Asymmetric Error Correction Models for the Oil-Gasoline Price Relationship

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Summary

The existing literature on price asymmetries does not systematically investigate the sensitivity of the empirical results to the choice of a particular econometric specification. This paper fills this gap by providing a detailed comparison of the three most popular models designed to describe asymmetric price behaviour, namely asymmetric ECM, autoregressive threshold ECM and ECM with threshold cointegration. Each model is estimated on a common monthly dataset for the gasoline markets of France, Germany, Italy, Spain and UK over the period 1985-2003. All models are able to capture the temporal delay in the reaction of retail prices to changes in spot gasoline and crude oil prices, as well as some evidence of asymmetric behaviour. However, the type of market and the number of countries which are characterized by asymmetric oil-gasoline price relations vary across models. The asymmetric ECM yields some evidence of asymmetry for all countries, mainly at the distribution stage. The threshold ECM strongly rejects the null hypothesis of symmetric price behaviour, particularly in the case of France and Germany. Finally, the ECM with threshold cointegration finds long-run asymmetry for each country in the reaction of retail prices to oil price changes.

Keywords: Oil prices, Gasoline prices, Asymmetries, Error correction models

JEL Classification: C22, D40, Q40

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1. Introduction

The transmission of positive and negative changes in the price of oil to the price of gasoline is very relevant for both consumers, who tend to be very sensitive to the money they pay for the fuel consumed by their cars, and researchers, who are often requested to provide plausible explanations of the observed temporal behaviour of the oil-gasoline price relationship.

The notion that gasoline prices react quickly to oil price increases and slowly to oil price reductions is largely accepted among consumers. The levels recently hit by oil and gasoline prices and the present uncertainty in supply and reserve availability have contributed to reinvigorate the interest in the asymmetric transmission of changes in the price of oil to the price of gasoline. According to the latest Oil Market Report issued by the International Energy Agency, oil prices strengthened for most of January 2005 and then slightly declined in early February 2005. During the same period, gasoline prices recorded a rally. On Friday, 4th March 2005 Brent has been quoted 51.73 U.S. dollars per barrel in London, whereas in New York the price of WTI has reached 54 U.S. dollars. Moreover, the average price of the OPEC oil (which is based on seven different oil qualities) has hit the level of 48.36 U.S. dollars, while only on Wednesday, 2nd March 2005 it was quoted 47.01 U.S. dollars. On the product side, the Italian gasoline price at the pump is close to 1.20 Euros per litre, while gasoil has been quoted Euros 1.09: both are the maximum levels recorded over the last three months.

The literature looking for empirical evidence in support of asymmetries in the transmission mechanism is wide. This literature employs a variety of reduced-form dynamic regression models relating the price of gasoline to the price of oil. Findings vary across countries, time periods, frequency of the data, markets and models, but in general they fail to provide strong evidence that prices rise faster than they fall.

The aim of this paper is to address the following question: to what extent does the empirical evidence on price asymmetries depend on the specific model used to analyze the relationship between gasoline and oil prices? This question is particularly relevant, since the existing literature does not systematically investigate the sensitivity of the empirical results to the choice of a particular econometric specification. Actually, one of the few attempts to explain the variability of the empirical findings on price asymmetries goes back to Shin (1994), who nevertheless argues that the contradictory results are mainly due to the lack of homogeneity in the data, rather than to different models.

The present paper fills this gap by providing a detailed comparison of the three most popular models designed to describe asymmetric price behaviour, namely asymmetric error correction model (henceforth asymmetric ECM), autoregressive threshold ECM and ECM with threshold cointegration. In order to reduce the proportion of variability in the results due to different countries, periods of time, data frequencies and markets, each model is estimated on a common monthly dataset which describes the retail and wholesale gasoline markets of France, Germany, Italy, Spain and UK over the period 1985-2003.

The plan of the paper is as follows. An exhaustive review of the econometric literature on price asymmetries in the gasoline market is offered in Section 2. Section 3 describes the data and the econometric models used in the empirical analysis. The results are presented and discussed in Section 4. Section 5 provides some concluding remarks.

2. Overview of the literature

Numerous attempts have been made to analyze the relationship between the price of crude oil and the price of gasoline (or other petroleum products). Studies typically differ in one or more of the following aspects: the country under scrutiny; the time frequency and period of the data used; the stage of the transmission mechanism, i.e. either retail or wholesale, or both; the dynamic model employed in the empirical investigation.

The problem of a different response to price increases and decreases is first considered in Bacon (1991), where attention is paid to the U.K. gasoline market but limited to the second stage of the transmission chain (the ex-Rotterdam spot price is used as a proxy of the product price). Biweekly data are used for the period 1982-1989. The author finds that increases in the product price are full transmitted within two months, in the case of price reductions an extra week is necessary; changes in the exchange rate necessitate two extra weeks relative to product prices before being incorporated in retail gas prices.

Again the U.K. is the country studied by Manning (1991), who instead looks directly at the impact of changes in oil prices on retail prices. The data are monthly for 1973-1988 and an ECM specification allowing for asymmetry only in the dynamic part of the equation. It is found weak and non-persistent asymmetry in price changes, which is absorbed within four months. No formal tests of asymmetric price effects are however performed.

Karrenbrock (1991) employs 1983-1990 monthly data to study the empirical relationship between U.S. wholesale and (after tax) retail gasoline prices. Operationally, the author uses a distributed lags model to find that the length of time in which a wholesale price increase is fully reflected in the retail gasoline price is the same as that of a wholesale price decrease for premium and unleaded regular gasoline. Instead, wholesale price increase for leaded regular gasoline are passed along to consumer more quickly than price increases. Nevertheless, the author concludes, contrary to the popular belief that consumers do not benefit from wholesale gasoline price decreases, these are eventually passed along to consumers as fully as are wholesale gasoline price increases.

Kirchgässner and Kübler (1992) also look at Western Germany for the period 1972-1989 using monthly data. The authors consider the response of both consumer and producer leaded gasoline prices to the spot price of the Rotterdam market; they do so for two sub-periods, before and after January 1980. The methodology adopted is very rigorous, as the variables are tested for, respectively, unit roots, Granger causality, cointegration, and structural breaks. When cointegration cannot be rejected, both symmetric and asymmetric ECMs are fitted. Unfortunately, the asymmetry is permitted only for price changes, thus allowing only for a different response in the short-run but not in the long-run. Briefly stated, the results show that, while long-run reactions are not significantly different for the 1970s and the 1980s, there is considerable asymmetry in the former period but not in the latter in the short-run adjustment processes. In particular, reductions in the Rotterdam prices are transferred faster to German markets than increases.

Shin (1994) relates the average wholesale price of oil products to the price of oil in his investigation of the U.S. market using monthly data for the period 1982-1990. His dynamic model shows no evidence of asymmetric effect.

Again the U.S. attracts the interest of Duffy-Deno (1996), and in particular the downstream relationship between wholesale and net-of-tax-retail gasoline prices The data this time are weekly for 1989-1993 and the econometric model shows strong persistent asymmetries, with a complete adjustment in the case of price rises and incomplete for price falls.

Borenstein et al. (1997) study the U.S. gasoline market using weekly data for 1986-1992. The empirical investigation confirms the common belief that retail gasoline prices react more quickly to increases in crude oil prices than do decreases (4 weeks versus 8 weeks). An ECM is estimated but, like the previous paper, only asymmetry for price changes is permitted. The authors offer three possible interpretations of the presence of asymmetric gasoline price behaviour. The first justifies downward gasoline price stickiness in terms of the existence of a natural focal point for oligopolistic sellers when oil prices are falling. According to the second, production lags and inventories allow to a quicker accommodation of negative shocks to optimal future consumption than positive shocks. The third interpretation relates oil price volatility to the degree of competition in the retail market.

Balke et al. (1998) extend the work of Borenstein at al. (1997) by using two different model specifications with weekly data from 1987 through 1997. In particular the authors use a distributed lag model in the levels of prices with asymmetric effects and an ECM representation which allows for both long-run and short-run asymmetry. On the basis of an encompassing test this last specification is preferred. Both models involve three prices, with the wholesale price depending upon oil and spot prices and the retail price upon wholesale and spot prices. The author do not obtain unambiguous evidence concerning asymmetry, been weak in the specification in levels and moderate and persistent in the ECM.

Reilly and Witt (1998) come back to the U.K. market to revisit the evidence of Bacon (1991) and Manning (1991) with monthly data for 1982-1995 and emphasizing the role of the dollar-pound exchange rate and the potential asymmetries associated with it, in addition to those of crude oil prices. A restricted ECM is estimated which allows only for short-run asymmetry. The hypothesis of a symmetric response by petrol retailers to crude price rises and falls is rejected by the data, and so is for changes in the exchange rate.

Akarca and Andrianacos (1998) investigate the dynamic relationship between crude oil and retail gasoline prices during the last 21 years and show that, in February 1986, this relationship had drastically changed. Since then, the results suggest that gasoline prices include higher profit margins, they are substantially less sensitive to changes in crude oil prices, and are more volatile.

Brown and Yucel (2000) examine the market conditions underlying the asymmetric relationship between gasoline and crude oil prices. They find the observed asymmetry is unlikely to be the result of monopoly power. The remaining explanations for the asymmetry suggest that policies to prevent an asymmetric relationship between gasoline and crude oil prices are likely to reduce economic efficiency.

Other papers look at the experience of other countries. For example, Godby et al. (2000) study the Canadian market for both premium and regular gasoline. The analysis is based on weekly data for thirteen cities between 1990 and 1996. By noting that the asymmetric ECM specifications used in previous studies are misspecified if price asymmetries are triggered by a minimum absolute increase in crude cost, a Treshold AutoRegressive model within an ECM is implemented in the paper. On this basis the authors fail to find evidence of asymmetric pricing behavior.

Asplund et al. (2000) investigate the Swedish retail market by fitting a restricted ECM with asymmetries only on the short-run dynamic components. The data are monthly and cover the period 1980 through 1996. There is some evidence that in the short-run prices are stickier downwards than upwards. Also, prices respond more rapidly to exchange rate movements than to the spot market prices.

Borenstein and Shepard (2002) propose a model with costly adjustment of production and costly inventories, which implies that wholesale gasoline prices will respond with a lag to crude oil cost shocks. Unlike explanations that rely upon menu costs, imperfect information, or long-term buyer/seller relationships, this model predicts that futures prices for gasoline will adjust incompletely to crude oil price shocks that occur close to the expiration date of the futures contract. Examining wholesale price responses in 188 gasoline markets, they also find that firms with market power adjust prices more slowly than do competitive firms, which is consistent with the model.

Weekly retail gasoline prices in Windsor, Ontario, from 1989 to 1994 are analyzed by Eckert (2002). Retail prices appear to respond faster to wholesale price increases than to decreases, but exhibit a cyclic pattern inconsistent with a common explanation of response asymmetry. The author reconciles these observations through a model of price cycles. Prices on the downward portion of the cycle appear insensitive to costs, compared with price increases, supporting the theory that price decreases result from battles over market share. This pattern resembles a faster response to cost increases than to decreases, and the conclusion that asymmetry indicates a role for competition policy may be inappropriate.

Salas (2002) uses an ordered probit, a partial adjustment, and a vector ECM to characterize price adjustments in the Philippine retail gasoline market since its deregulation. He finds that pricing decisions of oil firms depend significantly on eight weeks of previous changes in crude cost. Moreover, the speed of adjustment of retail prices to their long-run equilibrium relation with crude cost has been following an accelerating trend but is vulnerable to intervening factors. Lastly, the empirical evidence suggests that pump prices respond more quickly and fully to increases in crude cost rather than to decreases.

Bachmeier and Griffin (2003) consider daily data and adopt an Engle-Granger two step approach. No evidence of asymmetry is found for the American wholesale gasoline market over the period 1985-1998. In contrast with Borenstein et al. (1997), who claim that gasoline prices rise quickly following an increase in the price of crude oil but fall slowly following a decrease, they estimate an ECM with daily spot gasoline and crude-oil price data over the period 1985-1998 and find no evidence of asymmetry in wholesale gasoline prices. The sources of the difference in results are twofold. First, a standard Engle-Granger two-step estimation procedure is used, whereas Borenstein et al. (1997) use a non-standard estimation methodology. Second, even with the same non-standard specification, the use of daily rather than weekly data yields little evidence of price asymmetry.

Bettendorf et al. (2003) analyse the retail price adjustments in the Dutch gasoline market. They estimate an asymmetric ECM on weekly price changes for the years 1996-2001. They construct five datasets, one for each working day. The conclusions on asymmetric pricing are shown to differ over these datasets, suggesting that the choice of the day for which the prices are observed matters more than commonly believed. In their view, the insufficient robustness of the outcomes might explain the mixed conclusions found in the literature. They also show that the effect of asymmetry on the Dutch consumer costs is negligible.

The paper by Galeotti et al. (2003) re-examines the issue of asymmetries in the transmission of shocks to crude oil prices onto the retail price of gasoline. The distinguishing features are: (i) use of updated and comparable data to carry out an international comparison of gasoline markets; (ii) two-stage modeling of the transmission mechanism, in order to assess possible asymmetries at either the refinery stage, the distribution stage or both; (iii) use of asymmetric ECM to distinguish between short-run and long-run asymmetries; (iv) explicit, possibly asymmetric, role of the exchange rate; (v) bootstrapping of F-tests of asymmetries, in order to overcome the low-power problem of conventional testing procedures. In contrast to several previous findings, the results generally point to widespread differences in both adjustment speeds and short-run responses when input prices rise or fall.

The classical menu-cost interpretation, according to which prices are sticky because price menu changes are costly, implies that the probability of a price change should depend on the past history of prices and fundamentals only through the gap between the current price and the frictionless price. Davis and Hamilton (2004) find that this prediction is broadly consistent with the behavior of nine Philadelphia gasoline wholesalers. Nevertheless, they reject the menu-cost model as a literal description of these firms' behaviour, arguing instead that price stickiness arises from strategic considerations of how customers and competitors will react to price changes.

The influence of oil price volatility on the degree of gasoline price asymmetry is studied by Radchenko (2004). The author measures oil price volatility and gasoline price asymmetry and examines the impulse response functions of gasoline price asymmetry to a shock in oil price volatility. His findings suggest a robust negative relationship between the two variables for the American retail market over the period march 1991 - February 2003.

Finally, Kaufmann and Laskowski (2005) analyze monthly data on the American petroleum market for the period January 1986 – December 2002, and use an asymmetric ECM approach. Their results suggest that, when utilization rates and the level of stocks are included in the model, the asymmetry between the price of crude oil and motor gasoline vanishes. Using the same specification of the model, they find asymmetries in the home heating oil market.

To summarize, the vast majority of the articles reported in this survey have studied markets of individual countries. The frequency of the data is typically either weekly or monthly, although sometimes biweekly data are also employed. In general the contributions surveyed consider the lower end of the market, the one in which the product is distributed and sold at the pump. The relevant prices involved are therefore some definition of the wholesale price and the retail price. The other prevailing type of analysis relates the price of crude oil to the pump price within a single, unique stage. Finally, the most recent papers almost invariably test for asymmetric price effects both in the short-run and long-run using dynamic econometric models which exploit the presence of cointegration between the relevant variables.

3. Data and econometric models

In this paper the transmission of changes in upstream prices to downstream prices is investigated at different stages of the process of price formation. We consider the price of crude oil *(CR)* together with the gasoline spot price *(SP)*, the before-tax gasoline retail price *(NR)* and the exchange rate between the U.S. dollar and individual national currencies *(ER)* for five European countries, namely France, Germany, Italy, Spain and U.K.[1] The sample period ranges from January 1985 to March 2003, and the frequency of observations is monthly. All prices are log-transformed and expressed in local currencies, with the exception of crude prices that are denominated in U.S. dollar per barrel.

In particular, the selected crude oil price is the Crude Oil Import Cost (average unit value, c.i.f.), and as a proxy for the ex-refinery gasoline price we use the spot price f.o.b. Rotterdam for the NW Europe. Both prices are from the International Energy Agency. The retail price is obtained as an average of the prices of leaded gasoline and unleaded gasoline. The weight of the first product is equal to one until January 1990 (April 1992 for Spain) and progressively decreases to zero in November 2001 (March 1997 for Germany).[2] The price of leaded gasoline is from the International Energy Agency until June 2000 (March 1997 for Germany) and from DATASTREAM for the remaining part of the sample. The unleaded gasoline price

1 The exchange rate between the U.S. dollar and the Euro is multiplied by the fixed parity for each country after January 1999.

2 This assumption reflects the fact that unleaded gasoline, while virtually absent in the retail market at the beginning of the sample, has become increasingly important during the period spanned by our investigation, and it has been recently the only type of gasoline available at the pump in the countries under analysis.

is from DATASTREAM. The exchanges rates series are obtained from the International Monetary Found for the first portion of the sample and from DATASTREAM since January 1999.

The vast majority of the empirical studies which have been surveyed in Section 2 is based on the concept of cointegration between output and input prices. In the broad class of cointegration models, the most popular specifications for the analysis of price asymmetries are the asymmetric ECM, the threshold ECM, and the ECM with threshold cointegration.

3.1 Asymmetric ECM

If the variables are integrated of order one, or I(1), they may form a linear combination which is stationary, or I(0). The Engle-Granger two-step procedure considers first the relationship among the variables x_j, $j = 1,..,m$, in levels:

$$x_{1t} = \beta_1 + \beta_2 x_{2t} + .. + \beta_m x_{mt} + \varepsilon_t \tag{1}$$

The augmented Dickey-Fuller (ADF) statistic can be used to ascertain whether the residuals, $\hat{\varepsilon}_t$, are stationary.[3] If this is the case, the relevant series are said to be cointegrated. Equation (1) can be considered a steady-state relation among the variables and included in a ECM of the form:

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

with Δ indicating the first difference operator, and p the lag-length.

Granger and Lee (1989) extended the ECM specification to the case of asymmetric adjustments. In order to allow for asymmetries, cointegration residuals and first differences on the *x*'s can be decomposed into positive and negative values. Therefore, model (2) can be written as:

[3] Relevant critical values are available in MacKinnon (1991).

$$\Delta x_{1t} = \alpha^+ \hat{\varepsilon}^+_{t-1} + \alpha^- \hat{\varepsilon}^-_{t-1} + \sum_{i=1}^{p} \lambda_i^+ \Delta x^+_{1t-i} + \sum_{i=1}^{p} \lambda_i^- \Delta x^-_{1t-i} + \sum_{i=0}^{p} \gamma_i^+ \Delta x^+_{2t-i} + \sum_{i=0}^{p} \gamma_i^- \Delta x^-_{2t-i} + \\ + .. + \sum_{i=0}^{p} \delta_i^+ \Delta x^+_{mt-i} + \sum_{i=0}^{p} \delta_i^- \Delta x^-_{mt-i} + u_t \tag{3}$$

The asymmetry in the adjustment speed is introduced by defining $\hat{\varepsilon}_t^+$ equal to $\hat{\varepsilon}_t$ if $\hat{\varepsilon}_t > 0$ and to zero if $\hat{\varepsilon}_t \leq 0$, while $\hat{\varepsilon}_t^-$ equals $\hat{\varepsilon}_t$ or zero when $\hat{\varepsilon}_t < 0$ or $\hat{\varepsilon}_t \geq 0$. Similarly, short-run asymmetry is captured by decomposing the first differences into $\Delta x^+_{jt-i} = x_{jt-i} - x_{jt-i-1} > 0$ and $\Delta x^-_{jt-i} = x_{jt-i} - x_{jt-i-1} < 0$, where $j = 1,..,m$ and $i = 0,.., p$.

Simple inspection of the sign, magnitude and statistical significance of the estimated coefficients offers a first insight on the presence of asymmetric price behaviour. However, in order to establish if the estimated coefficients of model (3) are statistically different, the (single or joint) hypotheses *H₀*: $\alpha^+ = \alpha^-$, $\lambda_i^+ = \lambda_i^-$, $\gamma_i^+ = \gamma_i^-$, .., $\delta_i^+ = \delta_i^-$ have to be formally tested. The asymmetric ECM has often been used as an appropriate framework for conventional F tests of both the hypothesis of symmetric adjustment to the long-run equilibrium and the hypothesis of short-run symmetry. A few recent studies (see Cook et al., 1998, 1999, and Cook, 1999) have shown that standard tests of symmetry are affected by low power in an ECM framework. The solution adopted in this paper is to boostrap the calculated F statistic and obtain the corresponding rejection frequencies via simulation (see also Galeotti et al., 2003).

3.2 Threshold autoregressive ECM

A popular generalization of equation (3) adds a threshold autoregressive (TAR) mechanism to the standard ECM. The resulting model is referred to as the TAR-ECM specification. While it is set to zero in the classical asymmetric ECM, the threshold parameter is consistently estimated using the TAR-ECM.

A two-regime TAR-ECM has the form:

$$\begin{aligned}\Delta x_{1t} &= \alpha\hat{\varepsilon}_{t-1} + \sum_{i=1}^{p}\lambda_i \Delta x_{1t-i} + \sum_{i=0}^{p}\gamma_i \Delta x_{2t-i} + .. + \sum_{i=0}^{p}\delta_i \Delta x_{mt-i} + \\ &+ \left(\alpha^*\hat{\varepsilon}_{t-1} + \sum_{i=1}^{p}\lambda_i^* \Delta x_{1t-i} + \sum_{i=0}^{p}\gamma_i^* \Delta x_{2t-i} + .. + \sum_{i=0}^{p}\delta_i^* \Delta x_{mt-i}\right) 1(q_t > \gamma) + e_t\end{aligned} \tag{4}$$

where p indicates the autoregressive order, q_t is the threshold variable, which is a continuous and stationary transformation of the data, and $\gamma \in \Gamma$ is the threshold parameter.[4] The region denoted by Γ is typically selected by sorting the observations on the threshold variable into an increasing order and by trimming the bottom and top 15% quantiles; the resulting model is well identified for all possible thresholds. The error term e_t is assumed to be a martingale difference sequence. The function $1(.)$ indicates whether or not the threshold variable is above the threshold. The regression coefficients are $(\alpha, \lambda_i, \gamma_i, .., \delta_i)$ if $q_t \le \gamma$, and $(\alpha + \alpha^*, \lambda_i + \lambda_i^*, \gamma_i + \gamma_i^*, .., \delta_i + \delta_i^*)$ if $q_t > \gamma$. Alternatively, if we define $Y_t = (\varepsilon_{t-1} \quad .. \quad \Delta x_{mt-p})'$, $Y_t(\gamma) = (Y_t' \quad Y_t' 1(q_t > \gamma))'$, $\theta_1 = (\alpha, .., \delta_p)'$, $\theta_2 = (\alpha^*, .., \delta_p^*)'$ and $\theta = (\theta_1' \quad \theta_2')'$, model (4) can be expressed as:

$$\Delta x_{1t} = Y_t(\gamma)'\theta + e_t \tag{5}$$

Since equation (5) is non-linear and discontinuous, the parameter estimates can be obtained by sequential conditional least squares. The procedure is as follows: for each possible value of the threshold (i.e. for each $\gamma \in \Gamma$), a regression of the form (5) is estimated with least squares; for each regression, the sum of squared residuals, $S(\gamma)$, is calculated; the threshold's estimate, $\hat{\gamma}$, is the argument that minimizes $S(\gamma)$; the slope estimates are the coefficients $\theta(\hat{\gamma})$ of the corresponding equation (see Hansen, 2000).

It is crucial to test the significance of the threshold autoregressive model (5) relative to the linear model (2). The null hypothesis in this case is $H_0 : \alpha^* = \lambda_i^* = \gamma_i^* = .. = \delta_i^* = 0$ for each i. Defining the selector matrix $R = (0 \quad \mathrm{I})$, $M(\gamma) = \sum Y_t(\gamma) Y_t(\gamma)'$ and $V(\gamma) = \sum Y_t(\gamma) Y_t(\gamma)' \hat{e}_t^2$,

[4] Since the original series are non-stationary, plausible thresholds are the exogenous variables in first differences or the error correction term.

where I is the identity matrix of appropriate dimension, we can write the pointwise heteroskedasticity-consistent Wald statistic as:

$$W(\gamma) = \left(R\hat{\theta}(\gamma)\right)' \left[R\left(M(\gamma)^{-1}V(\gamma)M(\gamma)^{-1}\right)R'\right]^{-1} R\hat{\theta}(\gamma) \tag{6}$$

which leads to the appropriate test statistic:

$$W = \sup_{\gamma \in \Gamma} W(\gamma) \tag{7}$$

The distribution of W in expression (7) is non-standard, as the threshold is not identified under the null hypothesis of linearity. This problem has been analyzed in different contexts by Andrews and Ploberger (1994) and Hansen (1996), among others. In particular, Hansen (1996) suggests a bootstrapping procedure to approximate the asymptotic distribution of (7). This procedure can be implemented as follows: i) draw a sample of random numbers $\eta_t \sim NID(0,1)$ and define $x_t^* = \hat{e}_t \eta_t$; ii) regress x_t^* on Y_t to obtain the restricted sum of squared residuals $\tilde{S}^*$; iii) regress x_t^* on $Y_t(\gamma)$ to obtain the unrestricted sum of squared residuals $S^*(\gamma)$; iv) compute $W^*(\gamma) = T\left(\tilde{S}^* - S^*(\gamma)\right)/S^*(\gamma)$, where T is the number of observations and $W^* = \sup_{\gamma \in \Gamma} W^*(\gamma)$. Repeat steps i)-iv) B times, and denote with W_b^* the calculated statistic corresponding to the b-th iteration. The p-value for W is given by:

$$\text{p-value} = \frac{1}{B}\sum_{b=1}^{B} 1\left(W_b^* \geq W\right)$$

A second relevant issue concerns the significance of the threshold estimate. Consider the null hypothesis $H_0 : \gamma_0 = \gamma$, where γ_0 is the true value and γ is a specified value. A likelihood ratio-type statistic is:

$$LR(\gamma) = T\left(S(\gamma) - S(\hat{\gamma})\right)/S(\hat{\gamma}).$$

This statistic has a non-standard distribution. In case of homoskedasticity, it is possible to show that:

$$LR(\gamma_0) \xrightarrow[d]{} \xi$$

where

$$\xi = \max_{s \in R}\left(2W(s) - |s|\right) \qquad \text{with } W(\nu) = \begin{cases} W_1(-\nu) & \nu < 0 \\ 0 & \nu = 0 \\ W_2(\nu) & \nu > 0 \end{cases}$$

$W_1(-\nu)$ and $W_2(\nu)$ being two independent standard Brownian motions on $[0,\infty)$. Critical values of ξ are reported in Hansen (1997). If the error term is heteroskedastic, the asymptotic distribution depends on a new nuisance parameter, which Hansen (1997) suggests to treat with non-parametric techniques.

3.3 ECM with threshold cointegration

Both asymmetric ECM and TAR-ECM are based on the Engle-Granger two-step approach, that is testing for the presence of cointegration among the relevant price series is implemented via an ADF test on the long-run residuals. However, if the adjustment to the long-run equilibrium is asymmetric, that is, if it depends on the sign of the shocks, the test for cointegration is misspecified (see Balke and Fomby, 1997). In order to overcome this problem, Enders and Granger (1998) replace the standard ADF auxiliary regression with the following TAR process:

$$\Delta\hat{\varepsilon}_t = \mathrm{I}_t \rho_1 \hat{\varepsilon}_{t-1} + (1 - \mathrm{I}_t)\,\rho_2 \hat{\varepsilon}_{t-1} + \nu_t \tag{8}$$

where $\hat{\varepsilon}_t$ are the residuals of the long-run equation (1).

The indicator function I_t is defined to depend on the lagged values of the residuals, according to the following scheme:

$$\mathrm{I}_t = \begin{cases} 1 & if\ \ \hat{\varepsilon}_{t-1} > 0 \\ 0 & if\ \ \hat{\varepsilon}_{t-1} \leq 0 \end{cases} \tag{9}$$

or on the lagged changes in $\hat{\varepsilon}_t$:

$$\mathrm{I}_t = \begin{cases} 1 & if \;\; \Delta\hat{\varepsilon}_{t-1} > 0 \\ 0 & if \;\; \Delta\hat{\varepsilon}_{t-1} \leq 0 \end{cases} \tag{10}$$

Equations (8)-(9) are referred to as TAR cointegration, while model (8)-(10) is named "momentum" TAR (or M-TAR) cointegration. The TAR model is designed to capture potential asymmetric "deep" movements in the residuals, while the M-TAR model is useful to take into account sharp or "steep" variations in $\hat{\varepsilon}_t$ (see Enders and Granger, 1998). As demonstrated by Sichel (1993), negative "deepness" (i.e. $|\rho_1| < |\rho_2|$) of $\hat{\varepsilon}_t$ implies that increases tend to persist, whereas decreases tend to revert quickly towards equilibrium. Since there is generally no presumption on whether to use TAR or M-TAR specifications, it is recommended to choose the appropriate adjustment mechanism via a model selection criterion, such as the Akaike information criterion (AIC).

The test for the presence of a threshold in the equilibrium correction mechanism is termed threshold cointegration test. If $\rho_1 = \rho_2$ the adjustment is symmetric, thus the Engle-Granger approach turns out to be a special case of equations (8) and (9). If the errors are serially correlated, equation (8) can be augmented with the lagged differences of $\hat{\varepsilon}_t$ as in the standard ADF test:

$$\Delta\hat{\varepsilon}_t = \mathrm{I}_t \rho_1 \hat{\varepsilon}_{t-1} + \left(1 - \mathrm{I}_t\right) \rho_2 \hat{\varepsilon}_{t-1} + \sum_{i=1}^{p-1} \sigma_i \Delta\hat{\varepsilon}_{t-i} + \nu_t \tag{11}$$

The threshold parameter does not need to be restricted to zero, as instead it is in models (9) and (10). If the threshold enters the model unrestrictedly, the problem of how to consistently estimate the threshold, or attractor, emerges. Tong (1983) shows that the sample mean of the cointegrating residuals is a biased estimator of the attractor. Chan (1993) demonstrates that a search procedure over all possible values of the attractor in order to minimize the sum of squared residuals yields a super-consistent estimator of the threshold. If, for example, the M-TAR is the selected model according to AIC, equation (10) becomes:

$$\mathrm{I}_t = \begin{cases} 1 & if \;\; \Delta\hat{\varepsilon}_{t-1} > \hat{\mu} \\ 0 & if \;\; \Delta\hat{\varepsilon}_{t-1} \leq \hat{\mu} \end{cases} \tag{12}$$

where $\hat{\mu}$ indicates the consistent estimate of the threshold.

Once equation (11) is estimated, the null hypothesis $H_0 : \rho_1 = \rho_2 = 0$ of no cointegration can be tested through a F test. Correct critical values depend on the number of observations, the number of lags in equation (11) and the number of variables in the cointegrating relationship (see Enders, 2001). The empirical distribution of the F test under the null hypothesis is tabulated for up to five variables, different sample sizes and order of the augmentation in Wane et al. (2004). If the null hypothesis is rejected (i.e. the series $\hat{\varepsilon}_t$ follows a TAR or a M-TAR model), $\hat{\rho}_1$ and $\hat{\rho}_2$ converge to a multivariate normal distribution. Therefore, the hypothesis of symmetric adjustment, i.e. $\rho_1 = \rho_2$, can be tested using a standard F distribution. The corresponding asymmetric error correction representation can be written as:

$$\Delta x_{1t} = \alpha_{up}\hat{\varepsilon}_{t-1}^{up} + \alpha_{down}\hat{\varepsilon}_{t-1}^{down} + \sum_{i=0}^{p}\gamma_i \Delta x_{2t-i} + .. + \sum_{i=0}^{p}\delta_i \Delta x_{mt-i} + \sum_{i=1}^{p}\lambda_i \Delta x_{1t-i} + \xi_t \qquad (13)$$

where $\hat{\varepsilon}_{t-1}^{up} = \mathrm{I}_t \hat{\varepsilon}_{t-1}$ and $\hat{\varepsilon}_{t-1}^{down} = (1 - \mathrm{I}_t)\hat{\varepsilon}_{t-1}$.

4. Empirical results and discussion

We estimate the asymmetric error correction models described in Section 3 to describe the gasoline-price relation in France, Germany, Italy, Spain and UK over the period 1985-2003.

In order to gain a deeper understanding of the movements of gasoline-oil price relation over time, we analyze the transmission of changes in the crude oil price directly to the gasoline price at the pump (single stage), as well as the relations crude spot price-gasoline spot price (first stage) and gasoline spot price-retail gasoline price (second stage). Therefore, three equations are estimated for each model and country.

Tables 1-5 refer to the asymmetric ECM. The estimated coefficients and corresponding t-statistics are reported in Tables 1-3, whereas Tables 4-5 present the results of testing for price asymmetries. Coefficients α^+ and α^- in Table 1 indicate asymmetric adjustment speeds, which measure long-run asymmetry, while the coefficients γ_i^+ and γ_i^-, i=1,…,p, account for short-run, or transitory, asymmetry. The results suggest that “positive” coefficients are generally larger, in absolute value, than their “negative” counterparts for both long-run and short-run, as well as in each stage. This finding is unexpected for long-run effects, where “positive” (α^{+}) and “negative” (α^{-}) coefficients are associated with adjustments to the equilibrium level from above and from below. In contrast, short-run estimates, which show that after two periods the effects of upstream price increases are larger than those of price decreases for all countries, reflect more closely the consumers’ perception of the actual effects of oil price variations on gasoline price changes.

If we concentrate on the two-stage analysis, some additional remarks emerge. First, the magnitude of coefficients is larger in the first stage than in the second stage. Second, lagged effects compensate for the large impact of contemporaneous oil price changes in the refinery stage, while the adjustment towards the equilibrium level is more gradual in the distribution stage. These findings reflect the differences between the refinery and distribution markets. The quotations of spot gasoline react immediately to the fluctuations in the price of oil. In contrast, retailers do not immediately transfer onto pump prices all the adjustments in wholesale prices (and thus in crude oil prices); rather, changes are distributed over time.

A cross-country comparison reveals significant differences, especially at the second stage. The adjustment to the long-run equilibrium appears to be larger from below than from above in the Italian and Spanish distribution markets. In contrast, the systematically larger impact of price increases over price reductions tends to compensate the insignificant adjustment from below to the steady-state level in the retail chain of France and U.K.. Surprisingly, gasoline prices in Germany seem to react more to price decreases and to positive gaps to the equilibrium, than to price increases and negative disequilibrium.

Table 2 considers the transmission of shocks in exchange rates to retail prices. In the first stage, only positive changes appear to be significant, with the only exception of Germany. This evidence suggests that producers are generally reluctant to transfer onto consumers those price reductions which originate from favourable movements in exchange rates. Interestingly, this evidence disappears in the single stage, and it is supportive of the idea of separately modelling production and distribution stages.

The estimated autoregressive coefficients, which enter the model when the lag-length is equal to, or larger than, one, are reported in Table 3. All the estimated coefficients have positive signs in the first stage and are generally negative in the second. Moreover, relevant differences between "positive" and "negative" coefficients, as well as among countries, arise in the second stage. In particular, the coefficients relative to positive lagged changes in gasoline prices are significant and negative for France and Italy, while negative changes are significant and exhibit positive coefficients in the case of U.K. Spain does not show relevant autoregressive asymmetries.

In order to verify whether the differences between the adjustment coefficients and short-run effects are significant, formal statistical testing is required. Table 4 reports the calculated conventional F test for the hypothesis of long-run and short-run asymmetries. Rejection of the null hypothesis H_0: $\alpha^+ = \alpha^-$ implies asymmetric long-run adjustment, whereas short-run asymmetries arise when at least one of the hypotheses H_0: $\gamma_i^+ = \gamma_i^-$, $\delta_i^+ = \delta_i^-$ or $\lambda_i^+ = \lambda_i^-$, $i = 0{,}1$, is rejected.[5] Table 4 shows that long-run asymmetries occur in 3 cases out of 15, while in 8 cases out of 51 short-run asymmetries are significant. If we compare different countries and stages, long-run asymmetries characterize only France and Italy (single stage), and Germany (second stage). The lagged price effects are asymmetric at the first stage in France and Germany, and at the second stage in France, Spain and U.K.. Moreover, the reaction to exchange rate variations is asymmetric in U.K. at the first stage. Finally, contemporaneous price asymmetries arise in France and U.K. at the single stage. Overall, the test suggests the presence of asymmetry in 11 cases, a number which is much smaller than expected, both in terms of how this phenomenon is perceived by the ordinary consumer and from a visual inspection of the estimated coefficients. However, due to the well documented lack of power of the F test in the context of asymmetric ECM, any straightforward interpretation of the results reported in Table 4 may be misleading. Following, among others, Galeotti et al. (2003), we believe that a more reliable picture of potential asymmetries in the oil-gasoline price relation can emerge by bootstrapping the F statistics. Table 5 presents the calculated rejection frequencies at 5% significance level based on 1000 replications. As in Cook et al. (1999), we look at the number of rejection frequencies which are larger than 15% and 58% ("high" rejection frequencies): these amount to 32 and 8 out of 64. In contrast with

[5] In order to economize space, F tests for symmetric short-run effects are reported for contemporaneous and one period lagged changes only.

the standard F tests, the simulated results suggest that each country is more likely to present asymmetries, particularly at the second and single stages.

To summarise, when using the asymmetric ECM approach to describe the price transmission mechanism in the gasoline markets of five European countries, we do find evidence to support the presence of asymmetric price behaviour almost in all countries, and mainly at the distribution stage. As pointed out by Borenstein et al. (1997), retail sales, in contrast with other segments of the oil market, are likely to be characterized by oligopolistic cooperation. Therefore, our results, which evidence that asymmetry is stronger in the second stage, can be explained in terms of reduced competition among retailers.

The two-regime TAR-ECM differs from the asymmetric ECM in two respects: it treats the threshold as an estimable parameter, rather than restricting it to zero, and it accounts only for short-run asymmetries. Tables 6-8 report the estimated value and significance of the coefficients of the TAR-ECM specification. Table 9 presents the estimated values of the threshold parameter, in addition to the calculated Wald statistic for the null hypothesis of no threshold effect and the corresponding approximated p-values. Figures 1-4 plot the adjusted likelihood ratio and the Wald statistics for France (single stage) and Italy (first stage).

An informal indicator of the presence of asymmetries in the oil-gasoline price relation is given by the number of times the estimated coefficients of the error correction term and of the short-run variations differ depending on the sign of short-run price changes, i.e. whether the threshold variable is above or below a specific estimated value. If we consider equation (4), the long-run adjustment is measured by α if the threshold variable is below the estimated threshold, while it is $\alpha+\alpha^*$ otherwise. Similarly, short-run coefficients are $(\lambda_i, \gamma_i, .., \delta_i)$ and $(\lambda_i+\lambda_i^*, \gamma_i+\gamma_i^*, .., \delta_i+\delta_i^*)$. Therefore, significant "differential" parameters $\alpha^*, \gamma_i^*, \delta_i^*$ and λ_i^* suggest the presence of price asymmetries.

Looking at the empirical results presented in Tables 6-8, the coefficients accounting for both long-run and short-run price asymmetries which are statistically significant at 5% are 24 out of 71. If we concentrate on Table 6, significant long-run asymmetries (i.e. α^*) arise in 4 cases out of 15, whereas short-run asymmetries (i.e. γ_i^*, $i = 0,1,2$) are found in 10 cases out of 28.

If we compare the estimated asymmetric coefficients across stages, the main differences are related to the sign of the coefficients γ_1 and to the optimal number of lags in each equation. The lagged short-run effects are negative and contribute to the reduction of the impact of contemporaneous changes in the first stage, while they are positive and tend to increase the cumulative effect of oil and wholesale price changes on gasoline prices in the second and single stage. Moreover, the short-run impact of spot price changes vanishes in one or two periods for the first and single stages, while it is generally distributed over three periods in the second stage. These findings are very close to the results obtained with the asymmetric ECM. Furthermore, it is worthwhile noticing that significant differences in long-run adjustments arise mainly in the second and single stages, while "differential" short-run effects characterize all stages and have positive sign, except for France in the second stage.

Table 7 reports the estimates of the exchange rate effects. All contemporaneous impacts (i.e. δ_0) are significant and positive, while lagged differential effects are positive and statistically significant in the first stage only. Coefficients δ_0^* and δ_0 have opposite signs in all countries and stages, again except for France in the single stage.

The autoregressive coefficients λ_1 reported in Table 8 are significant and positive in the first stage, whereas they are negative and significant in the second stage. In a few cases, autoregressive effects are different depending on the magnitude of contemporaneous changes in oil prices. Spain (second stage) excluded, significant coefficients λ_1^* and λ_1 have opposite signs.

The estimated parameter values depend on the estimated values of the threshold. The latter are calculated using a likelihood ratio approach, after adjusting the LR statistic for heteroskedasticity in the residuals.[6] As an illustration, Figures 1 and 3 present the plots of the adjusted LR against the estimated values of the threshold for France in the single stage and Italy in the first stage, respectively. Values of the threshold corresponding to a LR below the dotted line are not rejected by the data. It is worth observing that the interval of threshold values below the dotted line in Figure 1 is rather tight, while the threshold estimates seem to be less precise in Figure 3. As far as the other countries are concerned, LR plots are well-

[6] This adjustment has been obtained by calculating the LR sequence on the GLS residuals.

shaped (i.e. similar to Figure 1) in about 50% of the cases. The estimates of the threshold are reported in Table 9. Significant and positive threshold values are found in 4 countries, namely France and Germany in the first stage, Italy in the second stage and U.K. in the single stage.

In order to test the null hypothesis of linearity against the threshold model we use a heteroskedasticity-consistent Wald statistic. Figures 2 and 4 display the plots of the statistic against the threshold for France (single stage) and Italy (first stage). The calculated test, along with approximated p-values for each country and stage, are reported in Table 9. Rejection of the null hypothesis of symmetry at 5% significance level occurs for France in the refinery stage, for Germany and Italy in the distribution stage, and for France and Germany in the single stage. In addition, if we test for symmetry at 1% significance level, evidence of asymmetric pricing behaviour is found also for Italy and Spain in the first stage and for France in the second stage.

The overall picture which emerges from the estimation of the threshold ECM is that price asymmetries are present in 34% of the cases. Moreover, asymmetries are more likely a short-run phenomenon (35.7%) than a long-run feature of the oil-gasoline price relation (26.7%). If we compare these findings with the results from the asymmetric ECM (according to which asymmetric price behaviour characterizes only 16% of the cases, with 13.3% of long-run and 16.3% of short-run asymmetries), the TAR-ECM approach turns out to provide stronger support to non-linear pricing schemes in the oil market.

As illustrated in Section 4, a threshold specification of the error correction mechanism is needed to test for threshold cointegration. Tables 10-15 report the results obtained by estimating and testing the threshold cointegrating relationship. Estimates and test statistics are relative to the three possible formulations of the error correction terms, namely TAR, M-TAR and consistent M-TAR (MC-TAR hereafter), and are presented in Tables 10-12. The estimated coefficients of the asymmetric ECM with threshold cointegration are reported in Tables 13-15.

Tables 10-12 show that the M-TAR specification is generally superior to the basic TAR model, at least according to AIC. The sequential conditional OLS method is then used to consistently estimate the threshold parameter for the M-TAR model. Within the MC-TAR specification, the threshold cointegration tests reject the null hypothesis H_0: $\rho_1 = \rho_2 = 0$ in favour of asymmetric cointegration for each country and stage. Moreover, all p-values associated with the tests for the null hypothesis of symmetry are smaller than 5%, supporting the idea of asymmetric adjustments. The reported evidence of asymmetric cointegration leads to the estimation of the ECM with long-run asymmetric equilibrium. Long-run adjustments are allowed to differ depending on the previous period changes in the long-run error terms. The estimated long-run coefficients are presented in Table 12. The most relevant asymmetric effects appear in the single stage. The coefficients α_{down} are all strongly significant and generally larger, in absolute value, than the corresponding α_{up}, which are not even significant for Italy, Spain and U.K. (see Table 13). As for the first stage, all coefficients are significant and, in the case of Italy and Spain, the estimated adjustments from below to the equilibrium exceed the corresponding adjustments from above by more than 0.1. The differences between the estimated coefficients are smaller in the second stage. It is important to point out that, contrary to the asymmetric ECM, the ECM with threshold cointegration identifies long-run asymmetries of the expected sign, that is adjustments from below are found to be faster than adjustments from above.[7] This suggests that a threshold specification of the long-run mechanism provides a more plausible representation of the oil-gasoline price relationship.

If we compare the empirical findings across stages, the magnitude of the adjustment coefficients is larger for the first stage than for the second and single stages. Moreover, as in the cases of asymmetric ECM and threshold ECM, coefficients γ_0 (γ_1) are significant and positive (negative) in the first stage, while contemporaneous price effects are smaller and lagged price effects positive in the other stages. Finally, the temporal delay of the reaction of downstream prices to upstream price changes is larger in the distribution stage than at the refinery level.

Table 14 reports the estimated effects of exchange rate movements on prices. As expected, all coefficients are positive. The effects die out after one period in the first stage, while in two cases lagged effects are significant at the single stage. This behaviour is due to the larger time delay in the reaction of pump prices to cost (and therefore exchange rate) variations. Autoregressive parameters are presented in Table 15. In line with the results obtained by

[7] A comparison with the TAR-ECM, where the threshold variable is the short-run variation of upstream prices, is less informative, thus it is not presented.

estimating the asymmetric ECM and threshold ECM, the autoregressive coefficients are positive in the distribution stage, while, in general, negative in the second stage.

The results of the estimation of the threshold cointegration ECM show strong evidence of asymmetries in the transmission of oil price changes to retail prices (single stage). Adjustments toward the equilibrium between crude oil prices, gasoline retail prices and exchange rates are faster when changes in the deviation from equilibrium are smaller than the estimated threshold.

5. Conclusion

Contrasting evidence about price asymmetries in the oil-product price relationship has been found in the applied econometric literature. Different data, together with different econometric models, have been employed in different studies. One of the major causes of the very large volatility in the empirical findings is the heterogeneity of the econometric approaches used in the empirical applications. Thus, a thorough assessment of the impact of different econometric approaches on the results cannot be put off any longer.

In this paper the three most popular econometric models for price asymmetries are applied to the same dataset, namely asymmetric ECM, threshold ECM, and ECM with threshold cointegration. These models account for different aspects of the potentially asymmetric oil-product price relationship. The asymmetric ECM includes long- and short-run asymmetries, but it forces the threshold to be zero. The threshold ECM tests the existence of short-run asymmetric price behaviour, and it allows to consistently estimate the unknown threshold value. The ECM with threshold cointegration assumes that adjustments toward the long-run equilibrium differ depending on whether changes in the deviation from equilibrium are positive or negative. The dataset we use in the empirical application includes crude oil, spot and retail gasoline prices, together with exchange rates for France, Germany, Italy, Spain and U.K. over the period 1985-2003.

A detailed comparison of the results obtained by estimating each model highlights both similarities and differences. All models are able to find the temporal delay in the reaction of retail prices to changes in spot gasoline and crude oil prices, as well as some evidence of asymmetric behaviour. However, the type of stages and the number of countries which are characterized by asymmetric oil-gasoline price relations vary across models. The asymmetric ECM supports some evidence of asymmetry for all countries, mainly at the distribution stage. The threshold ECM strongly rejects the null hypothesis of symmetric pricing behaviour, particularly in the case of France (all stages) and Germany (distribution level). Finally, the ECM with threshold cointegration captures long-run asymmetry for each country in the reaction of retail prices directly to oil price changes.

Table 1. Asymmetric ECM - asymmetric adjustment speeds and short-run price asymmetries

	France	Germany	Italy	Spain	U.K.
			first stage: spot=f(crude, exchange rate)		
LR asymm. α^+	-0.374 (-4.667)	-0.373 (-4.609)	-0.305 (-4.577)	-0.268 (-3.653)	-0.261 (-3.515)
LR asymm. α^-	-0.254 (-2.702)	-0.274 (-2.826)	-0.231 (-2.702)	-0.286 (-3.392)	-0.242 (-2.509)
SR asymm. γ_0^+	0.822 (8.440)	0.823 (8.368)	0.881 (10.195)	0.910 (9.121)	0.819 (9.083)
SR asymm. γ_0^-	0.919 (9.109)	0.842 (8.418)	0.899 (9.926)	0.720 (7.595)	0.736 (7.832)
SR asymm. γ_1^+	-0.152 (-1.426)	-0.088 (-0.800)	-0.281 (-2.766)	-0.205 (-1.868)	-
SR asymm. γ_1^-	-0.599 (-4.826)	-0.523 (-4.388)	-0.601 (-5.488)	-0.462 (-4.179)	-
			second stage: retail=f(spot)		
LR asymm. α^+	-0.162 (-2.588)	-0.660 (-6.121)	0.001 (0.022)	-0.052 (-0.888)	-0.231 (-3.273)
LR asymm. α^-	-0.065 (-0.970)	-0.272 (-3.101)	-0.180 (-3.489)	-0.257 (-3.438)	-0.086 (-1.568)
SR asymm. γ_0^+	0.191 (3.465)	0.293 (3.956)	0.090 (2.634)	0.094 (2.271)	0.175 (3.348)
SR asymm. γ_0^-	0.119 (2.092)	0.339 (4.545)	0.139 (3.902)	0.184 (4.236)	0.065 (1.167)
SR asymm. γ_1^+	0.545 (8.723)	-	0.372 (8.501)	0.242 (4.506)	0.394 (6.337)
SR asymm. γ_1^-	0.329 (5.239)	-	0.371 (8.679)	0.422 (8.493)	0.182 (2.949)
SR asymm. γ_2^+	0.271 (3.524)	-	0.177 (3.173)	0.096 (1.742)	-
SR asymm. γ_2^-	0.161 (2.298)	-	0.176 (3.375)	0.174 (2.925)	-
SR asymm. γ_3^+	-	-	0.032 (0.612)	0.111 (2.093)	-
SR asymm. γ_3^-	-	-	0.189 (3.716)	0.080 (1.405)	-
			single stage: retail=f(crude, exchange rate)		
LR asymm. α^+	-0.454 (-4.572)	-0.406 (-3.673)	-0.229 (-3.412)	-0.237 (-2.825)	-0.165 (-2.383)
LR asymm. α^-	-0.180 (-1.865)	-0.309 (-3.352)	0.009 (0.226)	-0.167 (-2.175)	-0.154 (-2.634)
SR asymm. γ_0^+	0.439 (5.598)	0.406 (4.456)	0.263 (4.285)	0.184 (2.991)	0.277 (4.193)
SR asymm. γ_0^-	-0.012 (-0.139)	0.383 (3.992)	0.258 (3.955)	0.110 (1.821)	0.045 (0.629)
SR asymm. γ_1^+	0.244 (2.772)	-	-	0.196 (3.126)	0.213 (2.867)
SR asymm. γ_1^-	0.261 (2.807)	-	-	0.261 (4.271)	0.240 (3.265)

Notes: LR = long-run; SR = short-run; parameters α^+, α^-, γ_i^+ and γ_i^- refer to equation (3), where *m*=3 and $x_1=SP$, $x_2=CR$, $x_3=ER$ for the first stage; *m*=2, $x_1=NR$ and $x_2=SP$ for the second stage; *m*=3, $x_1=NR$, $x_2=CR$ and $x_3=ER$ for the single stage. For each parameter the estimated value and t-ratio (in brackets) are reported. The optimal number of lags in the asymmetric ECM is chosen to eliminate any residual autocorrelation. A "-" in correspondence to the *i*-th lag (*i*=1,2,3) indicates that the optimal number of lags is *i*-1.

Table 2. Asymmetric ECM - exchange rate asymmetries

	France	Germany	Italy	Spain	U.K.
	first stage: spot=f(crude, exchange rate)				
SR asymm. δ_0^+	1.170 (3.466)	1.112 (3.344)	1.098 (4.163)	1.235 (4.020)	1.673 (5.414)
SR asymm. δ_0^-	0.458 (1.544)	0.578 (2.015)	0.326 (1.112)	0.435 (1.328)	0.119 (0.385)
	single stage: retail=f(crude, exchange rate)				
SR asymm. δ_0^+	0.512 (1.804)	-0.217 (-0.655)	0.090 (0.436)	0.203 (1.011)	0.531 (2.303)
SR asymm. δ_0^-	0.605 (2.382)	0.501 (1.759)	0.683 (3.025)	0.184 (0.885)	-0.033 (-0.149)
SR asymm. δ_1^+	0.254 (0.919)	-	-	0.560 (2.807)	0.597 (2.566)
SR asymm. δ_1^-	-0.086 (-0.330)	-	-	0.311 (1.496)	0.148 (0.654)

Notes: LR = long-run; SR = short-run; parameters δ_i^+ and δ_i^- refer to equation (3), where m=3, x_1=SP, x_2=CR and x_3=ER for the first stage; m=2, x_1=NR and x_2=SP for the second stage; m=3, x_1=NR, x_2=CR and x_3=ER for the single stage. For each parameter the estimated value and t-ratio (in brackets) are reported. A "-" in correspondence to the i-th lag (i=1,2,3) indicates that the optimal number of lags is i-1.

Table 3. Asymmetric ECM - autoregressive asymmetries

	France	Germany	Italy	Spain	U.K.
	first stage: spot=f(crude, exchange rate)				
SR asymm. λ_1^+	0.220 (2.252)	0.201 (1.988)	0.305 (3.259)	0.209 (2.108)	-
SR asymm. λ_1^-	0.310 (3.085)	0.286 (2.875)	0.293 (3.096)	0.270 (2.731)	-
	second stage: retail=f(spot)				
SR asymm. λ_1^+	-0.458 (-4.499)	-	-0.324 (-3.239)	-0.197 (-2.064)	0.055 (0.636)
SR asymm. λ_1^-	-0.178 (-1.861)	-	-0.110 (-1.048)	-0.304 (-2.949)	0.314 (3.590)
SR asymm. λ_2^+	-0.220 (-2.710)	-	-0.294 (-2.956)	-0.164 (-1.742)	-
SR asymm. λ_2^-	0.167 (2.217)	-	-0.118 (-1.185)	-0.027 (-0.269)	-
	single stage: retail=f(crude, exchange rate)				
SR asymm. λ_1^+	-0.025 (-0.239)	-	-	-	0.100 (1.014)
SR asymm. λ_1^-	0.108 (1.180)	-	-	-	0.263 (2.635)

Notes: LR = long-run; SR = short-run; parameters λ_i^+ and λ_i^- refer to equation (3), where m=3, x_1=SP, x_2=CR and x_3=ER for the first stage; m=2, x_1=NR and x_2=SP for the second stage; m=3, x_1=NR, x_2=CR and x_3=ER for the single stage. For each parameter the estimated value and t-ratio (in brackets) are reported. A "-" in correspondence to the i-th lag (i=1,2,3) indicates that the optimal number of lags is i-1.

Table 4. Asymmetric ECM - computed F tests for asymmetric adjustment speeds and short-run effects

Null hypothesis	France	Germany	Italy	Spain	U.K.
	first stage: spot=f(crude, exchange rate)				
$\alpha^+ = \alpha^-$	0.666 (0.415)	0.446 (0.504)	0.342 (0.559)	0.020 (0.889)	0.018 (0.894)
$\gamma_0^+ = \gamma_0^-$	0.350 (0.554)	0.015 (0.904)	0.016 (0.898)	1.407 (0.236)	0.302 (0.582)
$\gamma_1^+ = \gamma_1^-$	5.957 (0.015)	5.708 (0.017)	3.795 (0.051)	2.233 (0.135)	-
$\delta_0^+ = \delta_0^-$	1.714 (0.190)	1.019 (0.313)	2.693 (0.101)	2.165 (0.141)	9.046 (0.003)
$\lambda_1^+ = \lambda_1^-$	0.335 (0.563)	0.291 (0.589)	0.007 (0.934)	0.160 (0.689)	-
	second stage: retail=f(spot)				
$\alpha^+ = \alpha^-$	0.862 (0.353)	5.494 (0.019)	3.438 (0.064)	3.479 (0.062)	1.846 (0.174)
$\gamma_0^+ = \gamma_0^-$	0.609 (0.435)	0.141 (0.707)	0.749 (0.387)	1.644 (0.200)	1.520 (0.218)
$\gamma_1^+ = \gamma_1^-$	4.937 (0.026)	-	9.17E-05 (0.992)	5.172 (0.023)	4.415 (0.036)
$\lambda_1^+ = \lambda_1^-$	3.803 (0.051)	-	1.918 (0.166)	0.560 (0.454)	3.339 (0.068)
	single stage: retail=f(crude, exchange rate)				
$\alpha^+ = \alpha^-$	2.809 (0.094)	0.318 (0.573)	6.363 (0.012)	0.265 (0.607)	0.011 (0.917)
$\gamma_0^+ = \gamma_0^-$	11.423 (0.001)	0.021 (0.886)	0.002 (0.963)	0.542 (0.462)	4.328 (0.038)
$\gamma_1^+ = \gamma_1^-$	0.015 (0.904)	-	-	0.429 (0.512)	0.052 (0.819)
$\delta_0^+ = \delta_0^-$	0.041 (0.840)	1.851 (0.174)	2.653 (0.103)	0.003 (0.955)	2.247 (0.134)
$\delta_1^+ = \delta_1^-$	0.562 (0.454)	-	-	0.522 (0.470)	1.399 (0.237)
$\lambda_1^+ = \lambda_1^-$	0.772 (0.380)	-	-	-	1.055 (0.304)

Notes: entries are the calculated F tests for the null hypothesis of symmetry, i.e. equality between the coefficients associated with error correction terms, price changes and exchange rate changes in equation (3), and the corresponding p-values (in brackets). Tests for symmetry are reported only for the long-run adjustments, contemporaneous and one period lagged changes. A "-" in correspondence to the *i*-th lag (*i*=1,2,3) indicates that the optimal number of lags is *i*-1.

Table 5. Asymmetric ECM – simulated F tests for asymmetric adjustment speeds and short-run effects

Null hypothesis	France	Germany	Italy	Spain	U.K.
	first stage: spot=f(crude, exchange rate)				
$\alpha^+ = \alpha^-$	0.133	0.117	0.094	0.065	0.054
$\gamma_0^+ = \gamma_0^-$	0.092	0.052	0.042	0.228	0.090
$\gamma_1^+ = \gamma_1^-$	0.709	0.688	0.503	0.321	-
$\delta_0^+ = \delta_0^-$	0.273	0.170	0.400	0.340	0.864
$\lambda_1^+ = \lambda_1^-$	0.085	0.085	0.056	0.065	-
	second stage: retail=f(spot)				
$\alpha^+ = \alpha^-$	0.165	0.669	0.461	0.480	0.299
$\gamma_0^+ = \gamma_0^-$	0.142	0.065	0.141	0.256	0.236
$\gamma_1^+ = \gamma_1^-$	0.627	-	0.059	0.641	0.577
$\lambda_1^+ = \lambda_1^-$	0.505	-	0.311	0.117	0.459
	single stage: retail=f(crude, exchange rate)				
$\alpha^+ = \alpha^-$	0.412	0.107	0.734	0.081	0.045
$\gamma_0^+ = \gamma_0^-$	0.926	0.061	0.05	0.130	0.557
$\gamma_1^+ = \gamma_1^-$	0.067	-	-	0.101	0.055
$\delta_0^+ = \delta_0^-$	0.055	0.28	0.368	0.045	0.331
$\delta_1^+ = \delta_1^-$	0.116	-	-	0.105	0.234
$\lambda_1^+ = \lambda_1^-$	0.145	-	-	-	0.165

Notes: entries are the simulated rejection frequencies, i.e. the percentage number of rejections (out of 1,000 replications) of the null hypothesis of symmetry using a F test at 5% significance level. A "-" in correspondence to the *i*-th lag (*i*=1,2,3) indicates that the optimal number of lags is *i*-1.

Table 6. TAR-ECM – two-regime adjustment speeds and short-run price effects

	France	Germany	Italy	Spain	U.K.
	first stage: spot=f(crude, exchange rate)				
LR effect α	-0.277 (-5.162)	-0.305 (-5.560)	-0.252 (-5.084)	-0.266 (-5.465)	-0.220 (-1.938)
LR "differential" effect $\alpha*$	-0.199 (-1.788)	-0.147 (-1.191)	-0.061 (-0.664)	0.043 (0.439)	-0.057 (-0.457
SR effect γ_0	0.901 (11.790)	0.845 (11.366)	0.920 (11.538)	0.750 (10.359)	0.938 (5.133)
SR "differential" eff γ_0^*	0.327 (1.650)	0.444 (2.134)	0.272 (1.706)	0.571 (2.846)	-0.067 (-0.337)
SR effect γ_1	-0.375 (-4.571)	-0.329 (-4.057)	-0.464 (-5.344)	-0.304 (-3.886)	-0.320 (-2.539)
SR "differential" effect γ_1^*	-0.005 (-0.030)	0.026 (0.150)	0.002 (0.012)	-0.217 (-1.134)	0.204 (1.323)
	Second stage: retail=f(spot)				
LR effect α	0.109 (1.626)	-0.200 (-2.429)	-0.117 (-3.588)	-0.196 (-4.474)	-0.163 (-4.508)
LR "differential" effect $\alpha*$	-0.296 (-3.689)	-0.383 (-3.657)	0.061 (0.723)	0.196 (2.398)	0.125 (1.305)
SR effect γ_0	0.201 (2.217)	0.498 (5.056)	0.156 (5.663)	0.132 (2.992)	0.132 (2.976)
SR "differential" effect γ_0^*	-0.010 (-0.093)	-0.191 (-1.529)	-0.060 (-0.836)	-0.115 (-1.546)	0.346 (2.796)
SR effect γ_1	0.645 (9.012)	-	0.294 (10.448)	0.285 (7.188)	0.259 (6.359)
SR "differential" effect γ_1^*	-0.270 (-3.191)	-	0.265 (4.624)	0.155 (2.369)	0.111 (1.311)
SR effect γ_2	0.409 (5.837)	-	0.123 (3.626)	0.096 (2.361)	-
SR "differential" effect γ_2^*	-0.340 (-3.846)	-	-0.034 (-0.277)	0.121 (1.592)	-
	Single stage: retail=f(crude, exchange rate)				
LR effect α	-0.383 (-4.557)	-0.298 (-5.330)	-0.204 (-2.832)	-0.777 (-5.131)	-0.197 (-5.005)
LR "differential" effect $\alpha*$	0.100 (0.946)	-0.247 (-1.746)	0.149 (1.947)	0.525 (3.330)	-0.078 (-1.097)
SR effect γ_0	0.101 (1.042)	0.372 (5.266)	0.417 (2.188)	0.329 (2.081)	0.197 (3.295)
SR "differential" effect γ_0^*	0.413 (3.158)	0.138 (0.746)	-0.159 (-0.805)	-0.127 (-0.766)	0.325 (2.802)
SR effect γ_1	0.090 (1.148)	-	-	-	-
SR "differential" effect γ_1^*	0.235 (2.201)	-	-	-	-

Notes: LR = long-run; SR = short-run; parameters α, α^*, γ_i and γ_i^* refer to equation (4), where m=3, $x_1=SP$, $x_2=CR$ and $x_3=ER$ for the first stage; m=2, $x_1=NR$ and $x_2=SP$ for the second stage; m=3, $x_1=NR$, $x_2=CR$ and $x_3=ER$ for the single stage. For each parameter the estimated value and t-ratio (in brackets) are reported. Reported t-ratios need to be compared with critical values of the normal distribution. A "-" in correspondence to the i-th lag (i=1,2,3) indicates that the optimal number of lags is i-1.

Table 7. TAR-ECM – two-regime exchange rate effects

	France	Germany	Italy	Spain	U.K.
	first stage: spot=f(crude, exchange rate)				
SR effect δ_0	1.025 (5.786)	0.993 (5.513)	0.839 (4.784)	0.946 (5.401)	1.022 (2.629)
SR "differential" effect δ_0^*	-1.140 (-2.321)	-0.571 (-1.262)	-0.455 (-1.194)	-0.424 (-0.791)	-0.203 (-0.464)
SR effect δ_1	-0.170 (-0.899)	-0.195 (-1.046)	-0.173 (-0.916)	-0.248 (-1.344)	-0.745 (-1.766)
SR "differential" effect δ_1^*	0.782 (1.670)	1.144 (2.158)	0.701 (1.825)	1.560 (2.885)	0.792 (1.683)
	single stage: retail=f(crude, exchange rate)				
SR effect δ_0	0.537 (3.007)	0.448 (2.705)	1.096 (3.163)	0.430 (1.104)	0.463 (3.308)
SR "differential" effect δ_0^*	0.209 (0.773)	-1.464 (-3.384)	-0.843 (-2.291)	-0.181 (-0.445)	-0.825 (-2.465)
SR effect δ_1	0.023 (0.114)	-	-	-	-
SR "differential" effect δ_1^*	0.104 (0.380)	-	-	-	-

Notes: LR = long-run; SR = short-run; parameters δ_i and δ_i^* refer to equation (4), where m=3, $x_1=SP$, $x_2=CR$ and $x_3=ER$ for the first stage; m=2, $x_1=NR$ and $x_2=SP$ for the second stage; m=3, $x_1=NR$, $x_2=CR$ and $x_3=ER$ for the single stage. For each parameter the estimated value and t-ratio (in brackets) are reported. Reported t-ratios need to be compared with critical values of the normal distribution. A "-" in correspondence to the i-th lag (i=1,2,3) indicates that the optimal number of lags is i-1.

Table 8. TAR-ECM – two-regime autoregressive effects

	France	Germany	Italy	Spain	U.K.
	first stage: spot=f(crude, exchange rate)				
SR effect λ_1	0.202 (2.836)	0.206 (2.852)	0.278 (3.769)	0.240 (3.346)	0.414 (3.211)
SR "differential" effect λ_1^*	0.191 (1.192)	0.198 (1.161)	0.064 (0.438)	0.046 (0.255)	-0.339 (-2.228)
	second stage: retail=f(spot)				
SR effect λ_1	-0.815 (-6.936)	-	-0.132 (-2.054)	-0.157 (-2.010)	0.159 (2.914)
SR "differential" effect λ_1^*	0.751 (5.346)	-	0.224 (0.939)	-0.232 (-1.623)	0.127 (0.879)
SR effect λ_2	-	-	-0.076 (-1.818)	-	-
SR "differential" effect λ_2^*	-	-	0.209 (2.372)	-	-
	single stage: retail=f(crude, exchange rate)				
SR effect λ_1	0.355 (3.564)	-	-	-	0.302 (4.815)
SR "differential" effect λ_1^*	-0.535 (-4.322)	-	-	-	0.110 (0.826)

Notes: LR = long-run; SR = short-run; parameters λ_i and λ_i^* refer to equation (4), where m=3, $x_1=SP$, $x_2=CR$ and $x_3=ER$ for the first stage; m=2, $x_1=NR$ and $x_2=SP$ for the second stage; m=3, $x_1=NR$, $x_2=CR$ and $x_3=ER$ for the single stage. For each parameter the estimated value and t-ratio (in brackets) are reported. Reported t-ratios need to be compared with critical values of the normal distribution. A "-" in correspondence to the i-th lag (i=1,2,3) indicates that the optimal number of lags is i-1.

Table 9. TAR-ECM – estimated thresholds and computed Wald tests

	France	Germany	Italy	Spain	U.K.
	first stage: spot=f(crude, exchange rate)				
Threshold γ	0.062*	0.073*	0.051	0.073	-0.050
Wald test	30.521	18.909	23.450	24.615	10.787
p-value	0.027	0.206	0.079	0.086	0.779
	second stage: retail=f(spot)				
Threshold γ	-0.039*	-0.009	0.071*	0.024	0.069
Wald test	25.565	15.175	27.618	20.024	12.856
p-value	0.069	0.041	0.023	0.119	0.287
	single stage: retail=f(crude, exchange rate)				
Threshold γ	0.002	0.071	-0.081	-0.085	0.051*
Wald test	30.092	26.514	12.731	13.644	22.961
p-value	0.040	0.005	0.213	0.169	0.041

Notes: A"*" indicates statistical significance at 5%. The calculated Wald statistics are testing the null hypothesis of linear ECM against the alternative of ECM with threshold specification. The asymptotic p-values of the tests are obtained via bootstrapping (1,000 replications).

Table 10. TAR, M-TAR and MC cointegrating relations - first stage

	France			Germany			Italy			Spain			U.K.		
	TAR	M-TAR	MC	TAR	M-TAR	MC	TAR	M-TAR	MC	TAR	M-TAR	MC	TAR	M-TAR	MC
ρ_1	-0.324	-0.341	-0.198	-0.329	-0.347	-0.203	-0.269	-0.289	-0.153	-0.253	-0.259	-0.179	-0.295	-0.299	-0.151
	(-5.424)	(-5.356)	(-3.589)	(-5.430)	(-5.230)	(-3.596)	(-5.130)	(-5.077)	(-3.320)	(-4.627)	(-4.572)	(-3.860)	(-4.999)	(-4.412)	(-2.787)
ρ_2	-0.316	-0.300	-0.546	-0.335	-0.316	-0.561	-0.265	-0.245	-0.555	-0.264	-0.256	-0.514	-0.267	-0.271	-0.535
	(-4.508)	(-4.576)	(-7.462)	(-4.636)	(-4.832)	(-7.612)	(-4.151)	(-4.239)	(-7.710)	(-4.264)	(-4.312)	(-6.301)	(-3.765)	(-4.386)	(-7.323)
AIC	-2.649	-2.650	-2.718	-2.630	-2.630	-2.702	-2.777	-2.779	-2.878	-2.609	-2.609	-2.668	-2.623	-2.623	-2.707
$\rho_1 = \rho_2 = 0$	23.436	23.557	32.784	23.983	24.052	33.754	20.742	20.928	34.364	18.823	18.811	26.555	18.502	18.505	29.552
$\rho_1 = \rho_2$	0.008	0.207	15.346	0.004	0.115	15.968	0.003	0.314	22.825	0.021	0.001	13.172	0.096	0.101	18.946
	[0.929]	[0.649]	[1E-04]	[0.951]	[0.734]	[1E-04]	[0.954]	[0.576]	[0.000]	[0.886]	[0.975]	[4E-04]	[0.757]	[0.750]	[0.000]

Table 11. TAR, M-TAR and MC cointegrating relations - second stage

	France			Germany			Italy			Spain			U.K.		
	TAR	M-TAR	MC	TAR	M-TAR	MC	TAR	M-TAR	MC	TAR	M-TAR	MC	TAR	M-TAR	MC
ρ_1	-0.278	-0.267	-0.064	-0.322	-0.240	-0.176	-0.158	-0.158	-0.521	-0.175	-0.241	-0.540	-0.271	-0.207	-0.474
	(-3.674)	(-3.577)	(-1.007)	(-3.516)	(-2.948)	(-2.581)	(-2.025)	(-2.063)	(-5.730)	(-2.563)	(-3.321)	(-5.497)	(-3.927)	(-3.254)	(-5.845)
ρ_2	-0.170	-0.178	-0.598	-0.211	-0.273	-0.556	-0.250	-0.255	-0.098	-0.279	-0.207	-0.136	-0.244	-0.304	-0.170
	(-2.153)	(-2.221)	(-6.715)	(-2.734)	(-3.132)	(-4.677)	(-3.659)	(-3.637)	(-1.661)	(-3.801)	(-2.939)	(-2.419)	(-4.116)	(-4.760)	(-3.247)
AIC	-2.486	-2.484	-2.605	-2.719	-2.715	-2.755	-3.071	-3.072	-3.146	-3.031	-3.026	-3.091	-2.698	-2.703	-2.746
$\rho_1 = \rho_2 = 0$	8.068	7.867	22.544	8.457	7.938	12.675	7.877	7.931	16.763	9.426	8.834	16.542	15.267	15.904	21.242
$\rho_1 = \rho_2$	1.142	0.767	28.191	1.060	0.092	8.910	0.897	0.999	17.511	1.224	0.131	14.364	0.100	1.214	10.561
	[0.286]	[0.382]	[0.000]	[0.305]	[0.762]	[0.003]	[0.345]	[0.319]	[0.000]	[0.270]	[0.718]	[2E-04]	[0.753]	[0.272]	[0.001]

Table 12. TAR, M-TAR and MC cointegrating relations - single stage

	France			Germany			Italy			Spain			U.K.		
	TAR	M-TAR	MC	TAR	M-TAR	MC	TAR	M-TAR	MC	TAR	M-TAR	MC	TAR	M-TAR	MC
ρ_1	-0.423	-0.414	-0.204	-0.380	-0.282	-0.920	-0.087	-0.058	-0.279	-0.303	-0.265	-0.721	-0.192	-0.134	-0.358
	(-5.803)	(-5.878)	(-3.165)	(-4.713)	(-3.776)	(-7.935)	(-1.854)	(-1.543)	(-5.210)	(-4.499)	(-3.898)	(-7.124)	(-3.347)	(-2.346)	(-5.321)
ρ_2	-0.306	-0.315	-0.637	-0.352	-0.450	-0.236	-0.099	-0.140	-0.032	-0.232	-0.266	-0.160	-0.154	-0.234	-0.111
	(-3.987)	-(3.995)	(-7.658)	(-4.943)	(-5.935)	(-4.239)	(-2.860)	(-3.402)	(-1.034)	(-3.644)	(-4.194)	(-3.294)	(-3.042)	(-4.413)	(-2.393)
AIC	-2.793	-2.805	-2.862	-2.606	-2.613	-2.729	-3.135	-3.146	-3.207	-3.191	-3.185	-3.298	-2.879	-2.882	-2.918
$\rho_1 = \rho_2 = 0$	24.784	25.252	34.331	23.325	24.741	40.466	5.725	6.876	14.076	16.763	16.394	30.799	10.226	12.104	16.462
$\rho_1 = \rho_2$	1.233	0.871	16.852	0.071	2.476	28.273	0.040	2.225	15.889	0.584	0.000	24.951	0.242	1.691	9.583
	[0.268]	[0.352]	[1E-04]	[0.790]	[0.117]	[0.000]	[0.842]	[0.137]	[1E-04]	[0.446]	[0.997]	[0.000]	[0.623]	[0.195]	[0.002]

Notes to Tables 10-12: entires for parameters ρ_1 and ρ_2 are the estimated values and t-ratios (in round brackets); AIC=Akaike Information Criterion; $\rho_1=\rho_2=0$ = null hypothesis for the threshold cointegration test [for critical values relative to TAR and M-TAR models refer to Wane et al. (2004), for the MC model refer to Enders (2001) ; $\rho_1=\rho_2$:= null hypothesis for the test of symmetry (p-values are reported in squared brackets).

Table 13. ECM with threshold cointegration - asymmetric adjustment speeds and short-run price effects

	France	Germany	Italy	Spain	U.K.
			first stage: spot=f(crude, exchange rate)		
LR asymm. α_{up}	-0.312 (-6.017)	-0.303 (-5.617)	-0.238 (-5.421)	-0.216 (-4.823)	-0.283 (-5.389)
LR asymm. α_{down}	-0.279 (-2.947)	-0.349 (-3.448)	-0.332 (-3.442)	-0.402 (-4.093)	-0.222 (-2.213)
SR effect γ_0	0.865 (15.452)	0.836 (14.972)	0.896 (17.983)	0.821 (15.081)	0.804 (14.714)
SR effect γ_1	-0.349 (-4.913)	-0.286 (-4.040)	-0.421 (-6.265)	-0.329 (-4.775)	-0.187 (-2.607)
			second stage: retail=f(spot)		
LR asymm. α_{up}	-0.122 (-2.980)	-0.216 (-3.891)	-0.068 (-0.976)	-0.203 (-2.745)	-0.174 (-2.579)
LR asymm. α_{down}	-0.085 (-1.077)	-0.231 (-2.387)	-0.083 (-2.431)	-0.123 (-3.276)	-0.141 (-3.871)
SR effect γ_0	0.162 (4.973)	0.237 (6.535)	0.113 (5.529)	0.132 (5.557)	0.134 (4.379)
SR effect γ_1	0.442 (10.901)	0.450 (9.426)	0.388 (14.328)	0.344 (11.183)	0.281 (7.813)
SR effect γ_2	0.180 (4.012)	-	0.151 (4.092)	0.108 (3.176)	-
SR effect γ_3	-	-	0.103 (2.963)	0.057 (2.382)	-
			single stage: retail=f(crude, exchange rate)		
LR asymm. α_{up}	-0.274 (-4.855)	-0.378 (-2.360)	-0.070 (-1.156)	-0.188 (-1.832)	-0.103 (-1.609)
LR asymm. α_{down}	-0.369 (-3.855)	-0.357 (-6.305)	-0.078 (-2.810)	-0.206 (-4.682)	-0.158 (-4.157)
SR asymm. γ_0	0.189 (3.989)	0.389 (7.554)	0.263 (7.263)	0.144 (4.209)	0.177 (4.532)
SR asymm. γ_1	0.309 (5.815)	-	-	0.231 (6.405)	0.224 (5.098)
SR asymm. γ_2	-0.128 (-2.711)	-	-	-	-

Notes: parameters α_{up}, α_{down} and γ_i refer to equation (13), where m=3, $x_1=SP$, $x_2=CR$ and $x_3=ER$ for the first stage; m=2, $x_1=NR$ and $x_2=SP$ for the second stage; m=3, $x_1=NR$, $x_2=CR$ and $x_3=ER$ for the single stage. For each parameter the estimated value and t-ratio (in brackets) are reported. A "-" in correspondence to the i-th lag (i=1,2,3) indicates that the optimal number of lags is i-1.

Table 14. ECM with threshold cointegration – exchange rate effects

	France	Germany	Italy	Spain	U.K.
	first stage: spot=f(crude, exchange rate)				
SR effect δ_0	0.824 (4.997)	0.818 (5.036)	0.741 (4.943)	0.826 (5.090)	0.832 (4.812)
	single stage: retail=f(crude, exchange rate)				
SR asymm. δ_0	0.624 (4.623)	0.194 (1.217)	0.387 (3.276)	0.197 (1.855)	0.243 (1.903)
SR asymm. δ_1	-	-	-	0.442 (4.166)	0.368 (2.806)

Notes: parameters δ_i refer to equation (13), where m=3, x_1=*SP*, x_2=*CR* and x_3=*ER* for the first stage; m=2, x_1=*NR* and x_2=*SP* for the second stage; m=3, x_1=*NR*, x_2=*CR* and x_3=*ER* for the single stage. For each parameter the estimated value and t-ratio (in brackets) are reported. A "-" in correspondence to the *i*-th lag (*i*=1,2,3) indicates that the optimal number of lags is *i*-1.

Table 15. ECM with threshold cointegration – autoregressive effects

	France	Germany	Italy	Spain	U.K.
	first stage: spot=f(crude, exchange rate)				
SR asymm. λ_1	0.237 (3.841)	0.225 (3.586)	0.282 (4.749)	0.241 (3.841)	0.153 (2.347)
	second stage: retail=f(spot)				
SR asymm. λ_1	-0.275 (-4.100)	-0.192 (-3.304)	-0.198 (-2.846)	-0.222 (-3.448)	0.197 (3.816)
SR asymm. λ_2	-	-	-0.182 (-2.696)	-	-
	single stage: retail=f(crude, exchange rate)				
SR asymm. λ_1	-	-	-	-	0.184 (3.105)

Notes: parameters λ_i are the corresponding coefficients in equation (13), where m=3 and x_1=*SP*, x_2=*CR* and x_3=*ER* for the first stage; m=2, x_1=*NR* and x_2=*SP* for the second stage; m=3, x_1=*NR*, x_2=*CR* and x_3=*ER* for the single stage. For each parameter the estimated value and t-ratio (in brackets) are reported. A "-" in correspondence to the *i*-th lag (*i*=1,2,3) indicates that the optimal number of lags is *i*-1.

Figure 1. Likelihood ratio test for the threshold - France (single stage)

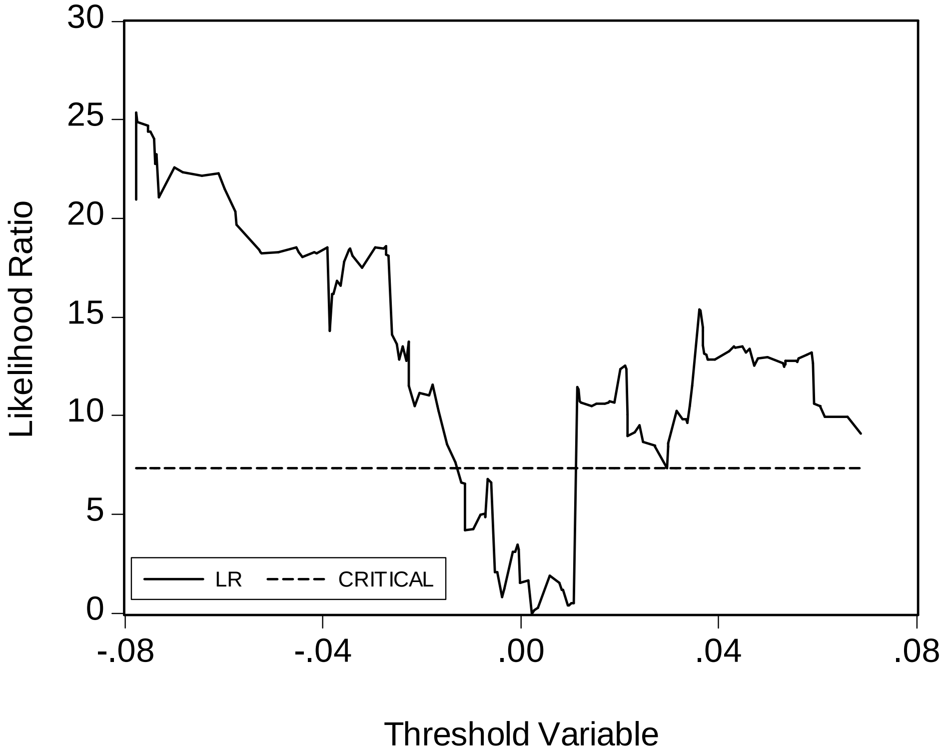


Figure 2. Heteroskedasticity-consistent Wald test – France (single stage)

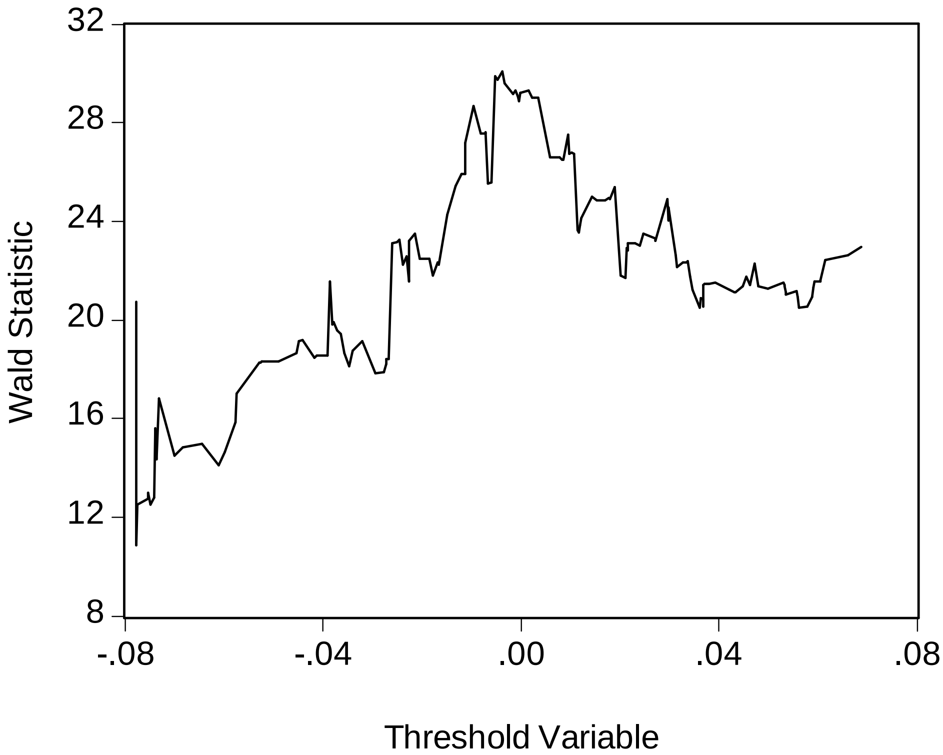

32
28
24
20
16
12
8
Wald Statistic
-.08
-.04
.00
.04
.08
Threshold Variable

Figure 3. Likelihood ratio test for the threshold - Italy (first stage)

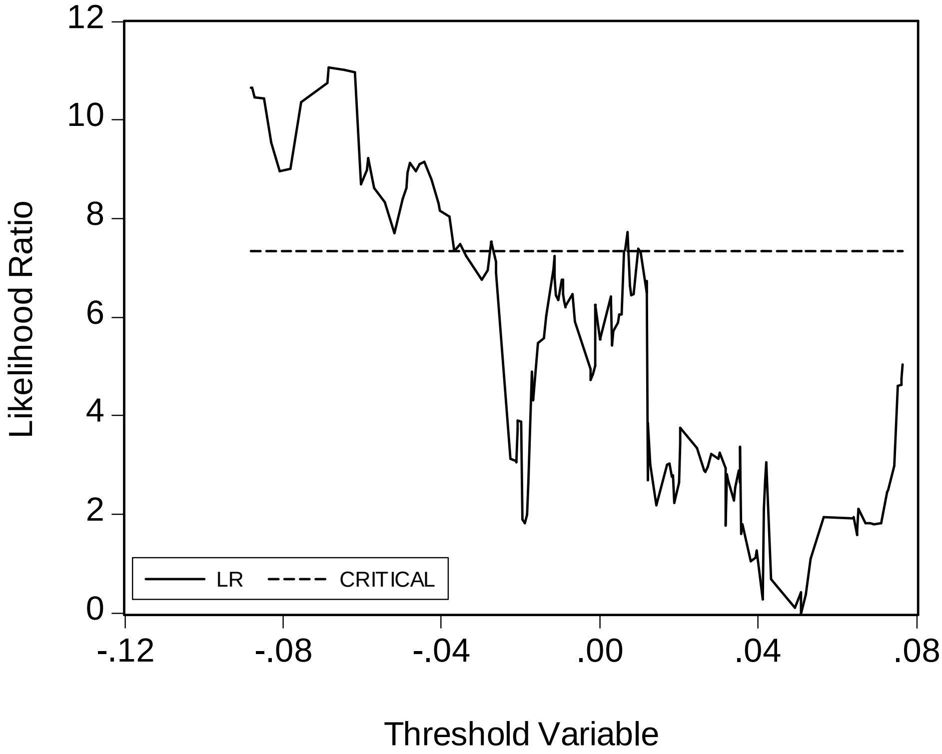
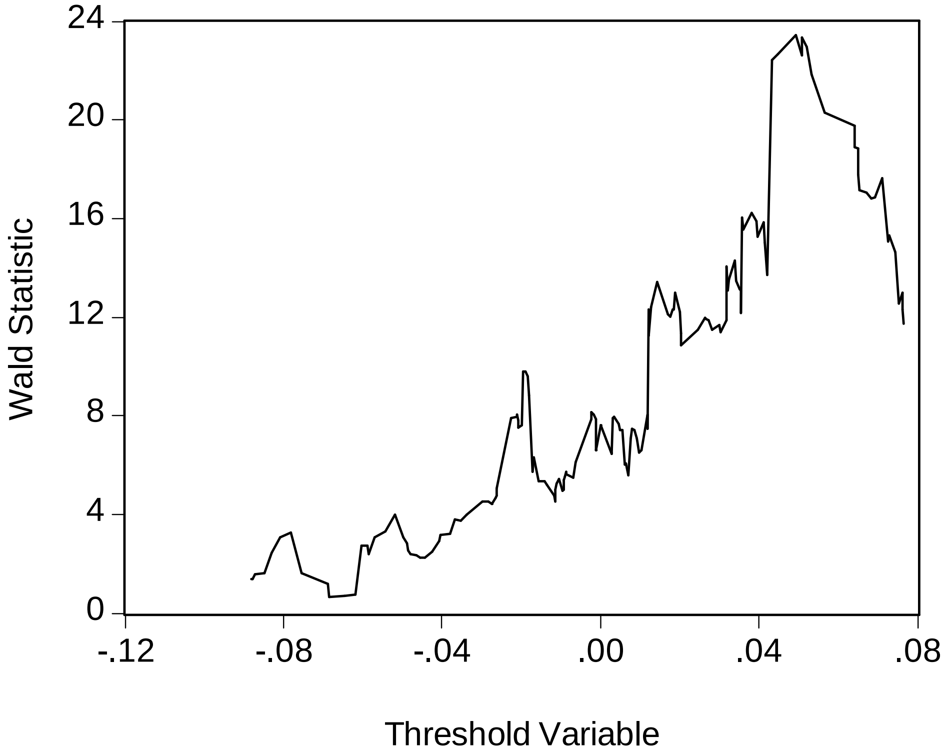


Figure 4. Heteroskedasticity-consistent Wald test – Italy (first stage)



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(lxv) This paper was presented at the EuroConference on “Auctions and Market Design: Theory, Evidence and Applications” organised by Fondazione Eni Enrico Mattei and sponsored by the EU, Milan, September 25-27, 2003

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(lxxiv) This paper was presented at the ENGIME Workshop on “Trust and social capital in multicultural cities” Athens, January 19-20, 2004

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