How competitive is the Spanish Gasoline Market? Testing for Asymmetries in the Oil-Gasoline Price Relationship

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Abstract

This paper presents an extended error correction model (ECM) testing for short and long-run asymmetries in the oil-gasoline price relationship, which is estimated in two stages. We are using an updated dataset with weekly observations from 1997 to the first quarter of 2008 for the Spanish market including the introduction of the Euro and the beginning of the latest oil crisis. After estimating the long-run equilibrium by cointegration techniques, the asymmetric ECM is specified. This paper distinguishes itself from the existing literature through an in depth analysis of the individual time series for dynamic comovement, unit root and structural break. We confirm a structural beak for the exchange rate time series, coinciding with the physical introduction of the Euro and split the total sample in two subsamples. Asymmetric price transmission is found for the exchange rate movement for the first stage and short-run asymmetries in the second stage of the value chain. The empirical results also indicate that in the second subsample (2002-2008) the market has become more competitive, as no asymmetric pricing pattern can be confirmed.

CONTENTS

1. Introduction

The Spanish Commission for Competition (CNC) recently published a report dealing with the competitiveness of the Spanish gasoline market and concludes that missing competition, together with market entry barriers are the main reasons for above average gasoline prices in Spain compared with other member states of the European Union.[1]

This statement is very representative for the long ongoing discussion about market frictions in the sector. Big oil companies are suspected to take advantage of their market power not only by setting market entry barriers, but by asymmetrically adjusting prices. Especially in periods of volatile international markets, this is an issue to which both public opinion and policy makers are rather sensitive. As it has been mentioned by the existing literature (for example Bornstein et al. (1997), Meyer and von Cramon-Taubadel (2004)), an asymmetric pricing pattern is in line with the presence of market power.

[1] The report was published on 3 September 2009 and can be accessed directly on the webpage of the Commission: http://www.cncompetencia.es/

In this paper we are going to investigate the transmission of positive and negative changes in the price of crude oil to the price of gasoline with focus on the Spanish Market. We use an updated dataset with weekly observations, reaching from 1997 to the end of the first quarter of 2008, representing more than 580 individual observations. The Spanish market has been proven especially interesting, as it was one of the last European markets that liberalized the gasoline and oil sector and market concentration is still high[2].

In order to test for the hypothesis of existing asymmetries, we first measure the direct influence of crude oil prices on the retail (net-of-tax) gasoline price. In a second step we allow for asymmetries along the value-chain, which means we first test the crude oil – spot market price relationship and then the sport-retail price relationship, both individually. It has been proven very convenient to model the vertical market structure in two stages in order to get more detailed information on possible asymmetries along the Marketing chain.

An extended error-correction model (ECM), first introduced by Granger and Lee (1989), is used which we estimate in two steps. For the first step, the time series are tested for unit-root and the long-run equilibrium is estimated applying cointegration techniques. In the second step, the ECM is specified. The main advantage of this methodology is the possibility to distinguish between asymmetries related to the adjustment speed towards the long-run equilibrium and asymmetries related to transitory (short-run) price movements. Furthermore, we consider possible asymmetries on the effect of the exchange rate, as crude oil is paid in US dollars whereas gasoline is accounted in Euro (Spanish Peseta before 1999).

The empirical findings show asymmetries in the exchange rate movements for the first stage and also for the single stage approach. Special interest however is paid to existing short-run asymmetries in the second stage reflecting the public opinion of possible market power abuse. Furthermore we analyze possible economic reasons for the existence of asymmetries and potential for policy implementation.

The paper is structured as follows. First, we start with a brief overview of the existing literature on price transmission with special focus on asymmetric price transmission (APT) in

[2] In 2003, the two biggest suppliers, Repsol-YPF and Cepsa-Elf, control more than 60% of the Spanish gasoline retail outlets (Cotín-Pilart et. al (2009))

the oil-gasoline price relationship. Section 2.2 describes the theoretical model, before reviewing the econometric methodology in Section 3.1. It follows a description of the data in part 3.2, before presenting the regression results and formal tests for symmetry in Section 4. The paper finishes with concluding remarks and policy recommendation.

2. Literature and theoretical framework

2.1. Overview of existing literature

Price theory and market efficiency play an important role in neo-classical economics, as a flexible price system is needed for efficient resource allocation. In the applied literature a wide variety of often conflicting theories exist dealing with reasons for asymmetric price transmission. Different economic explications have been proposed so far, mainly market power in the context of non-competitive markets and adjustment costs. Further reasons discussed in literature are political intervention, asymmetric information, and inventory management (see also Meyer and Cramon-Taubadel (2004)). Following their paper, asymmetries that arise due to different responses to positive and negative shocks can be distinguished into three groups: magnitude of price transmission, time of adjustment, and a possible combination of both.

Within the applied literature, price asymmetries have been mainly tested for agricultural products and the oil-gasoline price relationship. However some studies can also be found relating APT to the banking sector (see for example Hannan and Berger (1991), Neumark and Sharpe (1992), and Arbatskaya and Baye (2004) and more recent approaches for the electricity spot market (Wölfing (2008)). Within the agricultural framework the following markets received special interest: beef and pork (for example Boyd and Brorsen (1988), Hahn (1990) and Goodwin and Holt, (1999)), fruits and vegetables (Ward, 1982 and Pick et at. 1991), the dairy sector (Frigon et al. (1999), Serra and Goodwin (2003), Capps and Sherwell (2005) and fish (Gonzales et al. (2003), Jaffry (2005)).

According to Peltzman (2000), who uses a large sample of 77 consumer and 165 producer goods, asymmetry is more the rule than the exception and can be found in two third of all cases. Due to his results, APT can be found equally likely in oligopolistic and competitive markets, so no clear statement about market power can be made.

In more recent literature there have been attempts to describe consumer search models in non-competitive markets (for example Tappata (2008), Lewis (2003), Johnson (2002), and Bornstein et al. (1997)). One main problem comparing the existing empirical literature are missing standards for econometric specification and modeling. Today, a wide range of econometric specification is available to the researcher and the findings may depend to some extend on the model chosen (Frey and Manera (2005)).

The oil-gasoline price relationship has received wide focus by empirical research. Usually, the studies differ in one or more of the following characteristics: country under scrutiny, time frequency and period of the data used, stage of transmission mechanism and the econometric model employed. The following outline provides an overview about the most relevant papers.

Wide popularity has been received by papers focusing on the second stage of the distribution chain and therefore relating wholesale gasoline prices directly to retail prices. The first studies have been developed by Bacon (1991) and Karrenbrock (1991) for the UK and US market respectively, both concluding the existence of asymmetry. Further examples for the second stage transmission are Kirchgässner and Kübel (1992), Duffy-Deno (1996), Asplund et al. (2000), Eckert (2002), Johnson (2002) and Bettendorf et al. (2003).

Other papers focus on the direct or single stage approach, relating crude oil prices directly to the retail gasoline price. Manning (1991) uses monthly data from 1973-1988 for the UK and employs an ECM allowing for asymmetry in the dynamic part of the equation. In a later work, Reilly and Witt (1998) go back to these findings and estimate a restricted ECM with monthly data allowing for short-run asymmetries.

Recent papers use the single stage approach for more sophisticated econometric models such as Godby et al. (2000) that use a threshold autoregressive (TAR) model within the ECM, Salas (2002) who employs ordered Probit, partial adjustment, and vector error correction models, and Radchenko (2005) who examines the impulse response functions of gasoline price asymmetry to a shock in oil price volatility.

Bornstein et al. (1997) uses weekly and biweekly data from 1986-1992 and confirms the public opinion that retail gasoline prices react more quickly to increases in crude oil prices than to decreases. Three possible interpretations of the presence of asymmetric gasoline price behavior are developed in his paper, on which we will focus in Section 4.2.

In another paper, Borenstein and Shepard (2002) focus on the first stage of the distribution chain, and they find that firms with market power adjust prices more slowly than competitive firms.

Bachmeier and Griffin (2003) consider daily data with an ECM for the period 1985-1998 and compare their findings with Borenstein et al. (1997). However, they do not confirm asymmetry for the US wholesale gasoline market mainly for two reasons: First, the estimation procedure used is a standard Engle–Granger two-step estimation, whereas Borenstein et al. (1997) use a non-standard estimation methodology and second, the choice of daily data rather than weekly data. Besides, the work of Borenstein et al. (1997) was extended by Balke et al. (1998) by using two different model specifications and weekly data. The authors do not obtain unambiguous evidence concerning asymmetry.

Galeotti et al. (2003) and Grasso, Manera (2007) use monthly data to compare different European markets, including the Spanish one, testing for single, first, and second stage asymmetries. Whilst Galeotti et al. employ a standard asymmetric ECM, Grasso and Manera distinguish three different econometric specifications: namely asymmetric ECM, threshold autoregressive ECM and ECM with threshold cointegration, and find that the type of market and number of countries which are characterized by asymmetric oil-gasoline price relations vary across model.

Finally, Contin-Pilart et al. (2009) analyze the competitive situation in the Spanish market, measuring price asymmetries before and after the abolishment of the state regulation. As pointed out before, Spain was one of the last countries where the free market was not introduced before 1998. In 2003, the Spanish based refiners control about 70% of the service stations (Repsol-YPF 43.8%, Cepsa-Elf 18,7%, and BP 6.9%). In their paper, Cotín-Pilart et al. find evidence for political intervention, which indicates that the Spanish government and the big oil companies worked closely together to keep inflation targets on track and to prepare Spain for the free market. However, no asymmetry is found.

As mentioned in the introduction to this section, the existence of APT in literature depends on different factors and is rather unstable. But as shown in the paper of Grasso and Manera (2007) asymmetries are found in around 65% of all cases, while more than 50% of the studies have been conducted with data for the US.

2.2 Theoretical Model

Figure 1 illustrates the market structure we assume for the empirical research. It shows the vertical Marketing chain and possible asymmetries that may arise along the price transmission. Crude oil (the upstream product) is the input factor for refined gasoline. As mentioned by Borenstein et al. (1997) motor gasoline is just one of many products that can be made from refining crude oil, such as diesel fuel, kerosene, heating oil, etc. However, the possibility for substitution in the production process is rather limited. The Spot market used as a proxy for the distribution stage, represents the intermediate for the refined gasoline on the way to the retail gasoline stations (downstream).

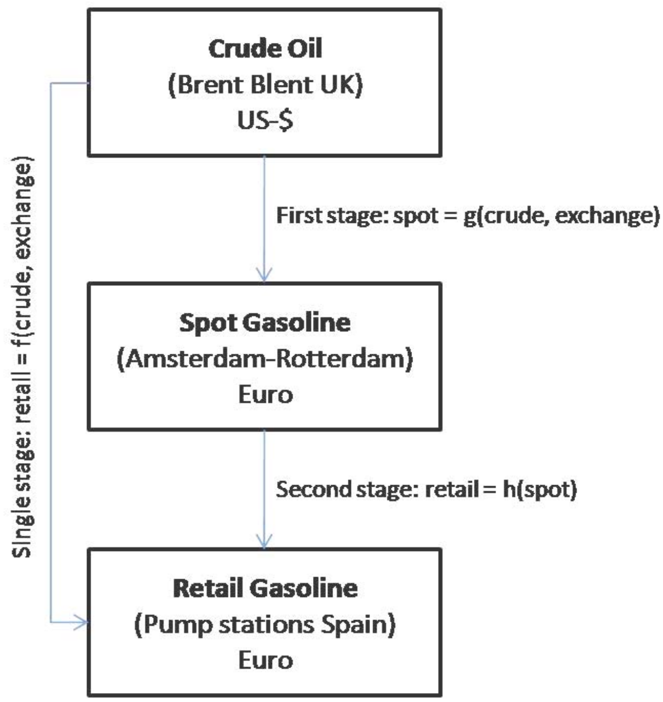

Figure 1, Market Structure

This setup allows us to test for three possible types of asymmetries. First, we test for single stage asymmetry, which shows the retail gasoline price as a function of crude oil and the exchange rate. Note that the exchange rate may be a source of asymmetry as crude oil is accounted in US-Dollars while the Spot Market in Amsterdam and the retail prices are measured in Euro. This setup will provide useful information on how the gasoline pump price behaves and how fast the adjustment takes place, when there are shocks to the input factors. Additionally, to get a deeper insight into the market movements, we want to check for asymmetry along the vertical structure. Therefore, we are going to measure possible asymmetries from the refineries (crude oil) to the distribution chain. In a second step, we test for possible asymmetries between the wholesalers and the final gasoline price at the service stations.

3. Empirical Model and Data Description

3.1 Econometric Approach

Most empirical studies mentioned in the literature review use a modern cointegration approach to investigate price asymmetries, namely an error-correction model (ECM). As first described in Hamilton (1994) and introduced by Engle and Granger (1987), a (nx1) vector time series $y_t = \begin{bmatrix} x_{1t} \\ x_{2t} \\ \vdots \\ x_{nt} \end{bmatrix}$ is said to be cointegrated if each of the series taken individually is I(1), that is, nonstationary with a unit root, while some linear combination of the variables in [a'y_t] is stationary, or I(0), for some nonzero (nx1) vector a. The long-run or equilibrium relationship between the series can be expressed in the general form

$$x_{1t} = a_1 + a_2 x_{2t} + ... + a_m x_{mt} + \varepsilon_t \, .$$

In our case, this equation relates output prices with input prices and the exchange rate. As mentioned in the theoretical model we want to test each stage of the production-distribution chain individually but also investigate the direct impact. In the single stage approach, retail prices can therefore be expressed in a reduced form as a function of crude oil prices and the exchange rate.

When we split the market according to the vertical structure, we examine in the first stage the effect of crude oil prices and the exchange rate on the spot price and in the second stage, the retail price as a function of the spot market. As both, the spot price and the retail gasoline price are measured in Euros, the exchange rate does not play any role in this second stage. According to this specification, we obtain three equilibrium relationships, whose residuals are tested stationary applying the Augmented-Dickey-Fuller (ADF) test. As Enders (2004) points out, the main feature of cointegrated variables is that their time paths are influenced by the extent of any deviation from long-run equilibrium. Thus an error correction term has to be introduced in the second estimation step to account for deviations from the (long-run) equilibrium. The general form of the error correction model (ECM) may therefore be written

$$\Delta x_{1t} = \alpha\,\hat{\varepsilon}_{t-1} + \sum_{i=1}^{p} \lambda_i \Delta x_{1t-i} + \sum_{i=0}^{p} \gamma_i \Delta x_{2t-i} + \ldots + \sum_{i=0}^{p} \delta_i \Delta x_{mt-i} + u_t$$

where Δ is the first-difference operator and *p* is the length of the lags. The error correction term $\hat{\varepsilon}_{t-1}$ is the lagged estimated residual from the equilibrium relationship obtained in the first step. The optimal lag length is determined by testing for significant lags in each equation. Additionally, we test each residual u_t for autocorrelation applying the Durbin-Watson d-statistic.

However, in the case of asymmetric adjustment, the idea is to allow different responses for positive and negative shocks. The above mentioned ECM was extended by Granger and Lee (1989) for the case of asymmetric adjustments, decomposing the cointegration residuals and the additional explanatory variables into positive and negative changes.

$$\Delta x_{1t} = \alpha^{+}\,\hat{\varepsilon}^{+}_{t-1} + \alpha^{-}\,\hat{\varepsilon}^{-}_{t-1} + \sum_{i=1}^{p} \lambda^{+}{}_{i} \Delta x^{+}{}_{1t-i} + \sum_{i=1}^{p} \lambda^{-}{}_{i} \Delta x^{-}{}_{1t-i} + \sum_{i=0}^{p} \gamma^{+}{}_{i} \Delta x^{+}{}_{2t-i} + \sum_{i=0}^{p} \gamma^{-}{}_{i} \Delta x^{-}{}_{2t-i}$$

$$+ \ldots + \sum_{i=0}^{p} \delta^{+}{}_{i} \Delta x^{+}{}_{mt-i} + \sum_{i=0}^{p} \delta^{-}{}_{i} \Delta x^{-}{}_{mt-i} + u_t$$

The asymmetry in the speed of adjustment can be introduced by defining

$\hat{\varepsilon}_t^{+} = \hat{\varepsilon}_t$ if $\hat{\varepsilon}_t > 0$, zero otherwise

$\hat{\varepsilon}_t^{-} = \hat{\varepsilon}_t$ if $\hat{\varepsilon}_t < 0$, zero otherwise.

Applying the same technique, short-run asymmetries can be captured by decomposing the first differences of the explanatory variables into positive and negative changes:

$\Delta x^{+}_{jt-i} = x_{jt-i} - x_{jt-i-1} > 0$, zero otherwise

$\Delta x^{-}_{jt-i} = x_{jt-i} - x_{jt-i-1} < 0$, zero otherwise,

where *j* represents the number of explanatory variables and *p* the length of the lags.

As assumed by Grasso et al. (2007), the inspection of the sign, magnitude and statistical significance of the estimated coefficients may offer a first insight about the existence of asymmetries. Nevertheless, the existence of asymmetries should be tested formally, subject to conventional F-test. Cook (1998) furthermore points out that in order to test for asymmetries in the error correction term, the coefficients ideally fulfill three criteria. First, they should possess plausible expected signs. Second, they must be statistically significant. Finally, they must be statistically different from each other.[3]

3.2 Data Description

Weekly observations for European crude oil prices (UK Brent Blend Spot Price, FOB) are obtained from the Energy Information Administration (EIA).[4] The gasoline spot prices are for the European market “Amsterdam-Rotterdam-Antwerp” and are obtained from the EIA as well. In order to employ the same type of gasoline for the spot market, we are using the 50ppm sulfur series, which was replaced by the ultra-low-sulfur fuel 10ppm recently, but is reported for the total sample period. Weekly retail gasoline prices for Spain for the most common type of gasoline (Euro95) are obtained from the Oil Bulletin published by the European Commission.[5] Contín-Pilart (2009) mentions that in Spain Euro95 accounts for about 80% of the gasoline sales. The Spanish data shows arithmetic average of the gasoline retail prices for more than 7500 petrol stations that are reported on a weekly basis to the national government.

[3] Cook et al. (1999) provide evidence from Monte Carlo experiments that standard tests of symmetry are affected by low power in reasonably large samples. Grasso et al. (2007) and Galeotti et al. (2003) overcome this problem by bootstrapping the F-statistics and obtaining the rejection frequencies by simulation, what we believe is not necessary in our case due the large sample size considered.

[4] Energy Information Administration: EIA, http://www.eia.doe.gov/

[5] European Commission: Energy, http://ec.europa.eu/energy/observatory/oil/bulletin_en.htm

Moreover, it should be pointed out that the crude oil (Brent) price is measured in US-Dollars per Barrel whereas retail gasoline in Spain is traded in Euro, Spanish Peseta before 2002. Thus, the exchange rate may play a relevant role for possible asymmetries. The series for the Dollar-Euro and Dollar-Peseta exchange rates are obtained from the Federal Reserve Bank of New York.[6] Weekly values are constructed by averaging the daily noon buying rates on each Friday of the week. All gasoline prices (spot and retail) are net-of-tax in order to provide easy comparability with other studies.

The following transformations have been applied to the data before conducting the empirical analysis. Retail gasoline prices, which are measured in Euro per 1000 liter (Spanish Pesetas for the period January 1997 - December 2001) are transformed to Euro per Barrel. Before January 1999, we use the exchange rate published by the European Oil Bulletin to transform the data into Euro.[7] After that period, the fixed exchange rate for the Spanish Peseta (1 Euro = 166.386 Pesetas) is taken into account.[8] The time series have been converted to barrels dividing by 1000 and multiplying by 158.987, the official barrel measure. Concerning the US-Dollar per Euro exchange rate before 1999, we use the Dollar-Peseta data and apply the above mentioned exchange rate from the European Oil Bulletin. The spot gasoline price, which is obtained from the US Government Agency EIA and therefore measured in Cents per Gallon, has been transformed first to Dollar per Barrel, multiplying by 42/100 (42 Gallons = 1 Barrel) and then transformed to Euro.

Finally, it should be mentioned that the time series for retail gasoline prices show some irregularities in the sampling period. The missing individual observations (not more than three values per year) have been recovered by constructing averages using the prior and the next week to the missing value. As Bornstein et al. (1997), we do not find our results affected by that irregularity. In the investigation, we therefore work with the full set of observations ranging from the first week of 1997 to week twelve 2008 (first quarter 2008) that counts up for 584 observations in total. Graph 1 shows the three time series in levels: gasoline and spot prices measured in Euro per barrel, Brent measured in Dollars per barrel.

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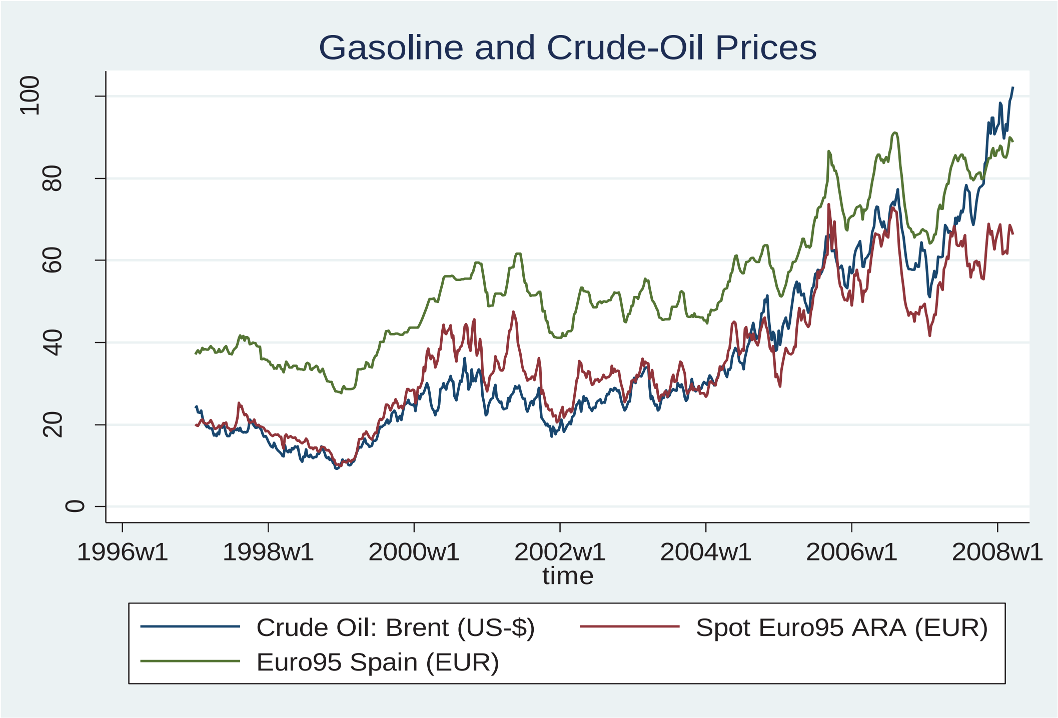
[6] Federal Reserve of New York: http://www.newyorkfed.org/

[7] For the period 1994-1998, the exchange rate from the national currencies to the European currency unit (ECU) is employed by the Oil bulletin. The author would like to thank Francis Soupart from the European Commission (CEC - DG TREN) for this useful information.

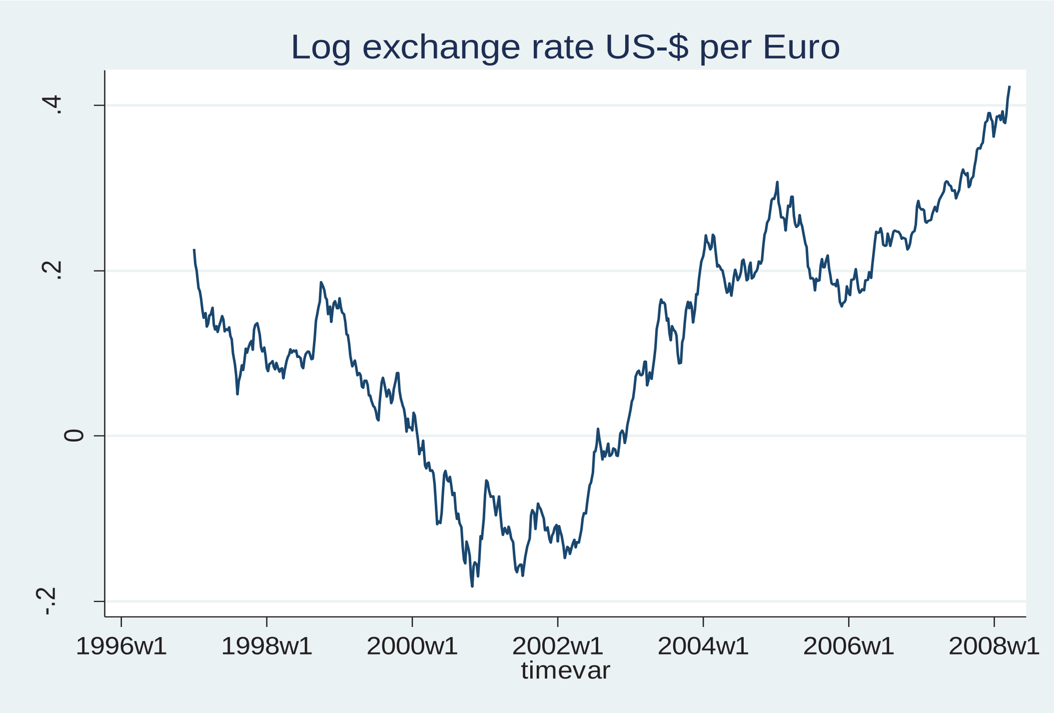
[8] European Central Bank: http://www.ecb.int/euro/intro/html/index.en.html



Graph 1, Gasoline and Crude Oil Prices in Levels

Graph 1 shows that the variables are not mean stationary, but rather have a strong comovement over time. This picture becomes even clearer when we observe the two gasoline time series for spot prices and retail prices, as they are both measured in Euro per Barrel.[9] For the period after 2007, crude oil prices are constantly higher than spot gasoline prices and at the end of the sample even above retail prices. This observation has to be related to the exchange rate movement. Graph 2 shows the exchange rate US-Dollar per Euro in logs for the total sample period. For the econometric analysis all variables have been log transformed.

[9] The data has been tested for comovement applying an dynamic approach developed by Den Hann (2000) who uses correlation of VAR forecast error at alternative forecast horizons. Appendix 1 shows the pairwise correlation of crude oil – spot, crude oil – retail and spot – retail for different time horizons.



Graph 2, Exchange rate US-Dollar per Euro in Logs

Graph 2 shows that there has been a significant devaluation of the Euro with respect to the Dollar until 2002. After that period however, the Euro recovered and reached an all-time high in 2008. The exchange rate evolution was responsible that Europe was not hit as hard as the United States by increasing crude oil prices.

On the other hand, the kink in Graph 2 indicates a structural change in 2002 (physical introduction of the Euro currency). We use the Chow-test to determine whether there has been a structural break for the single-stage but also for the two-stage approach and reject the null hypothesis of having equal parameters before and after the period 2001 week 52.[10] As a result, we estimate the asymmetric ECM not only for the sample period as a whole, but also for two subsamples: one reaching from 1997 week 1 to 2001 week 52 and the second from 2002 week 1 to 2008 week 12. This allows us to distinguish parameters for the three stages individually, but also for different sample periods.

[10] The corresponding test statistic can be found in Appendix 2.

All time series described here are found to be integrated of order one, I(1), applying augmented Dickey-Fuller (ADF) test, most of them with intercept and trend.[11] Table 1 describes the main summary statistics for the first log differences:

Variable specified in first log differences

	Obs	Mean	Std. Dev.	Min	Max
Crude (Brent)	583	0.002483	0.048483	-0.15576	0.198597
Spot (Euro95)	583	0.002057	0.048352	-0.20966	0.20947
Retail (Euro95)	583	0.0015	0.019966	-0.08371	0.089418
Exchange rate	583	0.000338	0.010609	-0.02991	0.028771

Table 1, Summary Statistics for crude prices, gasoline prices and exchange rate

Table 1 shows that all variables have a mean close to zero and it illustrates that Crude and Spot price show a considerable higher volatility in comparison with the retail price (standard deviation is more than double in both cases).

4. Empirical findings

4.1 Regression results

The ECM model described in Section 3.1 is estimated for the single stage and the two stage approach using data for Spain from the first week of 1997 to week 12 of 2008. As a first step the long-run equilibria have been estimated for all three stages individually, leading to three equilibrium relationships. Visual inspection of the residuals and the corresponding ADF statistics confirm that the residuals are stationary, which indicates that the series has been successfully cointegrated.[12] Table 2 shows the regression results. As pointed out in the data description part, in each stage, we distinguish three time periods: “Total Sample” refers to the whole sample from 1997 to 2008, “Subsample 1” refers to observations before the structural break (2001 week 52) and “Subsample 2” counts for observations from 2002 onwards.

[11] Appendix 3 includes the ADF statistics.
[12] Appendix 4 shows the first step (cointegration) regression results and the corresponding ADF statistics.

Block 1

1) Single stage: retail = f(brent, exchange)

	Total Sample: 1997 - 2008	**Subsample 1: 1997- 2001**	**Subsample 2: 2002 - 2008**
	579 observations R squared: 0.4078	257 observations R squared: 0.3868	324 observations R squared: 0.4857
Asymmetric adjustment speed			
α $^{(+)}$			
t-1	-0.0594 (.0221)	-0.0659 (.0375)	-0.0558 (.0259)
α $^{(-)}$			
t-1	-0.0768 (.0243)	-0.0629 (.0289)	-0.0848 (.0379)
Short-run asymmetries			
γ $^{(+)}$	0.1478 (.0312)	0.1187 (.0457)	0.2007 (.0442)
t-1	0.0633 (.0303)	0.0636 (.0319)	0.0425 (.0502)
t-2	0.0043 (.0274)	-0.0227 (.0357)	0.0486 (.0459)
t-3	-0.0082 (.0225)		
t-4	0.0477 (.0218)		
γ $^{(-)}$	0.1133 (.0272)	0.0498 (.0333)	0.2263 (.0505)
t-1	0.1047 (.0265)	0.0893 (.0369)	0.1034 (.0414)
t-2	0.0617 (.0283)	0.0897 (.0411)	0.0224 (.0513)
t-3	-0.0189 (.0255)		
t-4	0.0022 (.0244)		
Exchange rate asymmetries			
δ $^{(+)}$	-0.1307 (.1193)	0.1484 (.1809)	-0.5719 (.1474)
t-1	-0.3049 (.1269)	-0.4159 (.2126)	-0.2765 (.1628)
t-2	-0.2319 (.1129)		-0.2518 (.1392)
δ $^{(-)}$	-0.0628 (.1162)	-0.0504 (.158)	-0.0168 (.1661)
t-1	0.0293 (.1069)	-0.0890 (.1524)	0.1174 (.1553)
t-2	-0.1002 (.1179)		-0.0615 (.1673)
Autoregressive asymmetries			
λ $^{(+)}$			
t-1	0.2143 (.0675)	0.3364 (.1127)	0.1241 (.0776)
t-2	0.0500 (.0629)	-0.0625 (.0834)	0.0847 (.0751)
λ $^{(-)}$			
t-1	0.1055 (.0679)	0.0007 (.0823)	0.2011 (.1094)
t-2	0.2165 (.0711)	0.1904 (.0983)	0.2615 (.1121)
Constant term			
	0.0030 (.0026)	0.0021 (.0041)	0.0073 (.0025)

Block 2

2) First stage: spot = g(brent, exchange)

	Total Sample: 1997 - 2008	**Subsample 1: 1997- 2001**	**Subsample 2: 2002 - 2008**
	582 observations R squared: 0.4199	257 observations R squared: 0.5283	324 observations R squared: 0.4160
Asymmetric adjustment speed			
α $^{(+)}$			
t-1	-0.0795 (.0361)	-0.0402 (.047)	-0.0811 (.051)
α $^{(-)}$			
t-1	-0.1669 (.0395)	-0.2317 (.059)	-0.1283 (.0529)
Short-run asymmetries			
γ $^{(+)}$	0.4098 (.0894)	0.3165 (.1275)	0.5464 (.1272)
t-1	0.1357 (.0711)	0.1893 (.0853)	
γ $^{(-)}$	0.5970 (.0754)	0.4367 (.0998)	0.8625 (.1221)
t-1	0.1018 (.0923)	0.2211 (.1103)	
Exchange rate asymmetries			
δ $^{(+)}$	-0.5667 (.2865)	-1.2477 (.4419)	-0.6304 (.3383)
t-1	-0.5413 (.3106)	0.0158 (.479)	
t-2		0.8243 (.3525)	
δ $^{(-)}$	-1.6747 (.3262)	-2.0293 (.4476)	-1.2845 (.373)
t-1	0.5276 (.3339)	0.9028 (.4332)	
t-2		-0.4995 (.4355)	
Autoregressive asymmetries			
λ $^{(+)}$			
t-1	0.0989 (.0648)	0.1569 (.0903)	
λ $^{(-)}$			
t-1	0.0877 (.0811)	0.1338 (.1227)	
Constant term			
	0.0003 (.0046)	-0.0089 (.007)	0.0019 (.0046)

Block 3

3) Second stage: retail = h(spot)

	Total Sample: 1997 - 2008	**Subsample 1: 1997- 2001**	**Subsample 2: 2002 - 2008**
	581 observations R squared: 0.5319	258 observations R squared: 0.4473	324 observations R squared: 0.6554
Asymmetric adjustment speed			
α $^{(+)}$			
t-1	-0.0616 (.0242)	-0.0861 (.0418)	-0.0225 (.0252)
α $^{(-)}$			
t-1	-0.0545 (.021)	-0.0630 (.0236)	-0.1320 (.0486)
Short-run asymmetries			
γ $^{(+)}$	0.2360 (.0316)	0.2029 (.0372)	0.2558 (.0479)
t-1	0.0561 (.0272)	0.0825 (.0319)	0.0325 (.0332)
t-2	0.0353 (.0274)		0.0886 (.0295)
t-3			0.0296 (.0241)
γ $^{(-)}$	0.1710 (.0233)	0.0871 (.0303)	0.2702 (.0315)
t-1	0.1545 (.0259)	0.1509 (.0348)	0.1608 (.0333)
t-2	0.0390 (.0295)		0.0980 (.0355)
t-3			0.0556 (.0281)
Autoregressive asymmetries			
λ $^{(+)}$			
t-1	0.0675 (.0615)	0.1641 (.1104)	0.0000 (.071)
t-2	-0.0187 (.0556)		
λ $^{(-)}$			
t-1	0.0413 (.0662)	0.0047 (.0774)	0.0508 (.0957)
t-2	0.2079 (.0714)		
Constant term			
	0.0027 (.0015)	-0.0019 (.002)	0.0030 (.0019)

Robust standard errors (Huber-White Sandwich Estimator) reported in parenthesis ; α refer to the lagged asymmetric adjustment error obtained form the cointegration regression (Observations are split into positive and negative changes as stated in the econometric model; λ refer to autoregressive asymmetries while γ represent short run asymmetries and δ exchange rate asymmetries

Table 2, Regression results for single stage, first and second stage

The asymmetric adjustment speed to the long-run equilibrium is measured by α^+ and α^-, which represent the adjustment to the long run equilibrium from above and below, respectively. If prices deviate from the equilibrium values because of incomplete adjustment to past shocks or other market conditions, the current supply changes should move them back towards the equilibrium. Therefore, we expect the coefficients of the error correction term, α^+ and α^-, to be negative (see also Ratchenko (2005)). The short-run asymmetries can rather be seen transitory and refer to the corresponding upstream or input price. The positive coefficients γ^+ are associated with price increases and the negative ones (γ^-) with price reductions. In block 1) and 2) the short run asymmetries refer to changes in the Brent price and in block 3) they are related to the spot gasoline price. Another type of short-run asymmetry is measured by the exchange rate, which is only present in block 1) and 2). Autoregressive lagged asymmetries are captured by λ^+ and λ^- for positive and negative changes in the endogenous variable. The number of lags included in the estimation depends on two factors: first their individual significance and second the non-presence of autocorrelation in the residuals, which have been tested using the Durbin-Watson statistic. The d-statistic does not indicate significant autocorrelation for the residuals. [13]

The single stage estimates in block 1 do not show important asymmetries in the adjustment to the long-run equilibrium. All estimates possess the expected negative signs and are furthermore significant. Concerning the short-run asymmetries we find that for a given week the coefficients related to price increases are typically larger than the ones related to price decreases when considering the whole sample and the first subsample. However, as expected the lagged values at t-1 show the opposite pattern, which is a necessary condition to reach the long-run equilibrium. The coefficients associated with changes in the exchange rate are only significant for positive changes. The negative sign suggests that when the Euro appreciates with respect to the Dollar, and as a consequence the upstream product gets cheaper for the European market, changes are transmitted to the retail gasoline price. As a consequence, Spanish consumers take advantage of “favorable” changes in the exchange rate. For the full sample and subsample 1, the effect was biggest at time t-1, whereas for the last subsample the effect has been even stronger and presented itself the same week. As

[13] The corresponding d-statistic can be found in Appendix 5 and does not indicate significant autocorrelation; the only case where autocorrelation cannot be rejected at a 5% level of significance is the second subsample for the first stage.

autoregressive asymmetries are concerned, positive changes in lagged gasoline prices are bigger at a close time horizon, whereas in the negative changes the second lag exceeds the coefficients of the first. No clear statement about asymmetry can be drawn due to the lack of significance of the individual estimates.

Since we are interested in a deeper understanding of the gasoline-oil price relationship, we are going to test the two stages individually. The first stage represents the relationship from the refinery (crude oil) to the wholesale market. Once again, simple inspection of the estimates can give us first hints for the presence of asymmetries. The adjustment speed shows important differences in this setup, as demonstrated in the full sample period. Both estimates are significant and have a negative sign, although the speed of adjustment from below is at least twice larger than the adjustment speed from above. This means that when the current price is below the long-run equilibrium, the adjustment to the equilibrium is at least twice faster than when the price is above equilibrium. For the two subsamples, the asymmetry cannot be concluded as only the α^- are significant. However, it seems that in the first subsample, there were stronger asymmetries than in the second one.

Concerning the short-run asymmetries, the estimates related to negative price changes are typically larger than the ones related to positive changes, which means that the wholesale market is benefiting from short-run adjustments. The same is true for the exchange rate; however, in this case, a stronger asymmetry is present. For the whole sample, the coefficient related to negative changes in the exchange rate is three times larger than the one related to positive changes in the same week (both are significant). As the exchange rate is expressed as US-Dollar per Euro, the coefficient related to negative changes means that the US-Dollar appreciates with respect to the Euro. So, as both have a negative sign at time t, spot gasoline prices related to negative changes go down by a rate three times larger than for positive changes. On a first sight this seems rather odd, but as we can see at time t-1 for the total sample and for the first subsample, the lagged coefficients are changing signs, and an appreciation in the US-Dollar is translated into increasing spot gasoline prices. However, the magnitude is not as big as the one at time t. For the first stage, no autoregressive asymmetry is present.

Finally, the second stage shows the transmission from the spot market to the retail pump stations. We expect this stage to be less competitive than the first stage, as local market power may be an important feature. Once again, we first want to focus on the asymmetric adjustment speed to the long-run equilibrium. For the total sample and subsample one, only slight differences are present. However in subsample 2, where only the coefficient for adjustment from below is significant, we find important differences. Concerning the presence of short-run asymmetries, we find strong signals in favor of market power abuse. Positive coefficients typically exceed negative coefficients for the same week, which proves somehow the public opinion that the retail gasoline price raises faster, when the wholesale prices increases, as when the movement is downward. Although lagged effects seem to compensate for contemporaneous spot price changes. Some autoregressive asymmetry may be present in the second stage as well; however the coefficients are only significant in parts.

The overall results show that an analysis along the value chain may give a much richer insight into the behavior of the market agents for each stage. The higher coefficients in the first stage related to changes in crude oil show that prices in the input factor are almost instantaneously transmitted to the market prices, whereas in the second stage changes are distributed over more time periods. Focusing further on the exchange rate behavior, coefficients are way larger for the first stage as for single stage. Additionally, using two subsamples, we find that especially in short-run asymmetries coefficients in the second subsample exceed the estimates of the first, which induces that prices have been transmitted more directly to the buyers in recent years. One possible explanation of these changes might be that Spain was one of the last countries in the European Union where the state monopoly was abolished in October 1998 and the establishment of the free market took additional years to show effect.

4.2 Testing for Asymmetries

As mentioned in Section 3.1, to prove asymmetry, the estimates should be individually significant and significant different from each other. Table 3 shows the corresponding F-statistics, where each pair of coefficients are tested for equality using the Wald test. The null hypothesis H_0: $\alpha^+ = \alpha^-$, $\gamma^+ = \gamma^-$, $\gamma^+_{(t-1)} = \gamma^-_{(t-1)}$ and so on, is therefore tested pairs wise.

Block 1

F-statistic[1] for Single stage: retail=f(brent, exchange)

	Total Sample*: 1997 - 2008	**Subsample 1**: 1997- 2001**	**Subsample 2***: 2002 - 2008**
Asymmetric adjustment speed			
α (t-1)	0.20 (0.6540)	0.00 (0.9575)	0.30 (0.5850)
Short run asymmetries			
γ	0.48 (0.4873)	1.04 (0.3089)	0.10 (0.7518)
γ (t-1)	0.73 (0.3927)	0.19 (0.6622)	0.64 (0.4227)
γ (t-2)	1.54 (0.2158)	2.90 (0.0896)	0.11 (0.7355)
γ (t-3)	0.07 (0.7908)		
γ (t-4)	1.38 (0.2405)		
Exchange asymmetries			
δ	0.12 (0.7344)	0.47 (0.4952)	4.45 (0.0358)
δ (t-1)	2.70 (0.1007)	1.01 (0.3160)	2.16 (0.1431)
δ (t-2)	0.45 (0.5014)		0.57 (0.4495)
Autoregressive asymmetries			
λ (t-1)	1.09 (0.2973)	4.71 (0.0309)	0.29 (0.5921)
λ (t-2)	2.51 (0.1136)	3.02 (0.0833)	1.43 (0.2325)

*F (1,556), corresponding p-value in parentheses
**F (1, 240), corresponding p-value in parentheses
***F (1, 305), corresponding p-value in parentheses

[1] H0: Estimated parameters are equal (symmetry)

Block 2

F-statistic[1] for First stage: spot=g(brent, exchange)

	Total Sample': 1997 - 2008	**Subsample 1'': 1997- 2001**	**Subsample 2''': 2002 - 2008**
Asymmetric adjustment speed			
α (t-1)	1.85 (0.1744)	4.53 (0.0343)	0.29 (0.5924)
Short run asymmetries			
γ	1.76 (0.1856)	0.37 (0.5457)	2.16 (0.1430)
γ (t-1)	0.06 (0.8048)	0.04 (0.8469)	
Exchange rate asymmetries			
δ	4.53 (0.0338)	1.08 (0.2990)	1.16 (0.2826)
δ (t-1)	3.86 (0.0499)	1.36 (0.2450)	
δ (t-2)		3.81 (0.0522)	
Autoregressive asymmetries			
λ (t-1)	0.01 (0.9243)	0.02 (0.8927)	
λ (t-2)			

'F (1, 569), corresponding p-value in parentheses
''F (1, 242), corresponding p-value in parentheses
'''F (1.317), corresponding p-value in parentheses

Block 3

F-statistic[1] for Second stage: retail=h(spot)

	Total Sample+: 1997 - 2008	**Subsample 1++: 1997- 2001**	**Subsample 2+++: 2002 - 2008**
Asymmetric adjustment speed			
α (t-1)	0.04 (0.8494)	0.19 (0.6667)	2.97 (0.0856)
Short run asymmetries			
γ	1.93 (0.1652)	4.33 (0.0385)	0.04 (0.8399)
γ (t-1)	5.43 (0.0201)	1.69 (0.1943)	6.68 (0.0102)
γ (t-2)	0.01 (0.9371)		0.03 (0.8594)
γ (t-3)			0.36 (0.5465)
Autoregressive asymmetries			
λ (t-1)	0.08 (0.7826)	1.17 (0.2800)	0.19 (0.6599)
λ (t-2)	5.04 (0.0251)		

+F (1,568), corresponding p-value in parentheses
++F (1, 249), corresponding p-value in parentheses
+++F (1,311), corresponding p-value in parentheses

Table 3, Test for Asymmetries, F-statistics

The F-statistic for the single stage indicates slight asymmetries for the exchange rate in the total sample at t-1 and stronger asymmetries for the exchange rate in the second subsample at the same week. However, only those coefficients that are related to positive changes in the exchange rate are statistically significant. There is also evidence of asymmetries in the autoregressive system, especially in the first subsample, but once again there is a lack of pair wise significance.

By revising the first stage, we find asymmetries for the first subsample in the adjustment speed and no evidence of asymmetry in short-run movements. However important asymmetries show up in the exchange rate behavior, especially for the whole sample period. By testing for individual significance, the estimation results suggest the presence of asymmetries for the exchange rate at period t and t-1 (total sample). Nonetheless the adjustment process does not indicate presence of market power, as the wholesale benefits come from exchange rate movements (see Section 4.1).

For the second stage analysis, we find that there might be some asymmetry on the long-run equilibrium in the second subsample, but more important there are significant asymmetries on the short-run. These asymmetries are significant for the whole sample at one week lag, for subsample 1 within the same week and for subsample 2 at one week lag as well. In particular, the first two asymmetries mentioned are individually significant, show the expected signs and are statistically different from each other at the 5% level of confidence. Those asymmetries indicate that price increases are passed through on to the consumer faster than price decreases, what might be explained by less competitive markets in the second stage of the distribution chain. The asymmetries in the autoregressive system cannot be understood as significant, as the individual estimates are only significant in parts.

Bornstein et al. (1997) develop three important hypotheses to explain the asymmetric retail price adjustment. The first hypothesis states that prices are sticky downwards, as the old output price serves as a natural focal point for oligopolistic sellers. This statement is closely related to the trigger price model of oligopolistic coordination (see also Green and Porter (1984) and Tirole (1988)). Due to small retail margins, the gas stations have to pass through to the consumer a price increase in the wholesale market directly, whereas when the input price goes down, there is no need to react as fast. In the second hypothesis, they state that production lags and finite inventories of gasoline imply that negative shocks to the future optimal gasoline consumption path can be accommodated more quickly than positive shocks. However, this explanation especially proves asymmetries for the first stage, where crude oil is the main input product and the agents behave according to a consumption path.

Their last hypothesis, however, received more attention as it concerns consumer search related to the fact that volatility in crude oil prices, by creating a signal-extraction problem for consumers, lowers the search availability, finally leading to less competitive retail outlets. This effect is driven by the fact that the expected gain from search is smaller when crude oil prices are known to be volatile than when they are fairly stable. The gasoline retail chains might use this temporal decrease in the demand elasticity and take advantage of their market power to achieve temporally higher margins.

5. Conclusion

The applied literature on asymmetric price transmission in the oil-gasoline price relationship does not come to a clear-cut conclusion about the existence of asymmetries and their economic interpretation. The results mainly depend on the country under investigation, the frequency of the data and the econometric model used.

This paper analyzes the presence of asymmetric price transmission in the Spanish gasoline market and investigates the possibility of the presence of market power. Therefore, it provides a detailed analysis of the Spanish gasoline market by modeling the value chain in a two-stage approach.

The dynamic ECM applied in the econometric section allows us to distinguish between long-run and transitory asymmetries. We are using an updated dataset reaching form 1997 to the first quarter of 2008. This sample period includes the introduction to the free gasoline market in Spain (1998), the physical introduction of the Euro (2002) and the beginning of the 2008 oil crisis. Our empirical analysis shows a structural break coinciding with the Euro introduction in January 2002. As a consequence, the estimation is carried out for the whole sample period but also for two subsamples, which allows us to observe asymmetries along the value chain but also for different sample periods.

The empirical findings suggest the presence of asymmetries in the exchange rate adjustment for the first stage and short-run asymmetries in the second stage, which can be interpreted as presence of market power for the period 1997 to 2001 in the wholesale market. However, we do not find clear-cut evidence for the existence of asymmetries in subsample two, which might indicate that the market has become more competitive. The main question remains however, what can policy makers do against APT and a possible use of market power?

Romain et al. (2002) investigate in their paper the New York fluid milk markets and the impact of a price gouging law, setting a maximum markup for the distribution stage. They show that the introduction of the policy had a significant positive effect on the farm price transmission asymmetry. Furthermore, they point out that increased competition reduces the farm-retail price spread on the New York City and Upstate New York fluid milk market.

It should be noted that a maximum markup could only be implemented at a local or national level on the gasoline market, as it has been done in Spain during the transition period to the free market (1993 to 1998). In the case of crude oil and spot gasoline prices this would not be possible, as both are determined on the world market. However, as our empirical findings show, important short-run asymmetries arise typically in the second stage of transmission from wholesale to the pump station.

Moreover, the Spanish Ministry of Economics created a public website in the year 2000, where consumer prices for Spain are published on a monthly basis. This publication reports prices in the different states and autonomous regions, but also compares prices with other European countries and the evolution of crude oil prices and exchange rates. This attempt hopefully lowers significantly the search cost for the consumer and forces the oil companies to be more competitively. Furthermore, as mentioned by the before cited CNC report, additional market entries are necessary. An increasing number of competitors and a transparent information system would make an abuse of market power not impossible, but at least less likely in future periods.

A1. Graphs and Tables

A2. References

Arbatskaya, Maria & Baye, Michael R., 2004. Are prices 'sticky' online? Market structure effects and asymmetric responses to cost shocks in online mortgage markets, International Journal of Industrial Organization, Elsevier, vol. 22(10), pages 1443-1462,

Asplund, M., Eriksson, Friberg, R., 2000. Price adjustment by a gasoline retail chain. Scandinavian Journal of Economics 102, 101–121.

Bachmeier, L.J., Griffin, J.M., 2003. New evidence on asymmetric gasoline price responses. Review of Economics and Statistics 85, 772–776.

Bacon, R.W., 1991. Rockets and feathers: the asymmetric speed o adjustment of UK retail gasoline prices to cost changes. Energy Economics, 211–218.

Balke, N.S., Brown, S.P.A., Yücel, M.K., 1998. Crude oil and gasoline prices: an asymmetric relationship? Federal Reserve Bank of Dallas Economic Review, First Quarter, pp. 2–11.

Bettendorf, L., van der Geest, S.A., Varkevisser, M., 2003. Price asymmetries in the Dutch retail gasoline market. Energy Economics 25, 669–689.

Borenstein, S., Shepard, A., 2002. Sticky prices, inventories, and market power in wholesale gasoline markets. RAND Journal of Economics 33, 116–139.

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

Boyd, Milton S. and B. Wade Brorsen, 1988. Price Asymmetry in the U.S. Pork Marketing Channel. *North Central Journal of Agricultural Economics*, 10:103-109

Capps, Oral & Sherwell, Pablo, 2005. Spatial Asymmetry in Farm-Retail Price Transmission Associated with Fluid Milk Products, 2005 Annual meeting, July 24-27, Providence, RI 19316, American Agricultural Economics Association

Contin-Pilart I., Aad F. Correlje, M. Blanca Palacios, 2009. Competition, regulation, and pricing behaviour in the Spanish retail gasoline market, Energy Policy, Volume 37, Issue 1, , Pages 219-228

Cook, S., Holly, S., Turner, P., 1998. DHSY revisited: the role of asymmetries. Applied Economics 31, 775–778.

Cook, S., Holly, S., Turner, P., 1999. The power of tests for non-linearity: the case of Granger–Lee asymmetry. Economic Letters 62, 155–159.

Den Haan, Wouter J., 2000. The comovement between output and prices. Journal of Monetary Economics 46, 3-30.

Duffy-Deno, K.T., 1996. Retail price asymmetries in local gasoline markets. Energy Economics 18, 81–92.

Eckert, A., 2002. Retail price cycles and response asymmetry. Canadian Journal of Economics 35, 52–77.

Enders, Walter, 2004. Applied Econometric Time Series, John Wiley & Sons Inc, page, Second edition, pp. 335-337

Robert F. Engle and C. W. J. Granger, 1987. Co-Integration and Error Correction: Representation, Estimation, and Testing, *Econometrica* 55, 251-276.

Frigon, M., Doyon, M. and Romain, 1999. R., Asymmetry in Farm-Retail Price Transmission. In the Northeastern Fluid Milk Market, Food Marketing Policy Center, Research Report No. 45

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.M., Lintner, A., Stengos, T., Wandschneider, B., 2000. Testing for asymmetric pricing in the Canadian retail gasoline market. Energy Economics 22, 349–368.

Gonzales, F., Guillotreau, P., Le Grel L., Simioni, M., 2003. Asymmetry of price transmission within the French value chain of seafood products, Institut Nacional de la Recherche Agronomique,

Goodwin, B. y M. Holt, 1999. Price transmission and asymmetric adjustments in the U.S. beef sector. American Journal of Agricultural Economics 81, 630-37.

Granger, C.W.J., Lee, T.H., 1989. Investigation of production, sales and inventory relationships using multicointegration and non-symmetric error correction models. Journal of Applied Econometrics 4, S145–S159.

Grasso, M., Manera, M., 2007. Asymmetric error correction models for the oil-gasoline price relationship, Energy Policy 35, 156-177.

Green, E.J., Porter, R.H., 1984. Noncooperative collusion under imperfect price information. Econometrica 52, 87-100.

Hahn, W.F., 1990. Price transmission asymmetry in pork and beef markets. The Journal of Agricultural Economics Research 42, 21-30.

Hamilton, James D. 1994, Time Series Analysis, Princeton University Press, 1st edition, p. 571

Hannan, T.H., Berger, A.N., 1991. The rigidity of prices: Evidence from the banking industry. American Economic Review 81, 938-945.

Jaffry, S., 2002. Asymmetric price transmission: Case study of French hake value chain. IIFET Conference, Wellington, 19-22 August.

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

Karrenbrock, J.D., 1991. The behaviour of retail gasoline prices: symmetric or not? Federal Reserve Bank of St. Louis 73, 19–29.

Kirchgässner, G., Kübler, K., 1992. Symmetric or asymmetric price adjustments in the oil market: An empirical analysis of the relations between international and domestic prices in the Federal Republic of Germany, 1972–89. Energy Economics 14, 171–185.

Lewis, M., 2003, Asymmetric price adjustment and consumer search: An examination of the retail gasoline market, CSEM WP 120

Manera, M., Frey, G., 2005. Models of asymmetric price transmission. FEEM Working Paper No. 100.05

Manning, D.N., 1991. Petrol prices, oil price rises and oil price falls: some evidence for the UK since 1972. Applied Economics 23, 1535–1541.

Meyer, J., von Cramon-Taubadel S., 2004. Asymmetric price transmission: A survey. Journal of Agricultural Economics 55, 581-611.

Neumark, D., Sharpe S., 1992. Market structure and the nature of price rigidities: Evidence from the market for consumer deposits. Quarterly Journal of Economics, 657-680.

Peltzman, S., 2000. Prices rise faster than they fall. Journal of Political Economy 108, 466-502.

Pick D.H., Karrenbrock, J., Carman, H.F., 1990. Price asymmetry and marketing margin behavior: An example for California-Arizona citrus. Agribusiness 6, 75-84.

Radchenko, S., 2005/1. Oil price volatility and the asymmetric response of gasoline prices to oil price increases and decreases. Energy Economics 27, 708–730.

Reilly, B., Witt, R., 1998. Petrol price asymmetries revisited. Energy Economics 20, 297–308.

Romain, R., Doyon, M., Frigon, M., 2002. Effects of state regulations on marketing margins and price transmission asymmetry: Evidence from the New York City and Upstate New York fluid milk markets. Agribusiness 18, 301–315.

Salas, J.M.I.S., 2002. Asymmetric price adjustments and a deregulated gasoline market. Philippine Review of Economics 39, 38–71.

Serra, T, Goodwin, B., 2003, Price transmission and asymmetric adjustment in the Spanish Dairy Sector. Applied Economics 35, 1889 – 1899.

Tappata, Mariano E., 2008. Rockets and Feathers: Understanding Asymmetric Pricing.

Tirole, Jean. 1988. The Theory of Industrial Organization. The MIT Press, Cambridge, MA.

Ward, R.W., 1982. Asymmetry in retail, wholesale and shipping point pricing for fresh vegetables. American Journal of Agricultural Economics 64, 205-212.

Wölfing, N.M., 2008. Asymmetric price transmission in supply function equilibrium, carbon prices and the German electricity spot market. ZEW - Centre for European Economic Research Discussion Paper No. 08-040.

A3. Appendix

Appendix 1

Dynamic Comovement Analysis: Applying methodology developed by Den Haan (2000), using VAR forcast errors at different time horizons. The general VAR can be specified:

$$X_t = \mu + Bt + Ct^2 + \sum_{t=1}^{L} A_t X_{t-1} + \varepsilon_t$$

Where A is an N x N matrix of regression coefficients, μ, B and C are N-vectors of constants, ε is an N-vector of innovations and total number of lags included is equal to L. Note that the elements of ε are assumed to be serially uncorrelated but they can be correlated with each other. The differences between the forecasts and the realizations are used to construct the time series of the FC errors and to calculate the covariance.

1) Single Stage: brent – retail[1]

Dynamic correlation of brent ($/Barrel) and retail (EUR/Barrel)
Lag selection: AIC (-14.2768)
4 lags, constant included in the estimation of the VAR

Forcast horizon	**Correlation**	**Mean across samples**	**90% Confidence Interval**	
			Lower band	**Upper band**
1 week	0.3364	0.335	0.2659	0.4028
4 weeks	0.6512	0.6483	0.5739	0.7128
3 month	0.7495	0.7437	0.657	0.8147
6 month	0.7759	0.7697	0.6787	0.8439
1 year	0.7869	0.7805	0.6872	0.8552
2 years	0.7923	0.7858	0.6919	0.8606
4 years	0.795	0.7884	0.6942	0.8635

2) First Stage: brent – spot[1]

Dynamic correlation of brent ($/Barrel) and spot (EUR/Barrel)
Lag selection: AIC (-12.4792)
7 lags, constant included in the estimation of the VAR

Forcast horizon	**Correlation**	**Mean across samples**	**90% Confidence Interval**	
			Lower band	**Upper band**
1 week	0.5059	0.5053	0.4411	0.5676
4 weeks	0.7154	0.7151	0.6521	0.7692
3 month	0.8104	0.8073	0.7433	0.8615
6 month	0.8306	0.8274	0.7593	0.8838
1 year	0.8409	0.8375	0.7668	0.8939
2 years	0.8462	0.8427	0.7712	0.8991
4 years	0.8489	0.8453	0.7738	0.9021

3) Second Stage: spot – retail[1]

Dynamic correlation of spot (EUR/Barrel) and gas (EUR/Barrel)
Lag selection: AIC (-14.5993)
10 lags, constant included in the estimation of the VAR

Forcast horizon	**Correlation**	**Mean across samples**	**90% Confidence Interval**	
			Lower band	**Upper band**
1 week	0.5618	0.56	0.4992	0.6203
4 weeks	0.8398	0.838	0.7997	0.8736
3 month	0.926	0.9226	0.8911	0.9462
6 month	0.9423	0.9383	0.9046	0.9623
1 year	0.9502	0.9463	0.913	0.9695
2 years	0.9543	0.9504	0.9177	0.9729
4 years	0.9563	0.9525	0.9202	0.9747

[1] variables specified in logs, unit-root imposed, bootstrapped standard errors (2000 replications), Lag length defined by Aikaike-Information-Criterium (AIC)

Appendix 2

Chow-test for the existence of a structural break in period 2001, week52 (observation 260)

The general annotation for the Chow-test is[14]

$$\frac{\dfrac{ess\_c-(ess\_1+ess\_2)}{k}}{\dfrac{ess\_1+ess\_2}{N\_1+N\_2-2k}} \sim F(k, N\_1+N\_2-2k)$$

Where:
- ess_c:_ error sum of squares for complete sample
- ess_1: error sum of squares for first subsample
- ess_2: error sum of squares for second subsample
- k: number of estimated parameters
- N_1 and N_2: number of observations for the two groups

1) Single stage[1]		2) First stage[2]		3) second stage[3]	
ess_c:	0.180242	ess_c:	0.841482	ess_c:	0.135137
ess_1:	0.080176	ess_1:	0.400296	ess_1:	0.066954
ess_2:	0.088543	ess_2:	0.393704	ess_2:	0.059457
N_1:	259	N_1:	259	N_1:	259
N_2:	324	N_2:	324	N_2:	324
k:	7	k:	7	k:	5
F(7,569):	**5.551502**	**F(7,569):**	**4.86104**	**F(5,569):**	**7.910015**
p-value:	**0.0000**	**p-value:**	**0.0000**	**p-value:**	**0.0000**
[1]retail=f(brent, exchange)		[2]spot=g(brent, exchange)		[3]retail=h(spot)	

[14] Chow (1960), see also: http://www.stata.com/support/faqs/stat/chow.html

Appendix 3

Augmented Dickey-Fuller (ADF) test statistic

1) Variables specified in logs:

	Test statistic	1% critical value	5% critical value	10% critical value
1) Natural logarithm of retail gasoline price (Euro95), 583 observations[1]				
Z(t)	**-1.791**	-3.96	-3.41	-3.12
corresponding p-value	**0.709**			
2) Natural logarithm of gasoline spot price (ARA-Euro95), 583 observations[1]				
Z(t)	**-2.455**	-3.96	-3.41	-3.12
corresponding p-value	**0.351**			
3) Natural logarithm of Brent Blent (FOB), 583 observations[2]				
Z(t)	**0.01**	-3.43	-2.86	-2.57
corresponding p-value	**0.9593**			
4) Natural logarithm of exchange rate: US-$ per Euro, 583 observations[2]				
Z(t)	**0.267**	-3.43	-2.86	-2.57
corresponding p-value	**0.9758**			

[1] *including intercept and trend*

[2] *including intercept*

2) Variables specified in first log-differences:

	Test statistic	1% critical value	5% critical value	10% critical value
1) First log differences: retail price for gasoline (Euro95), 582 observations[1]				
Z(t)	**-15.723**	-2.333	-1.647	-1.283
corresponding p-value	**0.0000**			
2) First log differences: spot price for gasoline (ARA-Euro95), 582 observations[1]				
Z(t)	**-19.823**	-2.333	-1.647	-1.283
corresponding p-value	**0.0000**			
3) First log differences of Brent Blent (FOB), 582 observations[2]				
Z(t)	**-21.684**	-3.43	-2.86	-2.57
corresponding p-value	**0.0000**			

4) First log-differences of exchange rate: US-$ per Euro, 582 observations[2]				
Z(t)	**-18.79**	-3.43	-2.86	-2.57
corresponding p-value	**0.0000**			

[1] *including intercept and drift*

[2] *including intercept*

Appendix 4

Estimation of long-run (equilibrium) relationships

1) Long run equilibrium Single Stage:

gas = $\beta_0 + \beta_1$ *brent + β_2 *er

Number of observations: R-squared:	584 0.9648
β_0 2.0433 (0.01617)	
β_1 0.5695 (0.00501)	
β_2 -0.4142 (0.01983)	

all variables in logs: gas (retail gasoline price), brent (UK Brent blent FOB), er (exchange rate: US-$ per EUR), standard errors in parenthesis

2) Long run equilibrium First Stage:

spot = $\beta_0 + \beta_1$ *brent + β_2 *er

Number of observations: R-squared:	584 0.9662
β_0 0.4149 (0.02487)	
β_1 0.9236 (0.0077)	
β_2 -0.9282 (0.30498)	

all variables in logs: spot (spot gasoline price), brent (UK Brent blent FOB), er (exchange rate: US-$ per EUR), standard errors in parenthesis

3) Long run equilibrium Second Stage:

gas = $\beta_0 + \beta_1$ *spot

Number of observations: R-squared:	584 0.9712
β_0 1.7669 (0.15711)	
β_1 0.6275 (0.00448)	

all variables in logs: gas (retail gasoline price), spot (spot gasoline price) standard errors in parenthesis

Corresponding ADF-statistic for the residuals:

	Test statistic	1% critical value	5% critical value	10% critical value
1) ADF test statistic for residuals (Single Stage)[1]				
Z(t)	**-6.518**	-3.43	-2.86	-2.57
p-value	**0.0000**			
2) ADF test statistic for residuals (First Stage)[1]				
Z(t)	**-6.567**	-3.43	-2.86	-2.57
p-value	**0.0000**			
3) ADF test statistic for residuals (Second Stage)[1]				
Z(t)	**-6.113**	-3.43	-2.86	-2.57
p-value	**0.0000**			

[1] include intercept

Appendix 5

Durbin-Watson test statistic for autocorrelation of the residuals

1) Single stage: retail =f(brent, exchange)

1) Single stage: retail = f(brent, exchange)			
	Total Sample: 1997 - 2008	Subsample 1: 1997- 2001	Subsample 2: 2002 - 2008
	d-statistic(23,579)	d-statistic(17,257)	d-statistic(19,324)
DW*	2.016349	2.00462	2.016585
lags	1	1	1
Chi^2	0.896	0.152	0.715
df**	1	1	1
$p > Chi^2$	0.3438	0.6971	0.3979

* DW: Durbin Watson d-statistic

**degrees of freedom

2) First stage: spot =g(brent, exchange)

2) First stage: spot = g(brent, exchange)			
	Total Sample: 1997 - 2008	Subsample 1: 1997- 2001	Subsample 2: 2002 - 2008
	d-statistic(13,582)	d-statistic(15,257)	d-statistic(7,324)
DW*	1.991219	1.984429	1.792355
lags	1	1	1
Chi^2	0.129	0.065	4.11
df**	1	1	1
$p > Chi^2$	0.72	0.7986	0.0426

* DW: Durbin Watson d-statistic

**degrees of freedom

3) Second stage: retail =h(spot)

3) Second stage: retail = h(spot)			
	Total Sample: 1997 - 2008	Subsample 1: 1997- 2001	Subsample 2: 2002 - 2008
	d-statistic(13,581)	d-statistic(9,258)	d-statistic(13,324)
DW*	2.011587	1.98296	2.026055
lags	1	1	1
Chi^2	0.989	0.057	2.391
df**	1	1	1
$p > Chi^2$	0.3199	0.8106	0.122

* DW: Durbin Watson d-statistic

**degrees of freedom