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Asymmetric error correction models for the oil–gasoline price relationship

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Abstract

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 behavior, namely asymmetric ECM, autoregressive threshold ECM and ECM with threshold cointegration. Each model is estimated on a common monthly data set 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 behavior. 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 prescribes that long-run price asymmetries are most likely to be found in the second stage of the transmission chain. Conversely, the ECM with threshold cointegration suggests that long-run price asymmetries vary across countries and markets. Short-run price asymmetries are captured by the asymmetric ECM specification and the TAR-ECM. The latter model suggests that all European countries are likely to be affected by asymmetries at the distribution stage, while the results obtained with the asymmetric ECM are mixed.

Keywords: Oil and gasoline prices; Asymmetries; Error correction models

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 behavior 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, the present uncertainty in supply and reserve availability, and growing world energy demand 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, crude oil prices strengthened since mid May 2005 up to the end of August 2005. During the same days, the price of gasoline has risen even more sharply, supporting the notion that input price increases are immediately transmitted to output prices. WTI crude oil daily posted price was quoted below 44US$ per barrel ($/bbl) in mid May 2005, it was above 58$/bbl in late July 2005, and it recorded quotations above 66$/bbl in late August 2005 (i.e. +14% in 1 month). IPE Brent crude oil daily closing price was between 45$/bbl and 50$/bbl in mid May 2005, and it jumped to more than 65$/bbl in late August (that is, +8% between July and August). Conversely, NYMEX unleaded gasoline daily price was below 1.50US&dollar per gallon ($/gal) in mid May 2005, it recorded 1.75$/gal in late July 2005 and soared to 2.50$/gal in late August 2005 (i.e. +43% in 30

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days). At the end of the first decade of October 2005 WTI and Brent prices were around 60$/bbl, while unleaded gasoline price was back to the end-of-July level of 1.75$/gal. The scenario depicted above rises at least one major question: when oil prices fall, are gasoline prices likely to decrease immediately, or will they respond with some delay?

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, models, tax regimes and petroleum fiscal systems. This explains the reason why there is no clear-cut evidence in the empirical literature 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 behavior, 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 data set which describes the retail and wholesale gasoline markets of France, Germany, Italy, Spain and UK over the period 1985–2003. The empirical results obtained from the different econometric specifications are compared, the predictions of each model are interpreted from a policy perspective, and each model is evaluated in terms of its ability to capture specific types of asymmetric price behavior.

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; whether the price data are gross or net of taxes.

The problem of a different response to price increases and decreases is first considered in Bacon (1991), where attention is paid to the UK 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, while 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 gasoline prices.

Again the UK 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 is used which allows for asymmetry only in the dynamic part of the equation. Weak and non-persistent asymmetry in price changes is found, which is absorbed within four months. No formal tests of asymmetric price effects are performed.

Karrenbrock (1991) employs 1983–1990 monthly data to study the empirical relationship between US 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 increases 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 during the former period only 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 US market using monthly data for the period 1982–1990. His dynamic model shows no evidence of asymmetric effect.

Again the US attracts the interest of Duffy-Deno (1996), and in particular the downstream relationship between wholesale and net-of-tax-retail gasoline prices. Data 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 US 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 to 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 behavior. 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 et al. (1997) by using two different model specifications with weekly data from 1987 to 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- and short-run asymmetries. 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 authors do not obtain unambiguous evidence concerning asymmetry, which turns out to be weak in the specification in levels, while moderate and persistent in the ECM.

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

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

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 threshold autoregressive (TAR) 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 1980–1996. 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 US gasoline markets, they also find that firms with market power adjust prices more slowly than competitive firms.

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 common explanations 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 aggressive behavior over market shares. 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.

Johnson (2002) uses an asymmetric ECM approach to analyze retail price responses to changes in wholesale prices in 15 US cities. Weekly data from the beginning of July 1996 to the end of June 1998 support evidence of asymmetric responses.

Salas (2002) employs ordered Probit, partial adjustment, and vector error correction models to characterize price adjustments in the Philippine retail gasoline market since its deregulation in 1998. It is found that pricing decisions of oil firms depend significantly on 8 weeks of previous changes in crude costs. Moreover, the speed of adjustment of retail prices to their long-run equilibrium relation with crude cost has been following an accelerating trend but it is vulnerable to intervening factors. Lastly, empirical evidence suggests that pump prices respond more quickly and fully to increases in crude costs rather than to decreases.

Bachmeier and Griffin (2003) consider daily data and adopt an Engle–Granger two-step cointegration approach. No evidence of asymmetry is found for the US 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 after 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) analyze 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 data sets, one for each working day. The conclusions on asymmetric pricing are shown to differ over these data sets, suggesting that the choice of the day for which the prices are observed is extremely relevant. 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 price asymmetry on Dutch consumers 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 for the period 1985–2000. They carry out an international comparison of gasoline markets for five European countries. A two-stage modelling of the transmission mechanism is used in order to assess possible asymmetries at either the refinery stage, the distribution stage or both, while the asymmetric ECM is employed to distinguish between short- and long-run asymmetries. 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 menu-cost interpretation of sticky prices 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 9 Philadelphia gasoline wholesalers. Nevertheless, they reject the menu-cost model as a literal description of these firms' behavior, arguing instead that price stickiness arises from strategic considerations of how customers and competitors will react to price changes.

Radchenko (2005a) analyzes the stickiness of gasoline prices. The author uses a Markov-switching model and allows for different responses of retail prices to long-term and short-term shocks in upstream prices, as well as positive and negative variations in crude oil prices. Using weekly data from March 1991 to August 2002 for the US, he shows that the response of gasoline prices to short-term changes in oil prices is significantly different from the response to long-term oil price variations.

The impact of oil price volatility on the degree of price asymmetry is studied in Radchenko (2005b). 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 US retail market over the period March 1991–February 2003.

Finally, Kaufmann and Laskowski (2005) analyze monthly data on the US 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 model specification, they find asymmetries in the home heating oil market.

Table 1 summarizes the surveyed literature according to the following criteria: type of model, country, period, data frequency, stage of the production–distribution chain, type of retail gasoline price (i.e. gross or net of taxes) and the presence or absence of asymmetries. 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 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- and long-run using dynamic econometric models which exploit the presence of cointegration between the relevant variables. The most striking result is that more than 56% of the studies concentrate on US, and asymmetries are found in more than 65% of the cases.

It is widely acknowledged that taxes (both in the level and form) tend to affect the type of response of gasoline prices to crude oil price changes. Table 1 allows the reader to verify whether in the existing literature asymmetries actually vary according to the presence or absence of taxes in the retail gasoline prices. The results say that 66.7% of the studies which support the presence of asymmetric price behavior employ net-of-tax gasoline prices, that is asymmetries emerge more easily once the fiscal veil is removed. Moreover, according to Table 1, Germany is the only European country to support during the 1980s a symmetric reaction of the gross-of-tax gasoline price to oil price changes. This finding is consistent with the behavior of governments during periods of falling rude oil prices. Specifically, in the year from 1989 to 1993, a period in

Table 1
Survey of the literature

Study	Model	Country	Time	Data frequency	Stage of the production–distribution chain	Type of retail gasoline price	Asymmetry
Bacon (1991)	QQAM	UK	1982–1989	Biweekly	Second	Net-of-tax	Yes
Manning (1991)	ECM with short-run asymmetry	UK	1973–1988	Monthly	Single	Gross-of-tax	Yes
Karrenbrock (1991)	Linear adjustment	US	1983–1990	Monthly	Second	Net-of-tax	Yes
Kirchgässner and Kübler (1992)	ECM with short-run asymmetry	Germany	1972–1989	Monthly	Second	Gross-of-tax	Yes (1970s); No (1980s)
Shin (1994)	First differences	US	1982–1990	Monthly	First; Second	Net-of-tax Gross-of-tax	No (first); inconclusive (second)
Duffy-Deno (1996)	First differences	US	1989–1993	Weekly	Second	Net-of-tax	Yes
Borenstein et al. (1997)	ECM with short-run asymmetry	US	1986–1992	Weekly; biweekly	First; Second	Net-of-tax	Yes
Balke et al. (1998)	ARDL; asymmetric ECM	US	1987–1997	Weekly	First; Second; Single	Net-of-tax; Gross-of-tax	No (ARDL); Yes (ECM)
Reilly and Witt (1998)	ECM with short-run asymmetry	UK	1982–1995	Monthly	Single	Net-of-tax	Yes
Akarca and Andrianacos (1998)	Linear regression	US	1976–1996	Monthly	Single	Gross-of-tax	—
Godby et al. (2000)	Threshold ECM	Canada	1990–1996	Weekly	Single	Gross-of-tax	No
Asplund et al. (2000)	ECM with short-run asymmetry	Sweden	1980–1996	Daily; monthly	Second	Net-of-VAT	Inconclusive
Borenstein and Shepard (2002)	Lagged adjustment; VAR; PAM	US	1985–1995	Daily	First	—	Yes
Eckert (2002)	ECM with short-run asymmetry; switching regression	Canada	1989–1994	Weekly	Second	Net-of-tax	Yes (ECM); No (Switching regression)
Johnson (2002)	ECM with short-run asymmetry	US	1996–1998	Weekly	Second	Net-of-tax	Yes
Salas (2002)	Ord. Probit; PAM; VECM	Philipp.	1999–2002	Weekly	Single	Net-of- tax	Yes
Bachmeier and Griffin (2003)	ECM with short-run asymmetry	US	1985–1998	Daily	First	—	No
Bettendorf et al. (2003)	ECM with short-run asymmetry	Netherl.	1996–2001	Weekly	Second	Gross-of-tax	Inconclusive
Galeotti et al. (2003)	Asymmetric ECM	France; Germany; Italy; Spain; UK	1985–2000	Monthly	First; Second; Single	Net-of-tax	Yes
Davis and Hamilton (2004)	Structural dyn. regr.; Asym. Logit; Autoreg. Cond. Hazard	US	1989–1991	Daily	First	—	Yes
Radchenko (2005a)	Markov-switching model	US	1991–2002	Weekly	Second; Single	Gross-of-tax	Yes
Radchenko (2005b)	VAR; PAR	US	1991–2003	Weekly	Single	Gross-of-tax	Yes
Kaufmann and Laskowski (2005)	ECM with short-run asymmetry	US	1986–2002	Monthly	Single	Net-of-tax	Yes

Notes: "first" indicates the transmission of crude oil prices to spot gasoline prices. "second" indicates the reaction of retail gasoline prices to spot gasoline prices. "single" is the direct relationship between retail gasoline prices and crude oil prices. "net-of-tax" retail gasoline price = gross-of-tax gasoline price/(1-value added tax)-excise duty. "net-of-VAT" retail gasoline price = gross-of-tax gasoline price/(1 + value added tax). "asymmetry" is defined as the tendency of downstream prices to respond differently to increases in upstream prices than to decreases. "inconclusive" means that the paper provides mixed evidence about asymmetries. "—"indicates that the corresponding issue is not considered in the study.

which crude oil prices were in decline, most OECD member nation countries seized the opportunity to increase excise taxes on petroleum products. This was most evident among the European industrialized countries.[1]

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 US dollar and individual national currencies (*ER*) for five European countries, namely France, Germany, Italy, Spain and UK.[2] 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 US$/bbl.

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 North Western 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).[3] 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 is from DATASTREAM. The exchanges rates series are obtained from the International Monetary Fund 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. A necessary condition for cointegration is the integration of the series of output and input prices. Table A1 in the appendix reports the augmented Dickey–Fuller (ADF) statistics for each series. As expected, all price series are integrated of order one, or $I(1)$.

3.1. Asymmetric ECM

If the variables are $I(1)$, 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, $j = 1, \ldots, m$, in levels:

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

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

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

with Δ indicating the first difference operator, and p the lag-length. According to Table A2 in the appendix, the Engle–Granger test suggests the presence of cointegration between gasoline price and crude oil price at each stage of the production–distribution chain.

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

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

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 $\hat{\varepsilon}_t \leqslant 0$, while $\hat{\varepsilon}^-_t$ equals $\hat{\varepsilon}_t$ or zero when $\hat{\varepsilon}_t < 0$ or $\hat{\varepsilon}_t \geqslant 0$. 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 $j = 1, \ldots, m$ and $i = 0, \ldots, p$.

Simple inspection of the sign, magnitude and statistical significance of the estimated coefficients offers a first insight on the presence of asymmetric price behavior. However, in order to establish if the estimated coefficients of model (3) are statistically different, the (single or joint) hypotheses H_0: $\alpha^+ = \alpha^-$, $\lambda_i^+ = \lambda_i^-$, $\gamma_i^+ = \gamma_i^-$, …, $\delta_i^+ = \delta_i^-$ 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

[1] Taxes per barrel of oil equivalent increased by almost 70% in UK, and 50% in Italy over the period. France was the least aggressive in raising product taxes, although this was from a relatively high pre-1989 base level. The authors would like to thank an anonymous referee for suggesting this explanation.

[2] The exchange rate between the US$ and the Euro is multiplied by the fixed parity for each country after January 1999.

[3] 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.

symmetry. A few recent studies (see Cook et al., 1998, 1999; 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 Eq. (3) adds a 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

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

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 $(\alpha, \lambda_i, \gamma_i, \ldots, \delta_i)$ if $q_t \leqslant \gamma$, and $(\alpha+\alpha^*, \lambda_i+\lambda_i^*, \gamma_i+\gamma_i^*, \ldots, \delta_i+\delta_i^*)$ if $q_t > \gamma$. Alternatively, if we define $Y_t = (\varepsilon_{t-1} \quad \ldots \quad \Delta x_{mt-p})'$, $Y_t(\gamma) = (Y_t' \quad Y'1(q_t > \gamma))'$, $\theta_1 = (\alpha, \ldots, \delta_p)'$, $\theta_2 = (\alpha^*, \ldots, \delta_p^*)'$ and $\theta_1 = (\theta_1' \quad \theta_2')'$, model (4) can be expressed as

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

Since Eq. (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 TAR model (5) relative to the linear model (2). The null hypothesis in this case is $H_0: \alpha^* = \lambda_i^* = \gamma_i^* = \cdots = \delta_i^* = 0$ for each i. Defining the selector matrix $R = (0 \quad \mathrm{I})$, $M(\gamma) = \sum Y_t(\gamma) Y_t(\gamma)'$ and $V(\gamma) = \sum Y_t(\gamma) Y_t(\gamma)' \hat{e}_t^2$, where I is the identity matrix of appropriate dimension, we can write the pointwise heteroskedasticity-consistent Wald statistic as

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

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 $W^*(\gamma) = T(\tilde{S}^* - S^*(\gamma))/S^*(\gamma)$, 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 bth iteration. The p-value for W is given by

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

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 (LR)-type statistic is

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

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

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

where

$$\xi = \max_{s\in R}(2W(s) - |s|) \quad \text{with } W(v) = \begin{cases} W_1(-v) & v < 0, \\ 0 & v = 0, \\ W_2(v) & v > 0, \end{cases}$$

$W_1(-v)$ and $W_2(v)$ 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

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

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:

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

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

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

$$I_t = \begin{cases} 1 & \text{if } \hat{\varepsilon}_{t-1} > 0, \\ 0 & \text{if } \hat{\varepsilon}_{t-1} \leqslant 0 \end{cases} \tag{9}$$

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

$$I_t = \begin{cases} 1 & \text{if } \Delta\hat{\varepsilon}_{t-1} > 0, \\ 0 & \text{if } \Delta\hat{\varepsilon}_{t-1} \leqslant 0. \end{cases} \tag{10}$$

Eqs. (8) and (9) are referred to as TAR cointegration, while models (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 Eqs. (8) and (9). If the errors are serially correlated, Eq. (8) can be augmented with the lagged differences of $\hat{\varepsilon}_t$ as in the standard ADF test:

$$\Delta\hat{\varepsilon}_t = I_t\rho_1\hat{\varepsilon}_{t-1} + (1 - I_t)\rho_2\hat{\varepsilon}_{t-1} + \sum_{i=1}^{p-1}\sigma_i\Delta\hat{\varepsilon}_{t-i} + v_t. \tag{11}$$

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, Eq. (10) becomes

$$I_t = \begin{cases} 1 & \text{if } \Delta\hat{\varepsilon}_{t-1} > \hat{\mu}, \\ 0 & \text{if } \Delta\hat{\varepsilon}_{t-1} \leqslant \hat{\mu}, \end{cases} \tag{12}$$

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

Once Eq. (11) is estimated, the null hypothesis $H_0 : \rho_1 = \rho_2 = 0$ of no cointegration can be tested through an F test. Correct critical values depend on the number of observations, the number of lags in Eq. (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 an 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

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

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

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. Following the majority of the empirical literature, our study uses the net-of-taxes gasoline prices. Given the different fiscal systems which characterized the countries under scrutiny, this choice will ease the comparison of the empirical findings between countries.

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 2–6 refer to the asymmetric ECM. The estimated coefficients and corresponding t-statistics are reported in Tables 2–4, whereas Tables 5–6 present the results of testing for price asymmetries. Coefficients α^+ and α^- in Table 2 indicate asymmetric adjustment speeds, which measure long-run asymmetry, while the coefficients γ_i^+ and γ_i^-, $i = 1, \ldots, p$, 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- and short-run, as well as

Table 2
Asymmetric ECM—asymmetric adjustment speeds and short-run price asymmetries

	France	Germany	Italy	Spain	UK
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 Eq. (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 ith lag ($i = 1, 2, 3$) indicates that the optimal number of lags is $i-1$.

Table 3
Asymmetric ECM—exchange rate asymmetries

	France	Germany	Italy	Spain	UK
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 Eq. (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. "—" in correspondence to the ith lag ($i = 1, 2, 3$) indicates that the optimal number of lags is $i-1$.

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

Table 4
Asymmetric ECM—autoregressive asymmetries

	France	Germany	Italy	Spain	UK
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 Eq. (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 5
Asymmetric ECM—computed *F* tests for asymmetric adjustment speeds and short-run effects

Null hypothesis	France	Germany	Italy	Spain	UK
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 Eq. (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. "—" in correspondence to the *i*th lag ($i = 1, 2, 3$) indicates that the optimal number of lags is $i-1$.

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 UK. 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. This finding suggests that competition among retailers is higher in Germany than in Italy and Spain.

Table 6
Asymmetric ECM—simulated F tests for asymmetric adjustment speeds and short-run effects

Null hypothesis	France	Germany	Italy	Spain	UK
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 1000 replications) of the null hypothesis of symmetry using an F test at 5% significance level. "—" in correspondence to the ith lag ($i = 1, 2, 3$) indicates that the optimal number of lags is $i-1$.

Table 3 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 to model the production stage separately from the distribution stage.

The estimated autoregressive coefficients, which enter the model when the lag-length is equal to, or larger than, one, are reported in Table 4. 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 UK. 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 5 reports the calculated conventional F test for the hypothesis of long- and short-run asymmetries. Rejection of the null hypothesis $\alpha^+ = \alpha^-$ implies asymmetric long-run adjustment of the output price to the equilibrium level of the input price, whereas short-run asymmetries arise when at least one of the null hypotheses $\gamma_i^+ = \gamma_i^-$ (crude price symmetry), $\delta_i^+ = \delta_i^-$ (exchange rate symmetry) or $\lambda_i^+ = \lambda_i^-$(autoregressive symmetry), is rejected, with $i = 0, 1$ indicating the number of lags.[5] Table 5 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 the retail-crude relationship (single stage) for France and Italy, as well as the retail-spot relation (second stage) for Germany. The lagged price effects are asymmetric in the spot-crude relation (first stage) for France and Germany, and at the second stage in France, Spain and UK. Moreover, the reaction to exchange rate variations is asymmetric in UK at the first stage. Finally, contemporaneous price asymmetries arise in France and UK 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 5 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 6 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

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

larger than 15% and 58% ("high" rejection frequencies): these amount to 32 and 8 out of 64.In contrast with 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 summarize, 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 behavior 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 7–9 report the estimated value and significance of the coefficients of the TAR-ECM specification. Table 10 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. Figs. 1–4 plot the adjusted LR 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 Eq. (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 $(\lambda_i, \gamma_i, \ldots, \delta_i)$ and $(\lambda_i + \lambda_i^*, \gamma_i + \gamma_i^*, \ldots, \delta_i + \delta_i^*)$. 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 7–9, the coefficients accounting for both long- and short-run price asymmetries which are statistically significant at 5% are 24 out of 71. If we concentrate in Table 7, significant differences in long-run coefficients (i.e. α^*) arise in 4 cases out of 15, whereas heterogeneous short-run effects (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

Table 7
TAR-ECM—two-regime adjustment speeds and short-run price effects

	France	Germany	Italy	Spain	UK
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 α^*	−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 α^*	−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 α^*	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 Eq. (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. "—" in correspondence to the ith lag ($i = 1, 2, 3$) indicates that the optimal number of lags is i–1.

Table 8
TAR-ECM—two-regime exchange rate effects

	France	Germany	Italy	Spain	UK
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 Eq. (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. "—" in correspondence to the ith lag ($i = 1, 2, 3$) indicates that the optimal number of lags is i–1.

Table 9
TAR-ECM—two-regime autoregressive effects

	France	Germany	Italy	Spain	UK
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 Eq. (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. "—" in correspondence to the ith lag ($i = 1, 2, 3$) indicates that the optimal number of lags is i–1.

Table 10
TAR-ECM—estimated thresholds and computed Wald tests

	France	Germany	Italy	Spain	UK
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: "*" 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 (1000 replications).

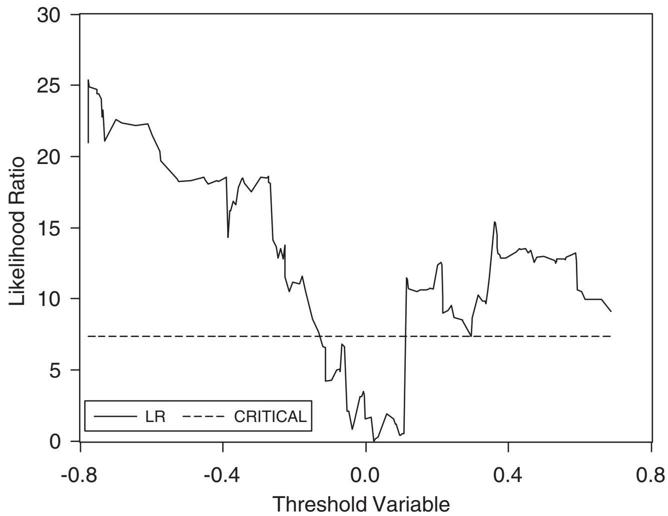

Fig. 1. Likelihood ratio test for the threshold—France (single stage).

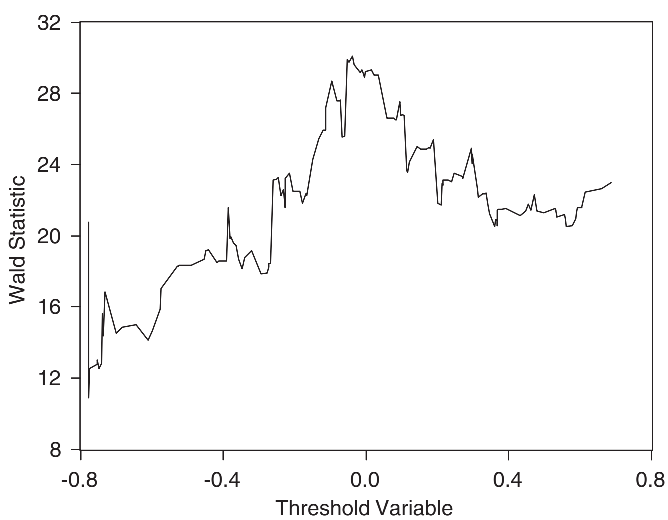

Fig. 2. Heteroskedasticity-consistent Wald test—France (single stage).

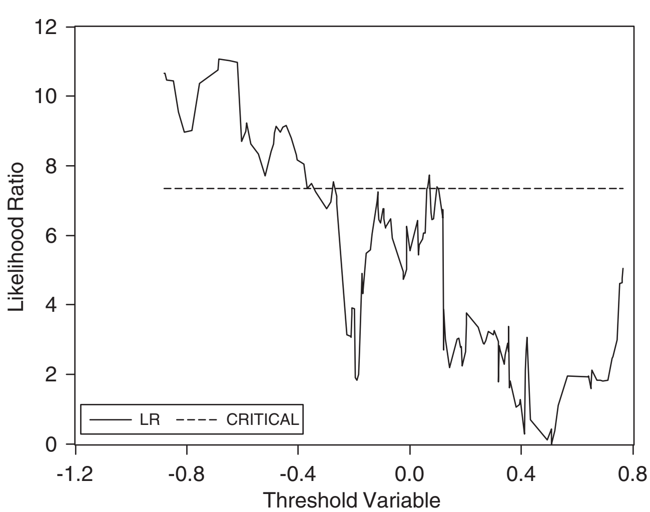

Fig. 3. Likelihood ratio test for the threshold—Italy (first stage).

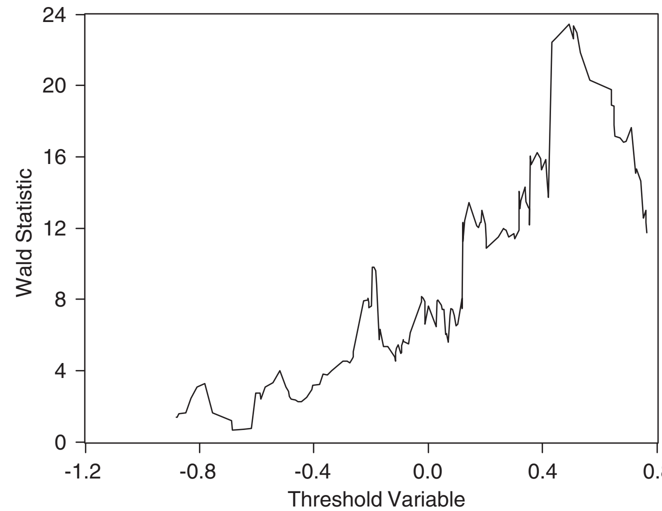

Fig. 4. Heteroskedasticity-consistent Wald test—Italy (first stage).

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 8 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 9 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 an LR approach, after adjusting the LR statistic for heteroskedasticity in the residuals.[6] As an illustration, Figs. 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 an 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 Fig. 1 is rather tight, while the threshold estimates seem to be less precise

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

in Fig. 3. As far as the other countries are concerned, LR plots are well-shaped (i.e. similar to Fig. 1) in about 50% of the cases. The estimates of the threshold are reported in Table 10. Significant and positive threshold values are found in 4 countries, namely France and Germany in the first stage, Italy in the second stage and UK in the single stage.

In order to test the null hypothesis of linearity against the threshold model we use a heteroskedasticity-consistent Wald statistic. Figs. 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 10. 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 behavior 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 behavior 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 11–16 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 11–13. The estimated coefficients of the asymmetric ECM with threshold cointegration are reported in Tables 14–16.

Tables 11–13 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 favor 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 14. The most

Table 11
TAR, M-TAR and MC-TAR cointegrating relations—first stage

	France			Germany			Italy			Spain			UK		
	TAR	M-TAR	MC-TAR	TAR	M-TAR	MC-TAR	TAR	M-TAR	MC-TAR	TAR	M-TAR	MC-TAR	TAR	M-TAR	MC-TAR
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 12
TAR, M-TAR and MC-TAR cointegrating relations—Second stage

	France			Germany			Italy			Spain			UK		
	TAR	M-TAR	MC-TAR	TAR	M-TAR	MC-TAR	TAR	M-TAR	MC-TAR	TAR	M-TAR	MC-TAR	TAR	M-TAR	MC-TAR
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 13
TAR, M-TAR and MC-TAR cointegrating relations—single stage

	France			Germany			Italy			Spain			UK		
	TAR	M-TAR	MC-TAR	TAR	M-TAR	MC-TAR	TAR	M-TAR	MC-TAR	TAR	M-TAR	MC-TAR	TAR	M-TAR	MC-TAR
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 11–13: 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 14
ECM with threshold cointegration—asymmetric adjustment speeds and short-run price effects

	France	Germany	Italy	Spain	UK
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 Eq. (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. "—" in correspondence to the ith lag ($i = 1, 2, 3$) indicates that the optimal number of lags is $i-1$.

Table 15
ECM with threshold cointegration—exchange rate effects

	France	Germany	Italy	Spain	UK
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 Eq. (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. "—" in correspondence to the ith lag ($i = 1, 2, 3$) indicates that the optimal number of lags is $i-1$.

Table 16
ECM with threshold cointegration—autoregressive effects

	France	Germany	Italy	Spain	UK
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 Eq. (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. "—" in correspondence to the ith lag ($i = 1, 2, 3$) indicates that the optimal number of lags is $i-1$.

relevant asymmetric effects appear in the single stage. In this case, 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 UK. As for the first stage, all coefficients are significant and, in the case of Italy and Spain, the estimated

Table 17
Summary of the empirical results

Stage	Asymmetric ECM			TAR-ECM			ECM with threshold cointegration		
	First	Second	Single	First	Second	Single	First	Second	Single
Panel (a) France									
LR asymmetries (price effect) (%)	13.3%	16.5%	41.2%	—	—	—	0%	100%	100%
SR asymmetries (price effect) (%)	40.05%	38.45%	49.65%	17%	80%	50%	—	—	—
SR asymmetries (exchange rate effect) (%)	27.3%	—	8.55%	—	—	—	—	—	—
Panel (b) Germany									
LR asymmetries (price effect)	11.7%	66.9%	10.7%	—	—	—	0%	0%	0%
SR asymmetries (price effect)	37%	6.5%	6.1%	33%	50%	33%	—	—	—
SR asymmetries (exchange rate effect)	17%	—	28%	—	—	—	—	—	—
Panel (c) Italy									
LR asymmetries (price effect)	9.4%	46.1%	73.4%	—	—	—	100%	0%	100%
SR asymmetries (price effect)	27.25%	10%	5%	0%	33%	33%	—	—	—
SR asymmetries (exchange rate effect)	40%	—	36.8%	—	—	—	—	—	—
Panel (d) Spain									
LR asymmetries (price effect)	6.5%	48%	8.1%	—	—	—	100%	100%	100%
SR asymmetries (price effect)	27.5%	44.85%	11.55%	33%	40%	33%	—	—	—
SR asymmetries (exchange rate effect)	34%	—	7.5%	—	—	—	—	—	—
Panel (e) UK									
LR asymmetries (price effect)	5.4%	29.9%	4.5%	—	—	—	0%	0%	100%
SR asymmetries (price effect)	9%	40.65%	30.6%	17%	25%	50%	—	—	—
SR asymmetries (exchange rate effect)	86.4%	—	28.25%	—	—	—	—	—	—

Notes: for the asymmetric ECM the calculated percentages are based on Table 6. For the TAR-ECM the calculated percentages are based on Tables 7–9. For the ECM with threshold cointegration the percentages correspond to Table 14. 100% (0%) = difference between adjustment speeds coefficients is always larger than (smaller than) 0.05 (statistically insignificant coefficients are not considered).

adjustments from below to the equilibrium exceed the corresponding adjustments from above by more than 0.05. The differences between the estimated coefficients are smaller in the second stage. It is worth pointing out that, contrary to the asymmetric ECM, the ECM with threshold cointegration identifies long-run asymmetries of the expected sign (i.e. adjustments from below are found to be faster than adjustments from above).[7] This result may suggest 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 15 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 behavior 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 16. In line with the results obtained by 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.

In Table 17, a comprehensive summary of the empirical results is presented. Results are summarized for each country separately, and classified according to the type of econometric model (asymmetric ECM, TAR-ECM, ECM with threshold cointegration), the type of asymmetry (long-run price effect, short-run price effect, short-run exchange rate effect) and stage of the price transmission chain (spot gasoline price-crude price, or first stage; retail gasoline price-spot gasoline price, or second stage; retail gasoline price-crude price, or single stage). The table entries represent the frequencies with which each simulated test

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

statistic rejects the null hypothesis of symmetry. Missing values imply that a particular model is not designed to capture a particular type of asymmetry. Table 17 is extremely useful, since it helps the reader to answer at least two interrelated questions: (i) which model is more suitable to account for asymmetric behavior in the oil price-gasoline price relationship? (ii) could the variability of results across the different stages of the production–distribution chain be interpreted in the light of the different degree of competition which characterizes each country?

The asymmetric ECM prescribes that long-run price asymmetries are most likely to be found in the second stage of the trasmission chain for Spain, UK and Germany, while asymmetries mostly affect the whole retail gasoline price-crude price relation for Italy and France. Conversely, the ECM with threshold cointegration suggests that long-run price asymmetries affect Spain in each stage of the price transmission mechanism; Italy and France show asymmetric price behavior in two of three stages (first and single for Italy; second and single for France), UK in the single stage only and Germany in none of them. Overall, the asymmetric ECM is more likely to conclude in favor of long-run asymmetries than the ECM with threshold cointegration. The policy implications which can be drawn from these two models are also quite different. According to the former, long-run asymmetries in European countries occur at the distribution level, a result which suggests a low degree of competion among retailers. For the latter, the picture is more articulated: at one extreme, Spain, Italy and France seem to require policy interventions to favor a higher degree of competition among producers and retailer; at the other extreme, the competitive structure of gasoline market in Germany is able to control potential asymmetries in the input–output price transmission mechanism.

Short-run price asymmetries are captured by the asymmetric ECM specification and the TAR-ECM. The latter model suggests that all European countries are likely to be affected by asymmetries at the distribution stage, with the exception of UK: in the short-run, gasoline prices rise faster than they fall, since the lack of competition among retailers increases the persistence of high gasoline prices. On the contrary, the results obtained with the asymmetric ECM are mixed. Short-run asymmetries are present at the distribution level only in Spain and UK, they affect the production stage in Italy and UK, while for France it is not possible to discriminate which part of the price transmission mechanism is responsible of asymmetric price reactions. To sum up, the TAR-ECM is more likely to support short-run asymmetries at the distribution level than the asymmetric ECM. These findings seem to indicate that, while the simpler asymmetric ECM is indicated for capturing long-run (i.e. cointegration-related) asymmetries, the presence of short-run price reactions should be investigated with models designed to account for non-linear short-run price dynamics, such as the TAR-ECM specification.

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 data set, 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 behavior, 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 data set we use in the empirical application includes crude oil, spot and retail gasoline prices, together with exchange rates for France, Germany, Italy, Spain and UK 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 behavior. 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 prescribes that long-run price asymmetries are most likely to be found in the second stage of the transmission chain for Spain, UK and Germany, while asymmetries mostly affect the whole retail gasoline price-crude price relation for Italy and France. Conversely, the ECM with threshold cointegration suggests that long-run price asymmetries affect Spain in each stage of the price transmission mechanism; Italy and France show asymmetric price behavior in two of three stages (first and single for Italy; second and single for France), UK in the single stage only and Germany in none of them. Overall, the asymmetric ECM is more likely to conclude in favor of long-run asymmetries than the ECM with threshold cointegration. The policy implications which can be drawn from these two models are also quite different. For instance, according to the asymmetric ECM, long-run asymmetries in European countries occur at the distribution level, a result which suggests a low degree of competion among retailers.

Short-run price asymmetries are captured by the asymmetric ECM specification and the TAR-ECM. The latter model suggests that all European countries are likely to be affected by asymmetries at the distribution stage, with the exception of UK: in the short-run, gasoline prices rise faster than they fall, since the lack of competition among retailers increases the persistence of high gasoline prices. On the contrary, the results obtained with the asymmetric ECM are mixed. Short-run asymmetries are present at the distribution level only in Spain and UK, they affect the production stage in Italy and UK, while for France it is not possible to discriminate which part of the price transmission mechanism is responsible of asymmetric price reactions.

To sum up, the TAR-ECM is more likely to support short-run asymmetries at the distribution level than the asymmetric ECM. These findings seem to indicate that, while the simpler asymmetric ECM is indicated for capturing long-run (i.e. cointegration-based) asymmetries, the presence of short-run price reactions should be investigated with models designed to account for non-linear short-run price dynamics, such as the TAR-ECM specification.

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Appendix A

Data of ADF unit root tests and cointegration tests are found in Tables A1 and A2, respectively.

Table A1
ADF unit root tests

	France	Germany	Italy	Spain	UK
Series CR_t					
$a \neq 0$	Yes	Yes	Yes	Yes	Yes
$b \neq 0$	Yes	No	No	Yes	Yes
p	2	2	4	2	3
ADF	−3.350	−3.228*	−3.168*	−3.672*	−3.996*
Series ΔCR_t					
ADF	−9.608**	−10.130**	−9.497**	−9.605**	−10.411**
Series SP_t					
$a \neq 0$	Yes	Yes	Yes	Yes	Yes
$b \neq 0$	No	No	Yes	No	Yes
p	2	2	2	2	2
ADF	−2.771	−2.824	−3.418	−2.288	−3.107
Series ΔSP_t					
ADF	−11.084**	−10.936**	−11.044**	−11.083**	−11.762**
Series NR_t:					
$a \neq 0$	Yes	Yes	Yes	Yes	Yes
$b \neq 0$	No	No	Yes	Yes	Yes
p	1	0	1	1	3
ADF	−2.346	−2.849	−3.678*	−3.323	−3.173
Series ΔNR_t					
ADF	−13.241**	−14.396**	−10.987**	−12.662**	−10.482**
Series ER_t					
$a \neq 0$	Yes	Yes	Yes	Yes	Yes
$b \neq 0$	Yes	Yes	No	No	Yes
p	3	3	1	1	3
ADF	−3.203	−3.280	−1.622	−2.808	−3.787*
Series ΔER_t					
ADF	−11.060**	−10.909**	−10.359**	−11.049**	−11.253**

Notes: *ADF* is the calculated t test for the null hypothesis of a unit root (i.e. $\gamma = 0$) in the series x_t from the augmented Dickey–Fuller regression: $\Delta x_t = a + bt + \gamma x_{t-1} + \sum_{i=1}^{p} \gamma_i \Delta x_{t-i} + \eta_t$. p is the order of the augmentation needed to eliminate any autocorrelation in the residuals of the *ADF* regression. * (**) indicate significance at 5% (1%) on the basis of the critical values by MacKinnon (1991).

Table A2
Cointegration tests

	France	Germany	Italy	Spain	UK
First stage					
a_0	0.527	0.670	1.443	1.202	0.617
a_1	0.809	0.816	0.839	0.819	0.844
a_2	1.132	1.124	0.899	0.913	0.901
a_3	—	—	—	—	—
p	1	1	1	1	1
ADF	−6.861***	−6.941***	−6.456***	−6.148***	−6.088***
Second stage					
a_0	−3.054	−2.877	0.586	−0.722	−3.294
a_1	0.682	0.554	0.494	0.510	0.524
a_2	—	—	0.003	0.002	—
p	2	2	2	2	1
ADF	−3.871**	−3.981***	−3.855*	−4.196**	−5.528***
Single stage					
a_0	−2.889	−2.689	−2.843	−0.695	−3.257
a_1	0.534	0.534	0.431	0.428	0.504
a_2	0.858	0.526	1.038	0.586	0.250
a_3	0.001	—	—	0.001	—
p	0	0	2	0	0
ADF	−6.949***	−6.840***	−3.386	−5.745***	−4.503***
Sample	1:1985–3:2003	1:1985–3:2003	1:1985–3:2003	1:1985–3:2003	
1:1985–3:2003					
T	219	219	219	219	219

Notes: first stage: $SP_t = a_0 + a_1 CR_t + a_2 ER_t + a_3 t + \varepsilon_t$. Second stage: $NR_t = a_0 + a_1 SP_t + a_2 t + \varepsilon_t$. Single stage: $NR_t = a_0 + a_1 CR_t + a_2 ER_t + a_3 t + \varepsilon_t$. *ADF* is the calculated t test for the null hypothesis of no cointegration (i.e. $\gamma = 0$) in the augmented Dickey–Fuller regression on $\hat{\varepsilon}_t$: $\Delta\hat{\varepsilon}_t = \gamma\hat{\varepsilon}_{t-1} + \sum_{i=1}^{p}\gamma_i \Delta\hat{\varepsilon}_{t-i} + \eta_t$. p is the order of the augmentation needed to eliminate any autocorrelation in the residuals of the *ADF* regression. * (**) [***] indicate significance at 10% (5%) [1%] on the basis of the critical values by MacKinnon (1991). T is the total number of observations. The t statistics on $\hat{a}_j$, $j = 0, \ldots, 3$, are invalid due to the unit root in the series and are not presented.

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