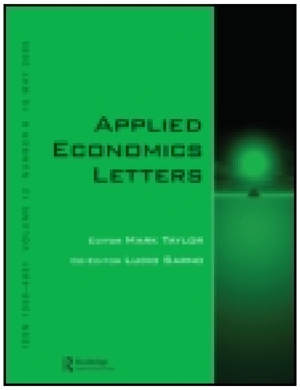

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Asymmetric gasoline price responses in France

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Asymmetric gasoline price responses in France

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This article examines the response of gasoline prices in France to shocks to crude oil prices in the international market. Using the Autoregressive Distributed Lag (ARDL) bounds testing approach of cointegration, we investigate potential price asymmetries in the French diesel and premium gasoline markets using weekly data over the period from May 1990 to April 2011. We find that gasoline prices gradually adjust towards a long-term equilibrium after a shock to the crude oil price, but this adjustment is lower when the crude oil price decreases than when it increases. We also find that the estimates differ slightly depending on the chosen gasoline price.

Keywords: gasoline prices; price asymmetries; ARDL cointegration

JEL Classification: Q41; Q43

I. Introduction

The last decade has been characterized by a 'breathtaking ascent' and increased volatility of crude oil prices (Smith, 2009, p. 145). These changes in gasoline prices have significantly affected the budgets of consumers and have increased the pressure of public opinion and political authorities on the actors of the oil industry. Producers, refiners, retailers and, frequently, multinational oil companies are regularly accused of using crude oil price changes to unreasonably increase their margins, especially in times of crisis, when the purchasing power of households is a critical issue. The frequent accusation is that, following an increase in the crude oil price, the oil industry rapidly adjusts gasoline prices upwards, but following a decrease in the crude oil price, the industry slowly adjusts prices downwards.

In this article, we address the following question: how do gasoline prices respond to shocks to the crude oil price? Numerous studies have analysed the relationship between the price of crude oil and the price of gasoline since the seminal work of Bacon (1991). A recent survey by Frey and Manera (2007) identified 34 empirical studies since 1991. Most of these studies find an asymmetric response of gasoline prices, but some articles challenge this result (Rao and Rao, 2008; Silva *et al.*, 2013) and thus continue the debate. This lack of consensus among economists may be explained by differences in methodologies, model data, periods and countries of interest; nonetheless, the results are not unanimous, and the question remains unresolved.

The literature review also suggests that there is a lack of empirical studies concerning the transmission of changes in the price of crude oil to the prices of gasoline in France. To the best of our knowledge, only one article focuses on France (Audenis *et al.*, 2002), and three articles consider the issue in France and other countries (Balabanoff, 1993; Galeotti *et al.*, 2003; Grasso and Manera, 2007). These studies all find an asymmetric response of gasoline prices to shocks to crude oil prices and also find that prices rise faster than they fall. However, all of the previous empirical studies pertaining to France have used monthly data prior to 2000. Therefore, we do not know the response of gasoline prices to the shocks to crude oil prices in the last decade and whether the previous results hold for weekly data.

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In this article, we investigate potential price asymmetries in the French diesel and premium gasoline markets. This article contributes to filling a gap in the recent literature by using updated and higher frequency data and the Autoregressive Distributed Lag (ARDL) bounds testing approach of cointegration.

II. Data and Methodology

This study uses weekly data for the period from 4 May 1990 to 29 April 2011, which yields a sample of 1096 observations. We considered two types of gasoline price: diesel (P_D) and premium (P_{SP95}). These data were collected from the French Ministry of Ecology, and they allow us to identify tax values for each type of gasoline (T_D and T_{SP95}). The chosen crude oil price was the Brent price, and the data were obtained from the US Energy Information Administration. We considered the Brent price in euro currency (B_E) and in dollar currency (B_D) to account for any possible effect of the exchange rate (R). All variables are in logarithm terms.

We examined the relationships between gasoline prices, crude oil prices, oil taxes and exchange rate. To empirically analyse these relationships, we used the ARDL bounds testing approach of cointegration developed by Pesaran *et al.* (2001). The ARDL approach has numerous advantages over standard cointegration methods. First, this approach is not restrictive because it is applicable irrespective of whether the underlying regressors are purely I(0), purely I(1) or fractionally integrated. Second, the ARDL approach allows the variables to exhibit different optimal lags. Third, compared with other multivariate cointegration techniques, the ARDL approach employs a single reduced form equation. Finally, it is relatively more efficient to determine the cointegration relation in small samples.

The bound testing procedure consists of estimating an unrestricted error correction model (in which c is a constant and ε is the error term) by Ordinary Least Squares (OLS) as follows:

$$\begin{aligned}\Delta P_i = {} & c_i + \alpha_{1i}P_{i,t-1} + \alpha_{2i}B_{j,t-1} + \alpha_{3i}T_{i,t-1} \\ & + \alpha_{4i}R_{i,t-1} + \sum_{k=1}^{p}\beta_{k,i}\Delta P_{i,t-k} + \sum_{k=1}^{l}\lambda_{k,i}\Delta B_{j,t-k} \\ & + \sum_{k=1}^{m}\tau_{k,i}\Delta T_{i,t-k} + \sum_{k=1}^{n}\rho_{i,k}\Delta R_{t-k} + \varepsilon_{it} \\ & \text{for } i = \{D, \mathrm{SP95}\} \text{ and } j = \{E, D\}\end{aligned} \tag{1}$$

The F-test is used to examine whether a long-term relationship exists among the variables by testing the significance of the lagged level variables. When the Brent price is expressed in euros, the null hypothesis of no cointegration is H_0: $\alpha_1 = \alpha_2 = \alpha_3 = 0$, whereas when the Brent price is expressed in dollars, the null hypothesis of no cointegration is H_0: $\alpha_1 = \alpha_2 = \alpha_3 = \alpha_4 = 0$. The computed F-statistics for cointegration are denoted as $F_{Pi}(P_i/B_j,T_i,R)$ and $F_{Pi}(P_i/B_j,T_i,R)$. Pesaran *et al.* (2001) tabulated two sets of critical values. The lower bound assumes that all of the regressors are I(0), and the upper bound assumes that all of the regressors are I(1). If the computed F-statistics lie above the upper bound, the null hypothesis is rejected, and cointegration is indicated. Conversely, if the computed F-statistics lie below the lower bound, the null hypothesis cannot be rejected; thus, the absence of cointegration is supported. If the statistics fall within the bounds, the inference is inconclusive.

After the cointegration is established, the ARDL long-term model can be estimated. The long-term model (in which u is the error term) is as follows:

$$\begin{aligned}P_i = {} & c_i + \sum_{k=1}^{p}\alpha_{1i}P_{i,t-k} + \sum_{k=0}^{q_1}\alpha_{2i}B_{j,t-k} + \sum_{k=0}^{q_2}\alpha_{3i}T_{i,t-k} \\ & + \sum_{k=0}^{q_3}\alpha_{4i}R_{t-k} + u_{i,t} \\ & \text{for } i = \{D, \mathrm{SP95}\} \text{ and } j = \{E, D\}\end{aligned} \tag{2}$$

The final step is to obtain the short-term dynamic parameters by estimating an error correction model associated with the long-term estimation. To incorporate an asymmetric response, we allow the short-term coefficient to vary depending on whether a change in a variable X is positive (ΔX^+) or negative (ΔX^-). Let $\Delta X_t^{\,+} = \mathrm{Max}\{0, \Delta X\}$ and $\Delta X_t^{\,-} = \mathrm{Min}\{0, \Delta X\}$. An asymmetric gasoline price response to a change in crude oil prices and to a change in oil taxes can be expressed as follows:

$$\begin{aligned}\Delta P_i = {} & \mu_i + \sum_{k=1}^{p}\left(\beta^+_{k,i}\Delta P^-_{i,t-k} + \beta^-_{k,i}\Delta P^-_{i,t-k}\right) \\ & + \sum_{k=0}^{l}\left(\lambda^+_{k,i}\Delta B^+_{j,t-k} + \lambda^-_{k,i}\Delta B^-_{j,t-k}\right) \\ & + \sum_{k=0}^{m}\left(\tau^+_{k,i}\Delta T^+_{i,t-k} + \tau^-_{k,i}\Delta T^-_{i,t-k}\right) \\ & + \sum_{k=0}^{n}\left(\rho^+_{k,i}\Delta R^+_{t-k} + \rho^-_{k,i}\Delta R^-_{t-k}\right) \\ & + \delta ECT_{t-1} + \nu_{i,t} \\ & \text{for } i = \{D, \mathrm{SP95}\} \text{ and } j = \{E, D\}\end{aligned} \tag{3}$$

where β^+, λ^+, τ^+ and ρ^+ are the short-term dynamic coefficients when changes in the variable are positive; β^-, λ^-, τ^- and ρ^- are the short-term dynamic coefficients when changes in the variable are negative; δ is the speed of adjustment; μ is a constant and ν is the error term.

III. Empirical Results

The ARDL bound test could lead to spurious results in the presence of I(2) variables because the critical values tabulated are simulated based on the assumption that the variables must be I(0) or I(1). We applied unit root tests and the results consistently show that the variables are stationary at the first difference; thus, the variables are I(1). We computed *F*-statistics using Akaike Information Criterion (AIC) to determine the lag structure. As reported in Table 1, the calculated *F*-statistics for diesel are higher than the upper bound critical value at the 1% level of significance, and the calculated *F*-statistics for SP95 are higher than the upper bound critical value at the 5% level of significance. These results suggest that a cointegration relationship exists between the gasoline prices and the other variables.

Table 1. Results of bounds *F*-test for cointegration

$F_{PD}(P_D/R_E,T_D)$	7.198***	
$F_{PSP95}(P_{SP95}/B_E,T_{SP95})$	4.428**	
	Lower bounds I(0)	Upper bounds I(1)
1% level	4.29	5.61
5% level	3.23	4.35
10% level	2.72	3.77
$F_{PD}(P_D/B_D,T_D, R)$	8.675***	
$F_{PSP95}(P_{SP95}/B_D,T_{SP95},R)$	4.573**	
	Lower bounds I(0)	Upper bounds I(1)
1% level	3.74	5.06
5% level	2.86	4.01
10% level	2.45	3.52

Notes: ***, ** and *Denote significance at the 1%, 5% and 10% levels, respectively.
The critical value bounds are obtained from Pesaran *et al.* (2001).

Hence, ARDL models could be established to determine the long- and short-term relationships. The orders of ARDL (1,0,0) and ARDL (1,0,0,0) were selected based on the AIC. The results obtained by normalizing the gasoline prices are presented in Table 2. As explained in Section II, taxes make up a large part of the gasoline prices. Crude oil prices explain only 20% or 25% of the prices. The exchange rate has a negative influence, which reflects the long-term appreciation of the euro, particularly since 2007.

The results of the short-term dynamic coefficients associated with the ARDL specifications are presented in Table 3. These short-term relationships are satisfactory because the Error Correction Terms (ECTs) related to the cointegration vector are significant and exhibit the expected negative sign. These coefficients correspond to the speed of the adjustment of gasoline prices to their equilibrium values determined by the long-term relationships. In this case, the coefficients are approximately 0.2, which reflects a relatively low adjustment to the long-term value. The estimates vary depending on whether the shock is positive or negative, showing an asymmetric response of gasoline prices. Changes in gasoline prices are sensitive to the most recent three changes in the case of a positive shock and only to the most recent change in the case of a negative shock. Moreover, the downward adjustment is lower than the upward adjustment. Thus, a 1% increase in the Brent price in euros implies an immediate increase of 0.12% in the diesel price, whereas a 1% decrease in the Brent price leads to an immediate decrease of only 0.07%. Changes in the exchange rate always have a negative effect on price developments. Being in the euro zone, where the crude oil price is traded in dollars, allows the exchange rate to have a damping effect on prices. We find that the estimates differ slightly depending on the chosen gasoline price. In particular, changes in the Brent price have more persistent effects on premium prices than on diesel prices.

Table 2. Estimated long-term coefficient-dependent variable (P_i)

	Diesel		SP95	
	$B_j = B_E$	$B_j = B_D$	$B_j = B_E$	$B_j = B_D$
c	0.822 (0.024)**	0.758 (0.023)**	1.049 (0.031)**	0.982 (0.031)**
B_j	0.269 (0.003)**	0.241 (0.003)**	0.200 (0.002)**	0.174 (0.002)**
T_i	0.685 (0.008)**	0.405 (0.007)**	0.682 (0.008)**	0.701 (0.008)**
R		–0.160 (0.008)**		–0.143 (0.007)**
Adjusted R^2	0.99	0.99	0.99	0.99

Note: ** and *Denote significance at the 5% and 10% levels, respectively.

Table 3. Estimated short-term coefficients

	Diesel		SP95	
	$B_j = B_E$	$B_j = B_D$	$B_j = B_E$	$B_j = B_D$
c	–0.000 (0.000)	–0.000 (0.000)	–0.000 (0.000)	–0.000 (0.000)
ECT(–1)	–0.175 (0.036)**	–0.173 (0.035)**	–0.199 (0.047)**	–0.197 (0.045)**
$\Delta P_i^+(-1)$	0.342 (0.048)**	0.248 (0.040)**	0.296 (0.055)**	0.300 (0.053)**
$\Delta P_i^+(-2)$	0.002 (0.025)	0.004 (0.024)	–0.011 (0.037)	–0.011 (0.034)
$\Delta P_i^+(-3)$	0.067 (0.024)**	0.075 (0.023)**	0.138 (0.034)**	0.139 (0.033)**
$\Delta P_i^-(-1)$	0.326 (0.044)**	0.326 (0.043)**	0.417 (0.058)**	0.402 (0.054)**
ΔB_j^+	0.116 (0.007)**	0.098 (0.007)**	0.075 (0.006)**	0.060 (0.007)**
$\Delta B_j^+(-1)$	0.024 (0.008)**	0.031 (0.008)**	0.025 (0.007)**	0.025 (0.007)**
$\Delta B_j^+(-2)$			0.021 (0.007)**	0.013 (0.006)**
ΔB_j^-	0.068 (0.007)**	0.053 (0.007)**	0.045 (0.006)**	0.032 (0.007)**
$\Delta B_j^-(-1)$	0.025 (0.008)**	0.019 (0.008)**	0.027 (0.007)**	0.021 (0.007)**
$\Delta B_j^-(-2)$			–0.014 (0.007)**	–0.011 (0.006)*
ΔT_i^+	0.828 (0.031)**	0.836 (0.031)**	1.000 (0.036)**	0.996 (0.035)**
$\Delta T_i^+(-1)$	–0.086 (0.043)**		–0.055 (0.051)	–0.073 (0.050)
$\Delta T_i^+(-2)$			–0.014 (0.050)	–0.005 (0.049)
$\Delta T_i^+(-3)$			–0.137 (0.049)**	–0.130 (0.049)**
ΔT_i^-	1.416 (0.078)**	1.361 (0.077)**	1.875 (0.091)**	1.840 (0.090)**
ΔE^+		–0.039 (0.023)*		–0.030 (0.020)
$\Delta E^+(-1)$		–0.076 (0.021)**		–0.043 (0.020)**
$\Delta E^+(-2)$		–0.046 (0.021)**		
ΔE^-		–0.012 (0.022)		–0.037 (0.019)*
$\Delta E^-(-1)$				–0.046 (0.019)**
Adjusted R^2	0.73	0.74	0.72	0.73
F-stat	248.7	208.2	231.7	174.8
DW	1.94	1.86	1.97	2.00
JB	399.18	413.49	150.85	147.26
BG LM	3.286	4.845	3.432	3.053
ARCH LM	7.929	9.981	5.178	5.151

Notes: BG LM, Breusch–Godfrey Lagrange Multiplier; DW, Durbin–Watson; JB, Jarque–Bera.
** and *Denote significance at the 5% and 10% levels, respectively.

We applied a number of diagnostic tests to the selected models. At the 5% significance level, none of the diagnostic tests exhibit any evidence of a violation of the classical linear regression model. Specifically, the model passes the Jarque–Bera test; thus, the errors are normally distributed. The Breusch–Godfrey Lagrange Multiplier (LM) and Autoregressive Conditional Heteroscedasticity (ARCH) LM tests show no evidence of serial correlation and heteroscedasticity.

IV. Conclusions

The empirical results of the study reveal that there is an asymmetric response of gasoline prices to shocks to the crude oil price. More specifically, we find that the adjustment of gasoline prices is lower when the crude oil price decreases than when it increases. Thus, our results indicate that there is asymmetry in the response of gasoline prices depending on the nature of the shock. Moreover, changes in the exchange rate always have a negative effect on price developments. Being in the euro zone, where the crude oil price is traded in dollars, allows the exchange rate to have a dampening effect on prices. Finally, we find that the estimates differ slightly depending on the chosen gasoline price. In particular, changes in the Brent price have more persistent effects on premium prices than on diesel prices.

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