econd generation*" tests for unit roots in panel data). We carry out the first part of the empirical analysis by examining the presence of cross-section dependence. We use the cross-section dependence test (CD test) proposed by Pesaran (2004).

As it is evident from Table 2 the relevant test strongly rejects the null hypothesis (p-value = 0.000) of cross-section independence for all the variables. In light of this evidence we proceed to test for unit roots using tests that are robust to cross-section dependence.

Table 2: Cross-section dependence test

Variable	**CD test**	**P-value**	**Correlation**	**Absolute (correlation)**
ln(GasNetPrice)	459.72***	0.000	0.963	0.963
ln(GasNetPrLC)	194.70***	0.000	0.456	0.780
ln(Brent)	643.95***	0.000	1.000	1.000
ln(BrentR)	627.60***	0.000	1.000	1.000
ln(DolrTWXin)	627.60***	0.000	1.000	1.000
ln(LCtoUSD)	253.12***	0.000	0.539	0.645

***Notes**:* Under the null hypothesis of cross-sectional independence the CD statistic is distributed as a two-tailed standard normal. Results are based on the test of Pesaran (2004). The p-values are for a one-sided test based on the normal distribution. Correlation and Absolute (correlation) are the average (absolute) value of the off-diagonal elements of the cross-sectional correlation matrix of residuals. GasNetPrice, is the net retail price of gasoline in Euros, GasNetPrLC, is the net retail price of gasoline in local currency, Brent is the Brent crude oil price in USD, BrentR is the Brent crude oil price in trade-weighted real dollars, DolrTWXin is the trade-weighted dollar exchange rate index, LCtoUSD denotes the units of local currency to USD dollar. All variables are expressed in natural logarithms. Significant at ***1% level of statistical significance.

5.2 Unit root and cointegration testing

To examine the stationarity properties of the variables in our models we use the second generation unit root tests for panel-data proposed by Breitung and Das (2005) and Pesaran (2007). The test results suggest that all the sample variables are integrated of order one (I-1).[13]

In order to investigate whether a long-run equilibrium relationship exists among the variables in our models we implement two cointegration tests proposed by Westerlund (2007) that allow for cross-section dependence. The results of the tests are presented in the following table; the critical values were created using a bootstrapping method. The results indicate that the first test rejects the null hypothesis of no cointegration for all three models. However, in some cases the second test that restricts the intercept to be the same across all countries fails to reject the null.[14]

[13] Due to space limitation the results of the unit root testing are available from the authors on request.
[14] The results though are sensitive to the selection of the lag structure of the model. Persyn and Westerlund (2008) point out that this sensitivity might occur in small datasets.

Table 3: Westerlund ECM panel cointegration tests

Equation	Statistic			
	G_τ	**G_α**	**P_τ**	**P_α**
lGasNetPrice = f (lBrent)	-4.955*** (0.000)	-45.127*** (0.000)	-26.008*** (0.000)	-43.164*** (0.000)
lGasNetPrLC = f (lBrentR)	-2.870*** (0.000)	-17.690*** (0.000)	-8.747 (0.998)	-6.020 (0.995)
lGasNetPrLC = f (lDolrTWXin)	-2.755*** (0.004)	-14.900*** (0.000)	-10.295 (0.849)	-7.307 (0.928)
lGasNetPrLC = f (lLCtoUSD)	-2.432 (0.309)	-16.267*** (0.000)	-13.922*** (0.001)	-15.702*** (0.000)

Notes: The test regression was fitted with a constant and trend and four lags and leads. The kernel bandwidth was set according to the rule $4(T/100)^{2/9}$. The null hypothesis assumes that there is no co-integration. The numbers in parentheses denote the p-values. GasNetPrice, is the net retail price of gasoline in Euros, GasNetPrLC, is the net retail price of gasoline in local currency, Brent is the Brent crude oil price in USD, BrentR is the Brent crude oil price in trade-weighted real dollars, DolrTWXin is the trade-weighted dollar exchange rate index, LCtoUSD denotes the units of local currency to USD dollar. All variables are expressed in natural logarithms. Significant at ***1% level of statistical significance.

5.3 Empirical results

We first proceed with the exposition of results generated from the benchmark linear specifications that will be contrasted with the threshold model. In this way, we will be able to draw the differences between these results and the traditional benchmark linear specifications in order to focus on issues that were depicted in the threshold model and are different from the linear baseline one (Polemis and Stengos, 2017).

From the following table, it is evident that nearly all of the variables are statistically significant in nearly all either of the specifications. However, the relevant signs of most of the regressors entering the linear models (symmetric and asymmetric ones) differ drastically revealing that the results are not robust. Specifically, examining the linear asymmetric model (see columns 3-6), it is evident that the crude oil positive coefficients are larger than their negative counterparts, indicating that the effects of upstream price increases are larger than those of price decreases. The relevant estimates for the positive coefficients range from 0.29 to 0.36, compared to 0.28 and 0.27 for the negative ones respectively. This means that a 10% increase (decrease) of the crude oil price will lead on average to a short-run increase (decrease) of the net retail gasoline price equal to 3.25% and 2.75% respectively.

Regarding the exchange rate terms included in the baseline linear model, some interesting results emerge. Specifically, the first exchange rate term ($\Delta \ln X^{W}$) representing the real effective exchange rate effect provides mixed results since the estimated coefficients when significant alternate their signs (see columns 3-6), revealing an inconsistent behaviour. On the contrary, the second exchange rate term ($\Delta \ln X^{lc}$) representing the nominal effective Euro trade-weighted exchange rate effect is positively correlated with the retail gasoline price in all of the specifications of the

ECMs. The relevant estimates for the positive coefficients are larger than their negative counterparts ranging from 0.49 to 0.51, compared to 0.33 and 0.32 for the negative ones respectively. Surprisingly the cointegation-terms (lagged crude oil and retail price) denoting the long-run relationship between the net retail gasoline price and its crude oil marker (Brent crude oil price or New York spot gasoline price) are not statistically significant (see columns 2-4). The same finding applies to the two error correction terms (see columns 5-6) representing the speed of adjustment toward the long-run equilibrium. All in all, the empirical findings suggest the absence of short-run and long-run price asymmetry.

Next we apply the necessary linearity tests of the benchmark linear specifications against the non-parametric alternative ones given in the threshold model. The tests we use are based on bootstrap critical values of a Wald type heteroskedasticity-consistent test of the null hypothesis against a TR alternative. Specifically all the bootstrapped tests reject linearity in favour of the threshold model with p-values equal to 0.000 in all cases. As a consequence and in alignment with the aforementioned results, the baseline model does not capture the nonlinear effects of the ERPT mechanism.

Therefore, we proceed to estimate the threshold model. As it is evident from the inspection of Table 4, we find that the optimal threshold level of the ERPT proxied by the trade-weighted dollar exchange rate index is almost identical in all of the four models (4.65).

Table 4: Baseline and threshold model results

	(1)	(2)	
Method	**OLS – Baseline Model**	**GMM – Threshold Model**	
Threshold	**-**	***4.6232***	
Regimes	**-**	**Low**	**High**
Constant	0.0000	0.0005	0.0015
	(0.9625)	(0.9804)	(0.6094)
$\Delta\ln(C_t^r)$	0.2914***	0.4612	0.4132***
	(0.0000)	(0.1999)	(0.0000)
$\Delta\ln(C_{t-1}^r)$	0.1618***	0.1425	-0.0337
	(0.0000)	(0.6379)	(0.4064)
$\Delta\ln(C_{t-2}^r)$	0.1482***	0.4906**	0.5583***
	(0.0000)	(0.0218)	(0.0000)
$\Delta\ln(X_t^W)$	-0.1145	2.3203	2.6203***
	(0.3663)	(0.1947)	(0.0000)
$\Delta\ln(X_{t-1}^W)$	0.0207	-0.2669	-3.3567***
	(0.8677)	(0.9262)	(0.0000)
$\Delta\ln(X_{t-2}^W)$	0.401***	1.4322	2.249***
	(0.0016)	(0.101)	(0.0000)
$\Delta\ln(X_t^{lc})$	0.4417***	-1.4858	0.0982
	(0.0000)	(0.1605)	(0.6484)
$\Delta\ln(X_{t-1}^{lc})$	0.3468***	-0.0092	1.4611***
	(0.0000)	(0.995)	(0.0000)
$\Delta\ln(X_{t-2}^{lc})$	0.1690***	-0.2144	−0.3303**
	(0.0004)	(0.7493)	(0.0349)
$\Delta\ln(R_{t-1}^{lc})$	−0.0952***	-1.0649***	0.1018*
	(0.0018)	(0.0000)	(0.0778)
$\Delta\ln(R_{t-2}^{lc})$	−0.0466**	-0.0601	0.0166
	(0.033)	(0.5902)	(0.7411)
$\ln(R_{t-1})$	-0.0001	-0.0004	-0.0002

	(0.5907)	(0.7487)	(0.6258)
$\ln(C_{t-1})$	0.0098*	-0.0560	-0.0230***
	(0.0647)	(0.7330)	(0.0060)
Adjusted R^2	0.347	-	
J Statistic	-	1.512	
D-W P-Value	0.8439	0.1380	
SupWald Statistic	-	46.3847	
SupWald Boot P-Value	-	0.0041***	
Observations	22,645	22,645	

Notes: *Column one refers to the pooled panel OLS results of the symmetric model (baseline). Column two presents the estimations of the dynamic GMM of Seo and Shin (2016). The threshold variable is the trade-weighted dollar exchange rate index,* $X_{W,t}$. *All variables are instrumented with its lag terms.* R_t^{lc} *is the net retail price of gasoline in local currency,* C_t^r *is the price of crude oil in trade-weighted real dollars,* $X_t^{W\$}$ *is the trade-weighted dollar exchange rate index,* $X_t^{lc/\$}$ *is the exchange rate of local currency units per dollar,* $R_{j,t}$ *is the net retail price of gasoline in Euros,* $C_{j,t}$ *is the price of crude oil in dollars,* $\Delta ln(R_{j,t}^{lc})$ *is the change in the log retail price in local currency from week t-1 to week t in country j and similarly for other difference terms. D-W denotes the Durbin-Watson test for autocorrelation in panel data. All models include time and country fixed effects.* *** *Significance at 1%* ** *Significance at 5%* * *Significance at 10%*

However, there is a prevailing issue of endogeneity. The latter is associated with the use of the exchange rate term which is treated as an endogenous covariate in our models. This could be explained by the fact that although it has been documented in the literature that exchange rate affects the level of retail gasoline prices (see among others Galeotti, 2003; Polemis, 2012; Polemis and Fotis, 2013) there is a possibility that the direction of causality might also be reversed. Moreover, it is almost certainly the case that ERPT and upstream pricing adjustment mechanism are not randomly determined among the EU-28 countries throughout the sample period, thus raising the concern that the coefficients of exchange rate and crude oil marker (Brent or New York spot gasoline prices) are biased.

To provide a credible identification strategy that would address this issue and allow interpreting the results in a causal way we followed two approaches. Firstly, we perform the necessary tests to detect endogeneity in the threshold model. The following table depicts the endogeneity test results (see Kourtellos et al, 2017). It is worth mentioning that, the proposed test for the endogeneity of the threshold variable (X_t^W), is valid regardless of whether the threshold effect is zero or not. Moreover, the test statistic is applicable regardless of whether the regressors are endogenous or exogenous. Under the null hypothesis, X_t^W is exogenous, while under the alternative hypothesis the threshold variable is endogenous. As it is evident from Table 5, the two bootstrap test statistics (White and Homo) reject the null hypothesis. This means that the threshold variable (trade-weighted dollar exchange rate index) is treated as endogenous in our TR model.

Table 5: Threshold endogeneous test results

Polynomial	**Wald (*White*)**	**Wald (*Homo*)**	**Boot P (White)**	**BootP (Homo)**	**GCV**
0	15.0911	20.2219	0.0000^{***}	0.0000^{***}	0.00039
1	21.4792	22.1012	0.0000^{***}	0.0101^{***}	0.000379
2	24.8343	24.0289	0.0202^{**}	0.0606^{*}	0.000377

Notes: This table presents the endogeneous tests results suggested by Kourtellos et al. (2016) at varying polynomials. Boot (White) and Boot (Homo) are corresponding bootstrap critical values at 5% significant level. *** Significance at 1%, ** Significance at 5%, * Significance at 10%

In the second stage and after having identified that the threshold variable is endogenous, we rely on the GMM model developed by Seo and Shin (2016).[15] As a consequence this may lead to biased results. Specifically, the main variable of interest is the trade-weighted dollar exchange rate index. Recall, that when entered linearly to the asymmetric model, the coefficients alternated their signs giving an indication of an inconsistent behaviour (see Table 4 column 1). On the other hand, the results for the

[15] We have also used three other panel threshold models namely Threshold Error Correction Model along the lines of Hansen (1999), Structural Threshold Error Correction Model developed by Kourtellos et al, (2016) and Semiparametric Structural Threshold Error Correction Model described in Kourtellos et al. (2017). However, they did not perform well since an (endogenous) threshold variable and endogenous regressors co-exist in the model. Therefore, the analysis relies solely on the GMM model. The results of these models are available upon request.

non-linear model with an endogenous threshold, do suggest a strong non-linear relationship between retail gasoline prices and exchange rate. The point estimates suggest that the level of real effective exchange rate is positively related to the level of net retail gasoline price. However, it is evident that the trade-weighted dollar exchange rate index is more important in the sample above the threshold (high regime) since the relevant coefficient (2.6203) is statistically significant. This means that a 10% increase (decrease) in the level of exchange rate leads to a 26.2% increase (decrease) in the retail gasoline price in the short-run. This finding gives sufficient evidence that for net EU exporting countries (high regime), fluctuations in the real effective exchange rate of the US against its major EU trading partners does affect the level of net retail gasoline prices and subsequently the asymmetric pricing mechanism. It is also worth mentioning that the magnitude of the relevant elasticity exceeds unity denoting that ERPT is almost complete. This finding runs contrary to the existing studies where the relevant estimated elasticity ranges from 0.4 to 0.6 (see for example Krugman, 1986; Helpman and Krugman, 1987; Feenstra, 1989; Goldberg and Knetter, 1997).

Notably, the other control variables have the expected signs and are all statistically significant for values above the threshold (high regime). Similarly to the linear model, the upstream oil price marker (Brent crude oil price) is positively correlated with the net retail gasoline price as it was expected. The relevant short-run price elasticity is estimated to 0.413. This means that a 10% increase (decrease) of the Brent crude oil price will lead to a short-run increase (decrease) of the net retail gasoline price equal to 4.13%. This pattern does not change since the input price coefficient remains statistically significant even when the number of lags is set to two (0.5583). Regarding, the second exchange rate term for the net exporting countries (high regime), we argue that the relevant coefficients are statically significant alternating their signs only when one and two lags are present (1.4611 and -0.3303 respectively). Surprisingly the lagged retail price cointegation term (lnR_{t-1}) is not statistically significant bellow and above the threshold.

Having estimated the GMM we proceed to capture possible asymmetries that arise from differential responses of net retail gasoline price changes to positive and negative fluctuations in the exchange rate. The test we use is based on bootstrap critical values of a Wald type heteroskedasticity-consistent test of the null hypothesis (no asymmetry) against the existence of an asymmetric gasoline adjustment mechanism (see for example Hansen, 1996; Godby *et al*, 2000; Li *et al*, 2002). In other words rejection of the null hypothesis implies that there is no significant threshold (no asymmetry). From the relevant table, we find that the null hypothesis is strongly rejected with a SupWald Bootstrapped P-value for the GMM equal to 0.0041. In this case, we can safely argue that gasoline asymmetry is present in the EU oil industry. These results are in alignment with some of the empirical studies reported in the literature (see for example Borenstein et al, 1997; Deltas, 2008; Polemis, 2012; Greenwood-Nimmo and Shin, 2013; Kristoufek and Lunackova, 2015; Polemis and Tsionas, 2017). One possible reason for this behaviour might be attributed to the fact that in such a case, the profit function is inherently asymmetric. If prices are too high, the costs to profit of a sub-optimal level of sales is partly offset by the higher price (and hence profit margin) of each unit sold. But if prices are too low, beyond some point the firm will be selling more units, and each of them at a loss, so that the quantity and price effects on profits reinforce rather than offset each other.

Lastly, all underlying estimated equations pass a battery of diagnostic tests. Specifically, the reported J-statistic test indicates that the instrument list satisfies the orthogonallity conditions in all of the specifications, since the null hypothesis that the over-identifying restrictions are valid cannot be rejected. Similarly, our estimated TR model does not suffer from autocorrelation since the relevant test (D-W test) cannot reject the null hypothesis.

6. Concluding remarks

This paper provides new insights into "*rockets and feathers*" hypothesis since it tries to investigate the impact of ERPT on asymmetric gasoline pricing mechanism. For this reason we use a large weekly panel of EU-28 countries over the period January 1994 to January 2015. Our pooled panel GMM threshold model follows the spirit of Seo and Shin (2016) and allows for the existence of a threshold effect with endogenous regressors.

In this study we use a bootstrap procedure to test the null hypothesis of a linear (symmetric) formulation against a TR alternative. Moreover, we provide a direct test for asymmetric behaviour around the estimated threshold. The results of the baseline model (expressed in symmetric and asymmetric formulation) compared with the threshold effects model that we use in the present study reveal significant differences in the interpretation of the key variable of interest (real effective exchange rate). This means that the baseline model does not capture the nonlinear effects stemmed from the existence of a threshold according to the bootstrapped P-values of the relevant linearity tests. As a consequence, the threshold model is better suited to assess these effects on gasoline price mechanism under two different regimes of ERPT (appreciation and depreciation).

The empirical findings reveal that the threshold variable expressed by the trade-weighted dollar exchange rate index is statistically significant only in the sample above the threshold (high regime). This means that for the net EU exporting countries, fluctuations in the real effective exchange rate of the US against its major EU trading partners does affect the level of pre-tax retail gasoline prices with the relevant elasticity exceeding unity (complete ERPT). Moreover, all the relevant statistical tests reject the null hypothesis that there is no significant threshold and thus an asymmetric adjustment gasoline mechanism prevails. Lastly, the results are rather robust when we account for the inclusion of the final (pump) retail gasoline price.

References

Aguerre, R.B., A.M. Fuertes and Phylaktis, K (2012). Exchange rate pass-through into import prices revisited: What drives it? *Journal of International Money and Finance*, 31 (4): 818-844

Akçelik, F., Ogünç F. (2016) Pass-through of crude oil prices at different stages in Turkey. *Central Bank Review* 16, 41-51.

Al-Abri, A.S., Goodwin, B.K (2009) Re-examining the exchange rate pass-through into import prices using non-linear estimation techniques: Threshold cointegration *International Review of Economics and Finance* 18, 142–161.

Aleema, A., Lahiani, A (2014) A threshold vector autoregression model ofexchange rate pass-through in Mexico, *Research in International Business and Finance* 30, 24– 33

Arellano, M. and Bond, S., (1991). Some Tests of Specification for Panel Data: Monte Carlo Evidence and an Application to Employment Equations, *Review of Economic Studies*, 582: 277-297.

Asimakopoulos, S., Karavias, Y., 2015. The impact of government size on economic growth: A threshold analysis. *Economics Letters*, doi: http://dx.doi.org/10.1016/j.econlet.2015.12.010

Auer, P., Schoenle P (2016) Market structure and exchange rate pass-through. *Journal of International Economics*, 98, 60-77.

Bagnai, A., Alexander C., Ospina, M. (2018) Asymmetries, outliers and structural stability in the US gasoline market. *Energy Economics*, 69, 250-260,

Bailliu, J. and E. Fujii (2004), Exchange Rate Pass-Through and the Inflation Environment in Industrialised Countries: an Empirical Investigation, Bank of Canada Working Paper Series, 04-21.

Bhattacharya, P.S., Karayalcin, C.A., Thomakos, D (2008) Exchange rate pass-through and relative prices: An industry-level empirical investigation. *Journal of International Money and Finance* 27, 1135–1160

Blair, B.F., Campbell, R.C., Mixon, P.A (2017) Price pass-through in US gasoline markets. *Energy Economics,* 65, 42-49.

Borenstein, S., Shepard, A. (1996) Dynamic pricing in retail gasoline markets. *RAND Journal of Economics* 27, 429-451.

Borenstein, S., Cameron, C.A., Gilbert, R. (1997) Do gasoline prices respond asymmetrically to crude oil price changes? *Quarterly Journal of Economics* 112, 305-339.

Breitung, J., Das, S. (2005) Panel unit root tests under cross-sectional dependence. *Statistica Neerlandica* 59, 414-433.

Campa, J.M., Minquez G. (2006). Differences in exchange rate pass-through in the euro area. *European Economic Review* 50, 121-145.

Campa, J. M. and L. Goldberg (2006a), Distribution Margins, Imported Inputs, and the Sensitivity of the CPI to Exchange Rates, NBER Working Papers, 12121.

Campa, J. M. and L. Goldberg (2006b), Pass Through of Exchange Rates to Consumption Prices: What has Changed and Why?, NBER Working Papers, 12547.

Caner, M., and Hansen. B. (2004). Instrumental Variable Estimation of a Threshold Model, *Econometric Theory* 20: 813–843.

Ceglowski, J. (2010) Exchange rate pass-through to bilateral import prices. *Journal of International Money and Finance* 29, 1637–1651.

Choudhri E.U., Faruqee, H., Hakura, D.S (2005) Explaining the exchange rate pass-through in different prices. *Journal of International Economics* 65, 349– 374

Choudhri, E. U. and D. S. Hakura (2002), Exchange Rate Pass-Through to Domestic Prices: Does the Inflationary Environment Matter?, IMF Working Paper WP/01/194.

Gagnon, J. E. and J. Ihrig (2004), Monetary Policy and Exchange Rate Pass-Through, *International Journal of Finance and Economics*, 9 (4): 315-38.

Christie, T. (2014). The effect of government spending on economic growth: Testing the non-linear hypothesis. *Bulletin of Economic Research*, 66(2), 183-204.

Douglas C, and Herrera, A.M. (2010). Why are gasoline prices sticky? A test of alternative models of price adjustment, *Journal of Applied Econometrics*, 25(6): 903-928.

Deltas, G. (2008) Retail gasoline price dynamics and local market power. *Journal of Industrial Economics* 56, 613-628.

Demetriades, P., James, G. (2011) Finance and growth in Africa: The broken link. *Economics Letters* 113(3), 263-265.

Devereux, M., Yetman, J (2010) Price adjustment and exchange rate pass-through. *Journal of International Money and Finance* 29,181–200

Dornbusch, R., (1987). Exchange rates and prices. *American Economic Review* 77, 93–106.

Eckert, A. (2002). Retail price cycles and response asymmetry. *Canadian Journal of Economics* 35, 52-77.

Eleftheriou, K., Nijkamp, P and Polemis, M. (2018). Asymmetric price adjustments in US gasoline markets: Impacts of spatial dependence on the 'rockets and feathers' hypothesis. *Regional Studies* (forthcoming)

Fabra N., and Reguant, M. (2014). Pass-Through of Emissions Costs in Electricity Markets. *American Economic Review*, 104(9): 2872-289

Feenstra, R.C. (1989) Symmetric Pass-Through of Tariffs and Exchange Rates Under Imperfect Competition: An Empirical Test, *Journal of International Economics*. 27(1,2), 25-45.

Federal Reserve Bank of St. Louis - FRED (2017). Board of Governors of the Federal Reserve System (US), Trade Weighted U.S. Dollar Index: Broad [TWEXB], retrieved from https://fred.stlouisfed.org/series/TWEXB, September 13, 2017.

Gagnon, J. E. and Ihrig, J. (2004). Monetary policy and exchange rate pass-through. *International Journal of Finance and Economics*. 9: 315–338. doi:10.1002/ijfe.253

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

Godby, R., Lintner, A., Stengos, T., Wandschneider, B. (2000) Testing for asymmetric pricing in the Canadian retail gasoline market, *Energy Economics,* 22(3): 349-368.

Goldberg, P., Knetter, M. (1997) Goods Prices and Exchange Rates: What Have We Learned? *Journal of Economic Literature*, 35(3), 1243-1272.

Goodwin, B.K., Harper, D. (2000) Price transmission, threshold behavior, and asymmetric adjustment in the U.S. pork sector. *Journal of Agricultural and Applied Economics* 32(3), 543-553.

Goodwin, B.K., Holt, M.T. (1999) Price transmission and asymmetric adjustment in the U.S. beef sector. *American Journal of Agricultural Economics* 81(3), 630-637.

Gopinath, G., Itskhoki, O., Rigobon R. (2010) Currency Choice and Exchange Rate Pass-through. *American Economic Review* 100(1), 304-336.

Greenwood-Nimmo, M., Shin, Y. (2013) Taxation and the asymmetric adjustment of selected retail energy prices in the UK. *Economics Letters* 121, 411-416.

Hansen, B.E., 1996. Inference when a nuisance parameter is not identified under the null hypothesis. *Econometrica* 64, 413-430

Hansen, B.E. (2000). Sample Splitting and Threshold Estimation. *Econometrica*, 68 (3) 575-603.

Heckman, J., (1979), Sample Selection Bias as a Specification Error, *Econometrica* 47, 153–161.

Helpman, E., Krugman, P., (1987) Market Structure and Foreign Trade: Increasing Returns, Imperfect Competition, and the International Economy, Vol. 1 of MIT Press Books. The MIT Press.

Hooper, P., and Mann C.L. (1989) Exchange Rate Pass-Through in the 1980s: The Case of U.S. Imports of Manufactures, *Brookings Papers on Economic Activity.* 1, 297-337

Schröder, M., and Hüfner, F.P. (2002). Exchange rate pass-through to consumer prices: a European perspective, ZEW Discussion Papers 02-20, ZEW - Zentrum für Europäische Wirtschaftsforschung, Center for European Economic Research.

Johnson, R.N. (2002) Search costs, lags and prices at the pump. *Review of Industrial Organization* 20, 33-50.

Kourtellos, A., Stengos, T. Ming, T.C (2016). Structural Threshold Regression, Econometric Theory, 32(04): 827-860.

Kourtellos., A., Stengos, T., Sun, Y (2017). Endogeneity in Semiparametric Threshold Regression. Working Paper Series 17-13, The Rimini Centre for Economic Analysis.

Kreinin, M.E. (1977) The Effect of Ex-change Rate Changes on the Prices and Volume of Foreign Trade. *International Monetary Fund Staff Papers*, 24(2), 297- 329.

Kremer, M, Lee, J., Robinson,J Rostapshova. O., (2013). Behavioral Biases and Firm Behavior: Evidence from Kenyan Retail Shops. *American Economic Review: Papers and Proceedings* 103 (3): 362-368.

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

Krugman, P., (1986). Pricing to market when the exchange rate changes. NBER Working Papers 1926. National Bureau of Economic Research.

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

Mann. J. (2016). Rockets and feathers meet markup margins: Applications to the oil and gasoline industry. *Canadian Journal of Economics, 49*, 772-788. 10.1111/caje.12213

McCarthy, J., (2006). Pass-through of Exchange Rates and Import Prices to Domestic Inflation in Some Industrialized Economies. Federal Reserve Bank of New York.

Persyn, D., Westerlund, J (2008). Error-correction–based cointegration tests for panel data, *Stata Journal*, 8(2): 232-241.

Pesaran, M.H. (2004) General diagnostic tests for cross section dependence in panels. Cambridge Working Papers in Economics 0435, Faculty of Economics, University of Cambridge.

Pesaran, M. (2007) A simple panel unit root test in the presence of cross-section dependence. *Journal of Applied Econometrics* 22, 265-312.

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

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

Polemis., M., Tsionas, M. (2016). An alternative semiparametric approach to the modelling of asymmetric gasoline price adjustment. *Energy Economics*, 56(C), 384-388.

Polemis., M., Tsionas, M. (2017). Asymmetric Price Adjustment in the U.S Gasoline Industry: Evidence from Bayesian Threshold Dynamic Panel Data Models, *International Journal of the Economics of Business*, 24:1, 91-128,

Polemis, M., Stengos, T (2017). Does Competition Prevent Industrial Pollution? Evidence from a Panel Threshold Model, Working Paper Series 17-07, The Rimini Centre for Economic Analysis

Radchenko, S. (2005) Oil price volatility and the asymmetric response of gasoline prices to oil price increases and decreases. *Energy Economics* 27, 708-730.

Radchenko, S., Tsurumi, H. (2006). Limited information Bayesian analysis of a simultaneous equation with an autocorrelated error term and its application to the U.S. gasoline market. *Journal of Econometrics* 133(1), 31-49.

Radchenko, S., Shapiro, D. (2011). Anticipated and unanticipated effects of crude oil prices and gasoline inventory changes on gasoline prices. *Energy Economics* 33(5), 758-769.

Seo, M.H. and Shin, Y., 2016. Dynamic Panels with Threshold Effect and Endogeneity. *Journal of Econometrics* 195 (2), 169-186.

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

Westerlund, J., 2007. Testing for error correction in panel data. *Oxford Bulletin of Economics and Statistics* 69, 709-748.

Wlazlowski, S., Giulietti, M., Binner, J., Milas, C. (2012) Price transmission in the EU wholesale petroleum markets. *Journal of Business and Economic Statistics* 30 (2), 165-172.

Wooldridge, J.M. (2002) *Econometric Analysis of Cross Section and Panel Data*. MIT Press, Cambridge, MA.

Woo, W (1984). Exchange Rates and the Prices of Nonfood, Nonfuel Products, *Brookings Papers of Economic Activity*, 2, 511-530.

Hahn, E. (2003), Pass-Through of External Shocks to Euro Area Inflation, ECB Working Paper Series, 243.

Faruqee, H. (2006), Exchange Rate Pass-Through in the Euro Area, IMF Staff Papers, 53(1): 63-88.

Yanagisawa (2012) Structure for pass-through of oil price to gasoline price in Japan. *IEEJ Energy Journal*, 7(3).

Yang, J., (1997). Exchange rate pass-through in U.S. manufacturing industries. *The Review of Economics and Statistics* 79, 95–104.