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Estimating Asymmetric Fuel Price Responses in Croatia

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Abstract

Background: According to many studies, the transmission of oil prices to retail fuel prices is asymmetric. Fuel prices react faster if oil prices rise and more slowly if oil prices fall. Different standard econometric procedures lead to different results. The Linex approach, which is based on formulating the non-linear adjustment cost function, reflects the theory. It uses the generalised method of moments to estimate the reaction functions, which demands many observations. **Objectives:** The paper investigates the price asymmetry in the Croatian retail fuel market using standard approaches and the Linex approach. **Methods/Approach:** The simple and dynamic asymmetry models, error correction models, threshold autoregressive co-integration, and the Linex approach are used to verify the hypothesis of asymmetric reactions of gasoline and diesel prices in Croatia. **Results** The results using the standard methods are mixed, while the Linex approach indicates price asymmetry, the size of which is measured with the average price bias. The results correspond to other studies worldwide. **Conclusions** The authors' preferred Linex approach detects price asymmetries, even with large data samples with frequent changes in trends and volatilities. According to the approach, the question is not whether prices are formed asymmetrically but the size of the asymmetry.

Keywords: retail prices; error correction model; threshold autoregressive cointegration; linear-exponential adjustment cost function; generalised method of moments

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Introduction

Fuel price shifts, especially their growth, are perceived very sensitively by the population in every country because they represent a critical factor in determining the transport price. On the other hand, transport services enter into the pricing of most other goods and services as a nonnegligible component. Therefore, every gasoline and diesel price increase is perceived very negatively, and a decrease in fuel prices, on the other hand, is usually overlooked. This psychological aspect is often reflected in the opinion that the reaction to the adjustment of gasoline prices is not the same in the case of a drop in oil prices on the stock exchange (or world markets) as in the case of an increase. This effect was named by Bacon (1991) as the "rockets and feathers phenomenon". What causes this perception of reality?

In addition to the population's feelings, four theoretical explanations for asymmetric fuel price-making are known. Borenstein et al. (1997) suggested the first three (a short review is also provided by Brown et al., 2000 and Radchenko, 2005), and Douglas et al. (2010) assume the last.

The first theoretical rationale is called oligopolistic coordination theory. Firms try to assure competitors they keep a tacit agreement by reacting asymmetrically to oil price changes.

The second explanation results from the cost of production and inventory adjustment. Firms spread the adjustment over time because adjusting production and inventory levels is costly (Borenstein et al., 2002).

The search theory is the third theoretical reason why prices are changing asymmetrically. The high oil price volatility allows the firm to take advantage of consumers' high search costs and temporarily increase its margin after the increase in the oil price.

The fourth explanation is the theory of strategic interactions (Okun, 1981). Price makers (including "selfish") try not to pit against themselves a minority of consumers who tend to "punish" firms for unjustified or insufficiently explained price increases (Rotemberg, 2011, p. 953). There is a tendency to use the crude oil price increase to adjust retail fuel prices by additional unexpected costs or the effects of an increase in demand. All the theories should be considered by the chosen methodology investigating price asymmetries.

The primary objective of this paper is to delve into the phenomenon of price asymmetry in the retail fuel market in Croatia. The study employs standard approaches but also introduces a unique method using the Linex adjustment cost function, which the authors have modified explicitly for this study (Szomolanyi et al. 2020, 2022a, 2022b). This innovative approach allows for estimating the average fuel price bias that directly results from asymmetric price making, a topic of significant interest to academic researchers and economists.

Frequent changes in trends and volatilities characterise energy prices. Therefore, many authors tend to split the data sample into more subsamples (Bagnai et al., 2018; Bumpass et al., 2019; Cipcic, 2021). This decision may come at the cost of too much loss of degrees of freedom. The methodological part will demonstrate that the generalised method of moments (GMM) is appropriate for estimating the reaction function of fuel prices derived by the Linex function. The asymptotic properties of GMMs demand larger datasets (a few hundred or more). The methodological part also discusses that the aforementioned theoretical justifications for asymmetric price transition can be mathematically formulated as a non-linear adjustment cost function. Therefore, the Linex approach corresponds to the theoretical starting points.

The research hypothesis is whether fuel prices react asymmetrically, i.e., faster on average to oil price increases than decreases. An alternative option is symmetrical price adjustment; the absolute values of the reactions are, on average, the same. The average price of fuels is higher for asymmetric reactions than for symmetric ones. The average bias is the difference between the average price for asymmetric responses and the average price for symmetrical reactions.

The central research question of this paper is to explore whether a systematic price bias exists in the Croatian retail fuel market. The entire sample of Croatian retail gasoline price data and an asymptotically consistent estimate of the econometric specification of the price reaction function derived from the theoretical basis of price asymmetries will be used. Notably, this approach, novel in the context of retail fuel pricing in Croatia, will be compared with estimates based on econometric methods commonly used in global studies.

The result is a comprehensive, complex analysis of the relationship between Croatia's gasoline and diesel retail prices and the crude oil prices in world markets. The procedures mentioned can also be used to analyse price asymmetry, for example, in agriculture, energy, or other markets. As Deltas et al. (2020) point out, the study's standard research approaches also lead to mixed results. According to the Linex adjustment cost function approach, the retail fuel prices adjust asymmetrically. The corresponding price bias is slightly higher for gasoline.

Retail fuel price decreases are more sluggish than price increases in Croatia. Such price-making systematically produces a bias, which is, on average, positive in Croatia. The average gasoline price bias is higher than diesel in Croatia. This result is in contrast with the European average. The average European retail fuel price biases are higher than those of Croatia.

Literature Review

According to many studies, starting with Bacon (1991), retail fuel prices respond to changes in crude oil prices asymmetrically. However, Perdiguero-Garcia (2013) provided a meta-analysis of price asymmetries in the gasoline market, finding that the significant variation in the outcomes reported makes drawing definitive conclusions difficult. Deltas et al. (2020) observed that using different methods, data samples, and frequencies can yield varying results when examining the pass-through of crude oil prices to retail fuel prices. Some studies split the sample into two or more sets to find evidence for the rockets' and feathers' effects (Bagnai et al., 2018; Bumpass et al., 2019; Cipcic, 2021)

Historically, studies have employed various econometric methods to detect asymmetric responses of output prices to input prices. For instance, Bacon (1991) tested the price asymmetry hypothesis using a quadratic adjustment function. Other approaches used in empirical studies of retail fuel price asymmetries include the distributed lag model (Karrenbrock, 1991), the first differences model (Duffy-Deno, 1996), cointegration techniques (Borenstein et al., 1997), and the vector autoregressive model (Balke et al., 1998; Kang et al., 2019). More recently, authors of empirical studies have utilised threshold models with multiple regimes (Douglas et al., 2010; Bagnai et al., 2018; Torrado et al., 2020; Gosinska et al., 2020), broadening this field's methodological landscape.

The most recent papers continue to show mixed results. Torrado et al. (2020) confirmed retail gasoline price asymmetries in Spain and Germany but not in France from 2011 to 2017. Gosinska et al. (2020) found asymmetric retail fuel pricing in Poland from 2000 to 2016. Bragoudakis et al. (2021) did not confirm asymmetric adjustments of gasoline prices to changes in oil prices in Greece after 2010. Using Greek data immediately after an unannounced and non-negligible increase in consumption taxes in 2010, Genakos et al. (2022) confirmed that the asymmetric gasoline price reaction is higher with lower competitiveness. Asane-Otoo et al. (2022) used daily German data from 2014 to 2018 and pooled-panel asymmetric error correction models to confirm the asymmetries in most cases. The authors note that temporal aggregation of station-level price data leads to inaccurate inferences and could account for the inconclusive findings in the literature.

In the context of the study, it is crucial to note the findings of Cipcic (2021), who investigated the asymmetric reactions of gasoline and diesel prices in several post-communist countries, including Croatia. Her results, spanning from January 2005 to June 2013, did not confirm asymmetry in the countries during the entire period under review. However, her findings did reveal asymmetric fuel price reactions in some countries from January 2009 to June 2013, underscoring the importance of temporal considerations in such analyses.

Methodology

During its evolution, the methodology of analysing asymmetric reactions of business prices has undergone extensive development, accompanied by many exciting ideas or modifications of existing methods, many of which are still used today. Due to its importance in estimating the development of prices, it is a constantly dynamically developing area within economic analyses. It is not technically possible to present all the methods used in one article, so this section categorises them into several primary groups, in which it presents the workhorses of each approach.

The primary group corresponding to the initial period of asymmetry research are Simple Models of Asymmetry, which use dummy binary variables to capture price increases and decreases. These are applied to products with critical factors determining the examined prices. There may be more factors, and the asymmetry effects may persist for extended periods, which is why these models have been dynamised.

The second group consists of Error Correction Models. They were used to reflect the impact of the non-stationarity of data-generating processes on the analysis, particularly the procedures for dealing with them. The possibility to model, in addition to short-term first differences, also the original levels of non-stationary variables meant a significant shift in methodology. This type of model allowed distinguishing between short-term and long-term asymmetric responses. Through gradual development, non-linear versions were also proposed, considering differences other than ordinary ones and solving the non-stationarity of processes.

The third type of model is the Threshold Autoregressive Cointegration Model. Using them, analysts responded to the criticism that, in the case of asymmetry, the classic cointegration test used in the previous type of models is inappropriate because it can lead to incorrect conclusions. Models from this group differ in how they search for the threshold value. For all three mentioned groups, analyses in the form of vector models were proposed, which examine the connection of the investigated commodities in several related markets.

The last of the methods (Linex) presented is an approach modified by the authors using the adjustment costs function. The principle of this approach is based on the idea that changes in economic processes reflecting shifts in input prices are not costless. A non-linear functional form can express asymmetric adjustment costs. In the beginning, four theoretical explanations of asymmetric price adjustment were presented. The adjustment costs function can formulate all four. The authors consider the approach advantageous because it reflects all the known theories of asymmetric price adjustment.

Using the Linex approach and the Netherlands and U.K. data, Pfann et al. (1993) reported that the costs of firing production workers are lower than the hiring costs. However, the opposite is true for the non-production workers (Adda et al., 2003, pp. 243-4). Surico (2007a, 2007b, 2008) used the approach to analyse the U.S. and EMU monetary policy asymmetries.

Simple Models of Asymmetry

First, a simple asymmetry model intuitively uses dummy variables of price increases and decreases. The product of price and each of these dummy variables leads to an estimate of a pair of asymmetry parameters, whose equality is tested by the F test. This procedure has been used since Tweeten et al. (1969).

The basic simple model of price asymmetry for fuel prices has the form:

$$y_t = \beta_0 + \gamma_0^+ x_t^+ + \gamma_0^- x_t^- + u_t \tag{1}$$

where y_t is regressand and the average weekly retail price of gasoline or diesel in time *t*; x_t^+ denotes the key regressor – the average weekly crude oil price in time *t* equals x_t if its value has increased over the last period and zero otherwise and x_t^- is the average weekly price of oil in time *t* equals x_t if its value has decreased over the last period and zero otherwise. The coefficients γ_0^+ and γ_0^- are the very parameters that, if the hypothesis of their equality is rejected, it means pricing asymmetry. Conversely, if the linear hypothesis of the equality of these parameters cannot be rejected, prices change symmetrically.

In addition to the primary determinant of the investigated fuel price - the oil price in this case, other essential factors influencing the price of fuels can be assumed. Several analyses confirmed the interconnectedness of gasoline and diesel prices in the fuel market, gradually leading to the development of multi-equation models.

The second model, which extends the basic model by an additional explanatory variable, is a model in the form:

$$y_t = \beta_0 + \gamma_0^+ x_t^+ + \gamma_0^- x_t^- + \delta_0 z_t + u_t \tag{2}$$

where z_t is another regressor (the price of another fuel – the gasoline price in the equation of diesel price and vice versa in time *t*, and the statistical significance of the coefficient δ_0 confirms the validity of the influence of another factor. Asymmetry in models (1) and (2) is present if the null hypothesis of $\gamma_0^+ = \gamma_0^-$ is rejected. The F test can test this linear hypothesis in the linear model.

The last presented models in this group are dynamic models, similar to Karrenbrock (1991), in which instead of the price level, its change is examined:

$$\Delta y_t = \beta_0 + \sum_{i=0}^{s} \gamma_i^+ \Delta x_{t-i}^+ + \sum_{i=0}^{q} \gamma_i^- \Delta x_{t-i}^- + u_t \tag{3}$$

The duration of the rise response may not be the same as the fall, which will be reflected in the difference between the *s* and *q* values. The cumulative effect of price variation can be tested with the hypothesis: $\sum_{i=0}^{s} \gamma_i^+ = \sum_{i=0}^{q} \gamma_i^-$. The F test can again be used to check for asymmetry.

Error Correction Models

The development of most commodities' prices tends to be non-stationary because their essential characteristics (distribution moments) change over time. For example, the price of the studied commodity grows over time because inflationary pressures act on it so that the average will grow over time. However, the non-stationarity of the processes generating the investigated variables can lead to spurious regressions and conclusions that identify factors that do not influence them as determinants of prices. The solution to this problem was the second model group, the Error Correction Model.

The basis of this methodology is an auto-regressive distributed lag model of order one with two variables:

$$y_t = \beta_0 + \beta_1 y_{t-1} + \gamma_0 x_t + \gamma_1 x_{t-1} + u_t \tag{4}$$

and if other essential factors are assumed, then with three (or more) variables:

$$y_t = \beta_0 + \beta_1 y_{t-1} + \gamma_0 x_t + \gamma_1 x_{t-1} + \delta_0 z_t + \delta_1 z_{t-1} + u_t \tag{5}$$

where y_t is the average weekly retail price of gasoline or diesel in time t; x_t is the average weekly crude oil price in time t; z_t is another relevant regressor in time t (for example, the price of another fuel); u_t is a stochastic term in time t and β_0, β_1, γ_0, γ_1, δ_0, and δ_1 are unknown parameters of this regression model.

The model (4) can be rewritten as the error correction model (Engle et al., 1987):

$$\Delta y_t = \beta_0 + \gamma_0 \Delta x_t + (\beta_1 - 1)\left[y_{t-1} - \frac{(\gamma_0 + \gamma_1)}{1-\beta_1} x_{t-1} \right] + u_t \tag{6}$$

and model (5) as the error correction (ECM) model:

$$\Delta y_t = \beta_0 + \gamma_0 \Delta x_t + \delta_0 \Delta z_t + (\beta_1 - 1)\left[y_{t-1} - \frac{(\gamma_0 + \gamma_1)}{1-\beta_1} x_{t-1} - \frac{(\delta_0 + \delta_1)}{1-\beta_1} z_{t-1} \right] + u_t \tag{7}$$

which contains the original (one period-lagged) variables in the levels (deviations from the long-run equilibrium) and their first differences (the short-run relationship). Suppose a positive unit change of the regressor has an identical influence on the regressand as a negative unit change. In that case, distinguishing between them is not needed. The overall response with one parameter for one regressor, as in the reversible models (6) and (7) can be estimated. If this restriction is not valid, the estimation results can be improved by specifying increases ($\Delta^+ x_t$ and $\Delta^+ z_t$) and decreases ($\Delta^- x_t$ and ($\Delta^- z_t$) of the explanatory variables as separate variables and also by separating the positive and negative deviations from the long-run equilibrium relationship.

The asymmetric irreversible error correction model (Granger et al., 1989):

$$\Delta y_t = \beta_0 + \gamma_0^+ \Delta^+ x_t + \gamma_0^- \Delta^- x_t + \lambda^+ e_{t-1} \times D(e_{t-1} > 0) + \lambda^- e_{t-1} \times D(e_{t-1} \le 0) + u_t \tag{8}$$

where $e_{t-1} = y_{t-1} - \frac{(\gamma_0 + \gamma_1)}{1-\beta_1} x_{t-1}$ is one period-lagged deviation from the long-run equilibrium relationship; $D(e_{t-1} > 0)$ is a dummy variable that equals one if $e_{t-1} > 0$ and equals zero otherwise; $D(e_{t-1} \le 0)$ is a dummy variable that equals one if $e_{t-1} \le 0$ and equals zero otherwise; λ^+ and λ^- are the corresponding adjustment parameters.

The asymmetric irreversible error correction (A-ECM) model is in the form:

$$\Delta y_t = \beta_0 + \gamma_0^+ \Delta^+ x_t + \gamma_0^- \Delta^- x_t + \delta_0 \Delta z_t + \lambda^+ e_{t-1} \times D(e_{t-1} > 0) + \lambda^- e_{t-1} \times D(e_{t-1} \le 0) + u_t \tag{9}$$

where $e_{t-1} = y_{t-1} - \frac{(\gamma_0 + \gamma_1)}{1-\beta_1} x_{t-1} - \frac{(\delta_0 + \delta_1)}{1-\beta_1} z_{t-1}$ is one period-lagged deviation from the long-run equilibrium relationship; $D(e_{t-1} > 0)$ is a dummy variable that equals one if $e_{t-1} > 0$ and equals zero otherwise; $D(e_{t-1} \le 0)$ is a dummy variable that equals one if $e_{t-1} \le 0$ and equals zero otherwise; λ^+ and λ^- are the corresponding adjustment parameters, β_0, γ_0^+, γ_0^-, and δ_0 are also parameters of this regression model.

The models (6) and (7) are obtained from models (8) and (9) using restrictions $\lambda^+ = \lambda^-$ and $\gamma_0^+ = \gamma_0^-$. The rejection of the hypothesis $\lambda^+ = \lambda^-$ (LR Symmetry) indicates asymmetry in adjusting the long-run equilibrium. The rejection of the hypothesis $\gamma_0^+ = \gamma_0^-$ (SR Symmetry) indicates short-term adjustment asymmetry. Both hypotheses can also be tested jointly. In cases where models (4) and (5) have a more extensive dynamic structure (γ_i), models (8) and (9) will also be more extensive.

Assume that there are cointegrating relationships between the crude oil price and the retail fuel prices for gasoline and diesel, individually or jointly. The vector error correction model (VECM) should be used to look for a long-term equilibrium relationship.

Similarly, the single-equation error correction model is auto-regressive, so the vector error correction model is a vector auto-regressive model. It can be shown by the vector auto-regressive model of order two denoted VAR(2):

$$\mathbf{y_t} = \mathbf{\Phi D_t} + \mathbf{\Pi}_1 \mathbf{y_{t\text{-}1}} + \mathbf{\Pi}_2 \mathbf{y_{t\text{-}2}} + \mathbf{u_t} \tag{10}$$

where $\mathbf{y_t}$ is the vector of variables (gasoline, diesel and oil prices) in time t; $\mathbf{D_t}$ is the matrix of deterministic terms (constant, trend, …) in time t; $\mathbf{u_t}$ is the vector of stochastic terms in time t and $\mathbf{\Phi}$, $\mathbf{\Pi_1}$ and $\mathbf{\Pi_2}$ are the matrices of unknown parameters of this model.

The model (10) can be rewritten as the vector error correction model (*VECM*) of order one:

$$\Delta \mathbf{y_t} = \mathbf{\Phi D_t} + \boldsymbol{\alpha\beta}^T \mathbf{y_{t\text{-}1}} + \mathbf{\Phi}_1 \Delta \mathbf{y_{t\text{-}1}} + \mathbf{u_t} \tag{11}$$

where $\boldsymbol{\alpha\beta}^T = \left(\mathbf{\Pi}_1 + \mathbf{\Pi}_2 - \mathbf{I}\right)$ and $\mathbf{\Phi}_1 = -\mathbf{\Pi}_2$. Matrix $\boldsymbol{\beta}$ is called a co-integration matrix with co-integration vectors as columns and matrix $\boldsymbol{\alpha}$ is called a loading matrix. The test of co-integration in VECM is realized by Johansen's procedure (Johansen, 1988, 1991) by the lambda trace statistics depending on the specification of the deterministic components $\mathbf{\Phi D_t}$ of model (11).

The asymmetric form of this irreversible vector error correction (*A-VECM*) model is:

$$\boldsymbol{\Delta y_t} = \boldsymbol{\Phi D_t} + \boldsymbol{\alpha}^+[\boldsymbol{\beta}^T \boldsymbol{y}_{t-1} \odot D(\boldsymbol{\beta}^T \boldsymbol{y}_{t-1} > \mathbf{0})] + \boldsymbol{\alpha}^-[\boldsymbol{\beta}^T \boldsymbol{y}_{t-1} \odot D(\boldsymbol{\beta}^T \boldsymbol{y}_{t-1} \le \mathbf{0})] + \boldsymbol{\Phi}_1^+ \boldsymbol{\Delta}^+ \boldsymbol{y}_{t-1} + \boldsymbol{\Phi}_1^- \boldsymbol{\Delta}^- \boldsymbol{y}_{t-1} + \boldsymbol{u_t} \tag{12}$$

where the multiplication operation $\odot$ in square brackets of model (12) does not represent the matrix product, but the element-wise product (product of elements in the same positions); $\boldsymbol{\beta}^T\mathbf{y_{t\text{-}1}}$ is the vector of one period lagged deviations from the long-run equilibrium relationships; $D(\boldsymbol{\beta}^T\mathbf{y_{t\text{-}1}} > \mathbf{0})$ is the vector of a dummy variable; its element equals 1 if corresponding element of $\boldsymbol{\beta}^T\mathbf{y_{t\text{-}1}}$ is positive and equals 0 otherwise; similarly $D(\boldsymbol{\beta}^T\mathbf{y_{t\text{-}1}} \le \mathbf{0})$ is the vector of a dummy variable; its element equals 1 if corresponding element of $\boldsymbol{\beta}^T\mathbf{y_{t\text{-}1}}$ is not positive and equals 0 otherwise; $\boldsymbol{\alpha}^+$ and $\boldsymbol{\alpha}^-$ are the loading matrices of corresponding adjustment parameters and $\mathbf{\Phi_1}^+$ and $\mathbf{\Phi_1}^-$ are also matrices with some pairs of the asymmetric parameters of this model. Model (11) is obtained from model (12) using restrictions $\mathbf{\Phi_1}^+ = \mathbf{\Phi_1}^-$ and $\boldsymbol{\alpha}^+ = \boldsymbol{\alpha}^-$.

Threshold Autoregressive Cointegration Models

Engle et al. (1987) approach is based on a symmetric long-run relationship. A different solution to the problem than Granger et al. (1989) was proposed by Enders et al. (1998), who introduced Threshold Autoregressive Cointegration (TAR). If the adjustment to the long-run equilibrium is asymmetric, the cointegration test is mis-specified. To overcome the problem, Enders et al. (2001) replace the standard augmented Dickey-Fuller test equation with the following threshold autoregressive process:

$$\Delta e_t = I_t \rho_1 e_{t-1} + \left(1 - I_t\right) \rho_2 e_{t-1} + \varepsilon_t \tag{13}$$

where e_t is the deviation from the long-run equilibrium relationship (residual). If the errors are serially correlated, equation (13) can be augmented with the lagged differences of e_t as in the standard augmented Dickey-Fuller test.

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

$$I_t = 1 \text{ if } e_{t-1} > 0 \text{ and } I_t = 0 \text{ otherwise} \tag{14}$$

alternatively, it is defined to depend on the lagged values of the first differences of residuals:

$$I_t = 1 \text{ if } \Delta e_{t-1} > 0 \text{ and } I_t = 0 \text{ otherwise} \tag{15}$$

The relationships (13) and (14) are called TAR cointegration. In contrast, the relationships (13) and (15) are known as momentum TAR (or M-TAR) cointegration. In M-TAR models, the threshold is placed on the variation of e_{t-1} rather than on e_{t-1}.

The null hypothesis $\rho_1 = \rho_2 = 0$ of no cointegration can be tested through an F test. The adjustment is symmetric for nonzero $\rho_1 = \rho_2$; thus, the Engle-Granger approach is a special case of (13) and (14).

In the case of the rejection of the null hypothesis in (13), the analysed variables are cointegrated, and the asymmetric ECM representation can be written as:

$$\Delta y_t = \lambda_{up} e_{t-1}^{up} + \lambda_{down} e_{t-1}^{down} + \sum_{i=0}^{p} \gamma_i \Delta x_{t-i} + \sum_{i=0}^{r} \delta_i \Delta z_{t-i} + \sum_{i=1}^{s} \alpha_i \Delta y_{t-i} + u_t \tag{16}$$

where $e_{t-1}^{up} = I_t e_{t-1}$, $e_{t-1}^{down} = (1 - I_t) e_{t-1}$. When ρ_1 is less than ρ_2, the increases tend to persist, whereas the decreases tend to revert quickly toward equilibrium.

Adjustment Cost Function in Linear-Exponential Form

In this part, the linear exponential adjustment cost function form is presented, and the reaction function specification is derived, as proposed by the authors in their earlier papers (Szomolanyi et al., 2020, 2022b). Consider the function in the form:

$$F\left[p_t, E_{t-1}(c_t)\right] = \frac{-\gamma\left[p_t - kE_{t-1}(c_t)\right] + \exp\left\{\gamma\left[p_t - kE_{t-1}(c_t)\right]\right\} - 1}{\gamma^2} \tag{17}$$

where p_t is the retail gasoline or diesel price, c_t is the crude oil price, E_{t-1} denotes the expectation conditional upon the information available at the time t–1, k is the technology coefficient. If the asymmetry coefficient γ is negative, the negative difference $p_t - kE_{t-1}(c_t)$ is costlier for the price-making firm than the positive.

The first-order condition of the firm choosing the output price to minimise the cost function (17) is the general form of the firm's reaction function.

$$\frac{-1 + \exp\left\{\gamma\left[p_t - kE_{t-1}(c_t)\right]\right\}}{\gamma} = 0 \tag{18}$$

By applying the l'Hôpital's rule, if γ tends to zero, the reaction function is linear:

$$\lim_{\gamma \to 0}(p_t) = kE_{t-1}(c_t) \tag{19}$$

Performing a second-order Taylor expansion of the exponential terms in (18), solving the equation for p_t and, prior to the generalised method of moments estimation (GMM) of the short-run relation, replacing the expected values with actual, and taking the first differences of the relation, the econometric specification of the firm's price reaction function is obtained:

$$\Delta p_t = k\Delta c_t - \frac{1}{2}\gamma\Delta\left[(p_t - kc_t)^2\right] + u_t \tag{20}$$

where Δ denotes the first difference operator, u_t is a stochastic term containing the first differences of terms of the third or higher orders of the expansion.

Assuming that the changes in the input prices Δc_t are a normally distributed process with zero mean and variance σ^2, the estimates of the average price biases are:

$$E(\Delta p_t) = -\frac{k^2\gamma}{2}\sigma^2 \tag{21}$$

The orthogonality conditions implied by the rational expectation hypothesis make the GMM a natural candidate to estimate the (20). Standard errors have been computed using the Newey-West procedure. The most important feature of the procedure explained by Newey et al. (1987) is its consistency in the presence of both heteroskedasticity and the autocorrelation of unknown forms.

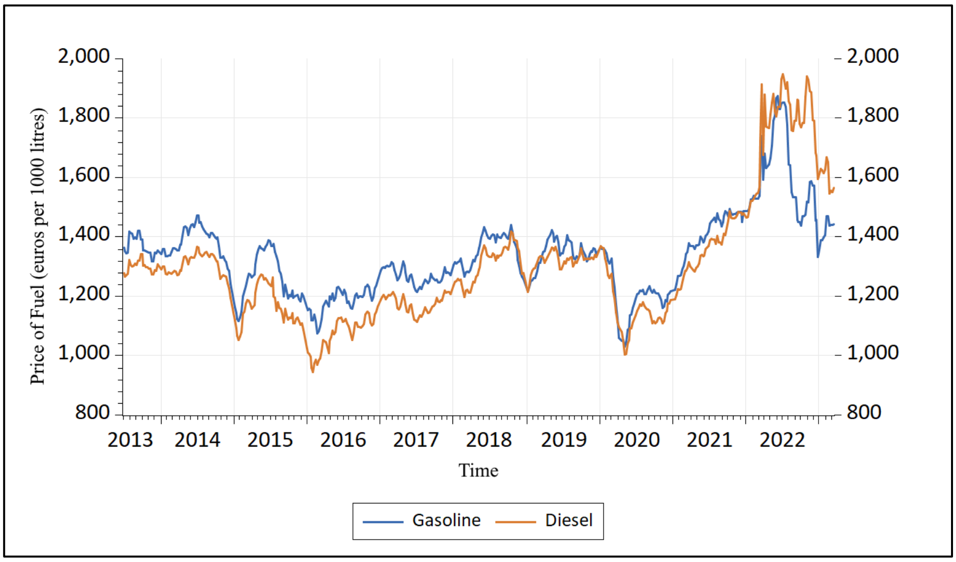
If retail gasoline and diesel prices are correlated, the relationship (20) can be estimated as the system of two equations, each for the given fuel price. The correlation is tested using the Breusch-Pagan test (Breusch et al., 1980).

Data

The analysis uses the weekly Croatian retail gasoline and diesel price data from the European Commission Weekly Oil Bulletin (energy.ec.europa.eu, 2023) and the daily Spot Brent crude oil prices obtained from the US Energy Information Administration website (eia.gov, 2023). The crude oil prices were converted to euros per 1000 litres using the daily exchange rate data series obtained from the European Central Bank (ECB.europa.eu, 2023). The daily data series were aggregated to the weekly by averaging.

The analysis uses the entire sample of published data, including the COVID-19 pandemic period. However, the common practice is sometimes different. Some studies split the sample into two or more sets to find evidence for the rockets' and feathers' effects (Bagnai et al., 2018; Bumpass et al., 2019; Cipcic, 2021).

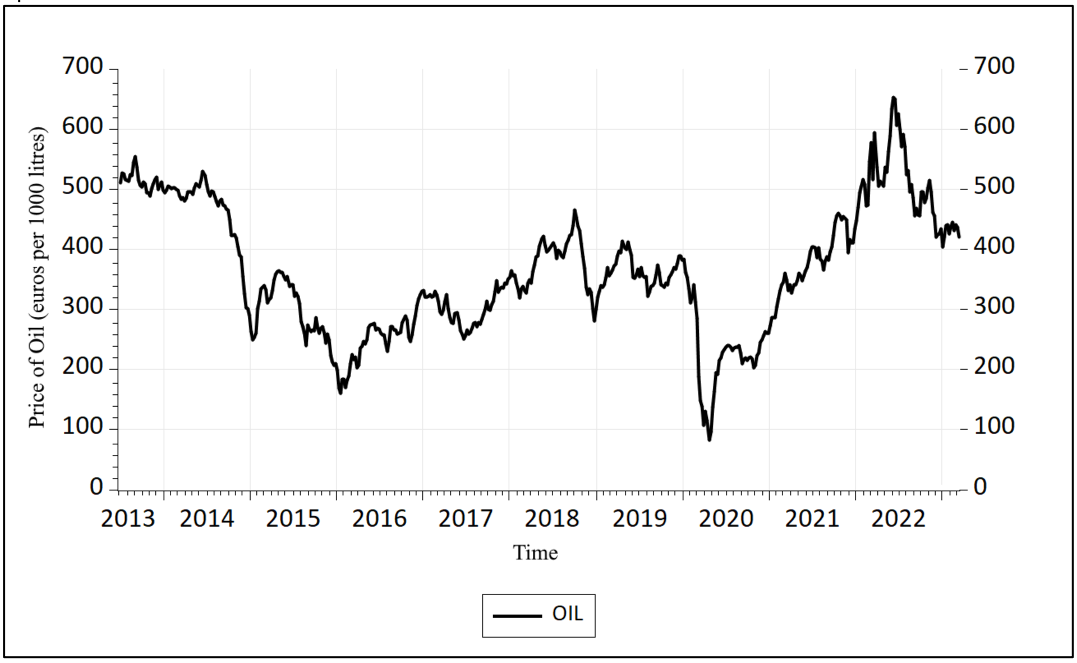
Figure 1
Retail Gasoline and Diesel Prices in Croatia – Time Series



Source: Author's illustration

Figure 1 graphically presents the weekly Croatian retail gasoline and diesel price data from January 2013 to March 2023, representing 483 data observations. The time series volatility has enlarged in the post-pandemic period. Since this period, the retail diesel price has been higher than the gasoline price almost continuously, which has happened only sporadically before. At the same time, both time series have practically identical courses.

Figure 2
Spot Brent Crude Oil Prices – Time Series



Source: Author's illustration

Figure 2 shows the weekly Spot Brent crude oil prices during the same period as fuel prices. All three data series have a very similar course, with a significant drop at the beginning of the pandemic period. Nevertheless, the decline in the third and fourth quarters of 2014, which continued slowly until 2016, was also interesting. The slowdown in global economic activity and shocks in the oil markets caused it. Fuel prices also copied this development.

When analysing data with similar characteristics, a changing trend, or volatility, it is customary to split the data set into several samples. Given the data used in the analysis, partitioning into changing trend periods would lead to selections with an insufficient data sample range, which is insufficient to use the Newey-West method and resistant to autocorrelation. Therefore, the entire range of available data is employed due to the adequacy of comparing individual procedures.

The given time series course indicates the non-stationarity of the processes generating these data. The non-stationarity was tested using the augmented Dickey-Fuller (ADF) unit root test (Dickey et al., 1981). The results of the ADF test for individual deterministic specifications (constant and trend, only constant and none) are shown for each variable and its first difference in Table 1. Test results in bold mean the rejection of non-stationarity (the first difference of all the data used). The null hypothesis of non-stationarity of the original variables was not rejected. This means that all prices are generated by non-stationary processes integrated by order one and are suitable for the search for co-integration, which is also indicated by an interconnected course.

Table 1
Augmented Dickey-Fuller Unit Root Test (Test of Non-stationarity)

Gasoline	Const	Trend	y_{t-1}	ΔGasoline	Const	Trend	y_{t-1}
const + trend	26.943	0.011	-0.022	const + trend	-0.834	0.002	**-0.911**
[τ stat]	[2.388]	[1.352]	[-2.511]	[τ stat]	[-0.354]	[0.292]	**[-18.90]**
const	23.436		-0.018	const	-0.237		**-0.911**
[τ stat]	[2.133]		[-2.149]	[τ stat]	[-0.202]		**[-18.92]**
none			0.0002	none			**-0.911**
[τ stat]			[-0.263]	[τ stat]			**[-18.94]**
Diesel	**Const**	**Trend**	y_{t-1}	**ΔDiesel**	**Const**	**Trend**	y_{t-1}
const + trend	18.594	0.019	-0.018	const + trend	-0.704	0.006	**-1.119**
[τ stat]	[2.019]	[1.748]	[-2.191]	[τ stat]	[-0.254]	[0.596]	**[-23.36]**
const	13.135		-0.010	const	0.731		**-1.118**
[τ stat]	[1.513]		[-1.450]	[τ stat]	[0.529]		**[-23.37]**
none			0.0003	none			**-1.117**
[τ stat]			[0.289]	[τ stat]			**[-23.39]**
Oil	**Const**	**Trend**	y_{t-1}	**ΔOil**	**Const**	**Trend**	y_{t-1}
const + trend	4.163	0.004	-0.015	const + trend	-1.020	0.004	**-0.854**
[τ stat]	[1.508]	[0.849]	[-2.179]	[τ stat]	[-0.726]	[0.741]	**[-19.37]**
const	5.094		-0.014	const	-0.118		**-0.853**
[τ stat]	[2.011]		[-2.141]	[τ stat]	[-0.167]		**[-19.36]**
none			-0.001	none			**-0.853**
[τ stat]			[-0.752]	[τ stat]			**[-19.38]**

Note: Test results in bold mean the rejection of non-stationarity. Values of tau statistics are in square brackets.
Source: Authors' calculations

Results

The results of the models from each group of methodological procedures are presented in individual tables. The section structure comes from the methodology section structure.

Simple Models of Asymmetry

Table 2 shows the results of estimations of models with the primary determinant oil price (1) and with an extended model supplemented by a change in the price of the second fuel (2). In the models, the corresponding autoregressive errors are estimated by appropriate methods to eliminate the autocorrelation problem. Table 3 shows the estimation results of dynamic models with a gradually increasing number of oil price lags included in the model. Standard errors have been computed using the Newey-West procedure.

The significance of the statistical tests is indicated by the number of asterisks, with three indicating statistical significance at the 1% level, two at the 5%, and one at the 10%. The probability values are enclosed in square brackets. The estimated parameters of all models in Table 1 are statistically significant, at least at the 5% significance level. Price symmetry was rejected in the models for gasoline prices but not diesel prices.

Table 2
Simple Models of Croatia's Fuel Prices Asymmetry

Simple Models	γ_0^+	γ_0^-	δ_0	Symmetry
gasoline model (1)	0.384***	0.395***	---	***F* = 6.182**
(std. err.)	(0.050)	(0.053)	---	**[0.013]**
diesel model (1)	0.351***	0.362***	---	*F* = 3.095
(std. err.)	(0.064)	(0.069)	---	[0.079]
gasoline model (2)	0.103**	0.109**	0.597***	***F* = 4.413**
(std. err.)	(0.045)	(0.046)	(0.016)	**[0.036]**
diesel model (2)	0.067**	0.069**	0.808***	*F* = 0.119
(std. err.)	(0.030)	(0.032)	(0.185)	[0.731]

Note: Three asterisks indicate statistical significance at the 1% significance level, two at the 5%. Test results in bold mean the rejection of symmetry. The probability values are in square brackets.
Source: Authors' calculations

Table 3
Dynamic Models of Croatia's Fuel Prices Asymmetry

Models with 1 Lag	Last Coefficients		Last Signif. Coefficients		Symmetry
	γ_1^+	γ_1^-		γ_1^-	
gasoline model (3)	-0.103	0.286*	-	0.286*	*F* = 0.373
(std. err.)	(0.227)	(0.167)	-	(0.167)	[0.542]
diesel model (3)	-0.195	-0.003	-	-	*F* = 0.024
(std. err.)	(0.350)	(0.263)	-	-	[0.878]
Models with 2 Lags	γ_2^+	γ_2^-	γ_2^+	γ_2^-	**Symmetry**
gasoline model (3)	1.049***	0.815***	1.049***	0.815***	*F* = 0.001
(std. err.)	(0.175)	(0.125)	(0.175)	(0.125)	[0.515]
diesel model (3)	1.357***	0.849***	1.357***	0.849***	*F* = 0.915
(std. err.)	(0.371)	(0.230)	(0.371)	(0.230)	[0.393]
Models with 3 Lags	γ_3^+	γ_3^-	γ_2^+	γ_3^-/γ_2^-	**Symmetry**
gasoline model (3)	-0.075	0.412**	1.100***	0.412**	*F* = 0.584
(std. err.)	(0.115)	(0.200)	(0.179)	(0.200)	[0.445]
diesel model (3)	-0.082	0.285	1.397***	0.800***	*F* = 0.342
(std. err.)	(0.197)	(0.204)	(0.386)	(0.253)	[0.559]
Models with 4 Lags	γ_4^+	γ_4^-	γ_2^+	γ_4^-/γ_2^-	**Symmetry**
gasoline model (3)	-0.184	0.292***	1.134***	0.292***	*F* = 2.634
(std. err.)	(0.199)	(0.113)	(0.190)	(0.113)	[0.105]
diesel model (3)	-0.097	0.146	1.414***	0.770***	*F* = 0.177
(std. err.)	(0.329)	(0.209)	(0.417)	(0.210)	[0.674]

Note: Three asterisks indicate statistical significance at the 1% significance level, two at the 5%, and one at the 10%. Test results in bold mean the rejection of symmetry. The probability values are in square brackets.
Source: Authors' calculations

As in the tables above, asterisks denote the statistical significance levels, and the probability values are square brackets. The significance of the parameters for lagged changes in the crude oil price indicates in the diesel price equation the dynamics of a maximum of two periods during both an increase and a price decrease, and in the gasoline price equation, two periods during a price increase, but up to four periods during a price decrease.

This result may indicate a certain temporal asymmetry of the reaction with gasoline. However, based on the comparison of the cumulated response, price symmetry was not rejected in any dynamic model either for the price of gasoline or for the diesel.

Error Correction Models

Table 4 shows the results of cointegration tests of models with a correction term (6) and models supplemented by the change in the price of the second fuel (7). Asterisks and square brackets denote the statistical significance and probability values. Engle-Granger (Engle et al., 1987) and Phillips-Ouliaris test (Phillips et al., 1990) in bold mean the non-rejection of cointegration. Both test procedures are based on tests of the unit root of the residuals of the cointegrating relationship. Engle and Granger use the augmented Dickey-Fuller test, and Phillips and Ouliaris use the Phillips-Perron statistics based on Newey-West estimate of standard error. In addition to these tests, the loading parameters $\lambda = \beta_1 - 1$ for (6) and (7) are estimated. In all models for gasoline prices, the variables are cointegrated. In contrast, in the models for diesel prices, cointegration was not rejected only in the trend model with no additional explanatory variables added.

Table 4
Error Correction Models of Croatia's Fuel Prices

ECM Gasoline	Engle-Granger	Phillips-Ouliaris	$\lambda = \beta_1 - 1$
model (6)	**τ = -4.429**	**τ = -4.163**	**-0.032*****
[p value]	[0.002]	[0.004]	(0.007)
model (6) + trend	**τ = -5.832**	**τ = -5.733**	**-0.072*****
[p value]	[0.000]	[0.000]	(0.000)
model (7)	**τ = -4.377**	**τ = -4.627**	**-0.079*****
[p value]	[0.002]	[0.001]	(0.017)
model (7) + trend	**τ = -4.662**	**τ = -4.839**	**-0.082*****
[p value]	[0.004]	[0.002]	(0.017)
ECM Diesel	**Engle-Granger**	**Phillips-Ouliaris**	$\lambda = \beta_1 - 1$
model (6)	τ = -2.991	τ = -2.586	-0.013
[p value]	[0.114]	[0.245]	(0.011)
model (6) + trend	**τ = -5.328**	**τ = -5.026**	**-0.063*****
[p value]	[0.000]	[0.001]	(0.021)
model (7)	τ = -2.939	τ = -3.342	-0.031*
[p value]	[0.270]	[0.130]	(0.017)
model (7) + trend	τ = -3.442	τ = -4.027	-0.066***
[p value]	[0.226]	[0.067]	(0.025)

Note: Three asterisks indicate statistical significance at the 1% significance level, two at the 5%, and one at the 10%. Engle-Granger (Engle et al., 1987) and Phillips-Ouliaris test (Phillips et al., 1990) in bold mean the non-rejection of cointegration. The probability values are in square brackets.
Source: Authors' calculations

Only models with unrejected cointegration are converted to asymmetric models (8) and (9). With their help, short-term and long-term symmetry is tested separately and then in a joint hypothesis. Table 5 shows the results of estimations and tests of asymmetric models with the correction terms (8) and (9). The test results in bold mean the rejection of long-run (LR), short-run (SR) or both symmetries.

Table 5
Asymmetric Error Correction Models of Croatia's Fuel Prices

A-ECM Gasoline	λ^+	λ^-	LR Symmetry	SR Symmetry	Both Sym.
model (8)	-0.044*	-0.016	*F* = 0.738	*F* = 0.537	*F* = 1.268
[p value]	(0.025)	(0.013)	[0.391]	[0.464]	[0.282]
model (8) + trend	-0.066***	-0.077***	*F* = 0.072	*F* = 1.711	*F* = 0.924
[p value]	(0.025)	(0.023)	[0.788]	[0.192]	[0.398]
model (9)	-0.071**	-0.077***	*F* = 0.017	*F* = 3.232	*F* = 1.629
[p value]	(0.028)	(0.025)	[0.895]	[0.073]	[0.197]
model (9) + trend	-0.075***	-0.079***	*F* = 0.006	*F* = 3.255	*F* = 1.704
[p value]	(0.028)	(0.025)	[0.939]	[0.072]	[0.183]
A-ECM Diesel	λ^+	λ^-	**LR Symmetry**	**SR Symmetry**	**Both Sym.**
model (8) + trend	-0.091**	-0.039	*F* = 0.926	***F* = 5.199**	*F* = 2.665
[p value]	(0.037)	(0.028)	[0.336]	[0.023]	[0.071]

Note: Three asterisks indicate statistical significance at the 1% significance level, two at the 5%, and one at the 10%. Test results in bold mean the rejection of long-run (LR), short-run (SR) or both symmetries. The probability values are in square brackets.
Source: Authors' calculations

Table 6 shows the key results of the cointegration test and the selection of the vector error correction model's deterministic scheme (11). The first part shows the values of Akaike information criteria (AIC) for the Johansen procedure's three most frequently used deterministic schemes. Based on the results of AIC and lambda trace statistics in the middle part, a model with two cointegrating vectors and a fourth deterministic scheme with a linear trend in their cointegrating relationships and an unrestricted constant was chosen.

Table 6
Vector Error Correction Model of Croatia's Fuel Prices

Akaike Information Criteria (AIC) by Rank and Model			
	Deterministic scheme		
Rank or No. of CEs	**2**	**3**	**4**
0	25.939	25.951	25.951
1	25.920	25.928	25.912
2	25.929	25.933	**25.896**
3	25.951	25.951	25.912
Cointegration Rank Trace Test - 4th deterministic scheme			
Hypoth. No. of CE(s)	**Eigenvalue**	**Trace Stat**	**p value**
None	0.0700	58.743	[0.001]
At most 1	0.0472	27.814	[0.028]
At most 2	0.0168	**7.224**	[0.322]
Test of restrictions:	**1.1260**	p value	[0.890]
Restricted estimate	**Gasoline** (i = 1)	**Diesel** (i = 2)	**Oil** (i = 3)
α_{i1}	-0.080***	0	0
(std. err.)	(0.016)	-	-
α_{i2}	0	-0.077***	0
(std. err.)	-	(0.016)	-

Note: Three asterisks indicate statistical significance at the 1% significance level. The probability values are in square brackets. The value of AIC in bold is the minimum of all AIC values.
Source: Authors' calculations

Furthermore, estimates of the loading matrix parameters supplemented after introducing acceptable restrictions have been included. The first cointegrating vector expressing the long-run relationship between the price of gasoline and oil affects only the equation of the price of gasoline. On the contrary, the second cointegrating vector expressing the long-run relationship between the price of diesel and oil affects only the equation of the price of diesel. This fact means that the parameters from (11) α_{21} and α_{31}, together with α_{12} and α_{32}, equal 0, which does not reject the restriction test equal to 1.126 with a p-value of 0.890 displayed at the bottom of Table 6.

Table 7 presents the results of testing short-run and long-run symmetry. Finally, both hypotheses are joined in the gasoline and diesel price equation of the estimated model from Table 6. Test results in bold mean the rejection of long-run (LR), short-run (SR), or both symmetries. The probability values are in square brackets. At the 5% significance level, only short-run symmetry in the gasoline equation is rejected.

Table 7
Asymmetric Irreversible Vector Error Correction Model of Croatia's Fuel Prices

SR symmetry	Gasoline	Diesel
statistics	***F* = 5.470**	*F* = 2.811
[p value]	[0.019]	[0.094]
LR symmetry	**Gasoline**	**Diesel**
statistics	*F* = 0.824	*F* = 0.677
[p value]	[0.364]	[0.411]
Both symmetries	**Gasoline**	**Diesel**
statistics	*F* = 5.474	*F* = 2.952
[p value]	[0.065]	[0.229]

Note: Test results in bold mean the rejection of long-run (LR), short-run (SR) or both symmetries. The probability values are in square brackets. *Source:* Authors' calculations

Threshold Autoregressive Cointegration Models

Table 8 shows the results of symmetry testing in TAR and M-TAR models. The most appropriate model was selected using the Schwarz information criterion from models with differences lagged by up to 8 periods. Standard errors have been computed using the Newey-West procedure. The table presents estimates with one (crude oil prices) and two explanatory variables (crude oil and other fuel prices). The latter are marked as Cross TAR and Cross M-TAR.

Table 8
TAR and M-TAR Models of Croatia's Fuel Prices Asymmetry

TAR Models	$\rho_1 = \rho_2 = 0$	$\rho_1 = \rho_2$	M-TAR Models	$\rho_1 = \rho_2 = 0$	$\rho_1 = \rho_2$
gasoline	***F* = 7.010**	*F* = 0.462	gasoline	***F* = 5.842**	*F* = 0.558
[p value]	[0.001]	[0.497]	[p value]	[0.003]	[0.455]
gasol. + trend	***F* = 12.10**	*F* = 0.105	gasol. + trend	***F* = 8.887**	*F* = 0.280
[p value]	[0.000]	[0.746]	[p value]	[0.000]	[0.597]
diesel	***F* = 4.560**	*F* = 0.006	diesel	*F* = 1.896	*F* = 0.614
[p value]	[0.011]	[0.940]	[p value]	[0.152]	[0.434]
diesel + trend	***F* = 5.349**	*F* = 0.055	diesel + trend	***F* = 4.621**	*F* = 0.015
[p value]	[0.005]	[0.815]	[p value]	[0.010]	[0.903]
Cross TAR	$\rho_1 = \rho_2 = 0$	$\rho_1 = \rho_2$	**Cross M-TAR**	$\rho_1 = \rho_2 = 0$	$\rho_1 = \rho_2$
gasoline	***F* = 5.679**	*F* = 0.001	**gasoline**	***F* = 10.06**	***F* = 8.255**
[p value]	[0.004]	[0.981]	[p value]	[0.000]	[0.004]
gasol. + trend	***F* = 6.955**	*F* = 0.016	**gasol. + trend**	***F* = 9.203**	***F* = 6.423**
[p value]	[0.001]	[0.900]	[p value]	[0.000]	[0.012]
diesel	***F* = 4.292**	*F* = 2.133	diesel	*F* = 3.010	*F* = 0.549
[p value]	[0.015]	[0.145]	[p value]	[0.051]	[0.459]
diesel + trend	***F* = 3.737**	*F* = 0.098	diesel + trend	***F* = 4.389**	*F* = 0.327
[p value]	[0.025]	[0.755]	[p value]	[0.013]	[0.568]

Note: F test result in bold represents the rejection of null hypothesis. The probability values are in square brackets. *Source:* Authors' calculations

F test result in bold represents the rejection of null hypothesis. The probability values are in square brackets. In the estimations of M-TAR models, in which the price of diesel appears as the second explanatory variable (Cross M-TAR), the symmetrical adjustment of gasoline prices is rejected (in the other TAR and M-TAR models, the symmetrical adjustment of fuel prices is not rejected).

Adjustment Cost Function in Linear-Exponential Form

Tables 9 and 10 show the estimation results using GMM. Table 9 shows the single-equation (20) estimates, and Table 10 shows the system estimate. The lagged first differences of retail fuel and crude oil prices (up to two lags) are used as instruments. Asterisks and square brackets denote the statistical significance and probability values. *J* test result in bold means the rejection of orthogonality. The probability values are in square brackets.

Table 9
Results of the Linex Adjustment Cost Function Approach – single equations GMM

GMM Models	*k*	γ	*J*	Bias
gasoline	1.579***	–0.001***	0.521	**0.3145**
(std.err.)	(0.186)	(0.0001)	[0.470]	
diesel	1.470***	–0.001***	0.00001	**0.2122**
(std.err.)	(0.392)	(0.0002)	[0.997]	

Note: Three asterisks indicate statistical significance at the 1% significance level. J test result in bold means the rejection of orthogonality. The probability values are in square brackets.
Source: Authors' calculations

The biases (21) in € per 1000 litres are computed in the tables' last column. Table 10 presents the test of the mutual correlation of retail fuel prices. According to the value of BP = 206.76, the noncorrelation of the stochastic terms in both equations are not reject.

Table 10
Results of the Linex Adjustment Cost Function Approach – system GMM

GMM Model	*k*	γ	*J*	Bias
gasoline	1.537***	–0.001***	0.004	**0.2944**
(std.err.)	(0.205)	(0.0001)	[0.998]	
diesel	1.352***	–0.001***		**0.1716**
(std.err.)	(0.433)	(0.0002)	BP = 206.76	

Note: Three asterisks indicate statistical significance at the 1% significance level. J test result in bold means the rejection of orthogonality. The probability values are in square brackets.
Source: Authors' calculations

The reaction function specification coefficients *k* and γ are statistically significant in all cases. As the hypothesis $\gamma < 0$ is not rejected, an asymmetry in the retail fuel price adjustment is indicated. According to the computed bias values in both estimates, the degree of asymmetry is estimated to be higher when gasoline prices are adjusted.

A challenge is to fill the international comparison results gap. A natural way is to compare the Croatian results with the European ones. However, there is no data on average fuel prices in the Eurozone or EU average. The possibility is to compute the average European retail fuel prices. The European Commission Weekly Oil Bulletin publishes the retail fuel prices for all European Union member states, but averaging this data without state weights does not respect the reality. The European Union is a heterogeneous group of economies with different business power and characteristics. The better way is to use a weighted average. Such weights could be the retail fuel consumption or GDP in constant prices for each economy. Unfortunately, the data exists only in quarter or annual frequencies. Due to the robustness of the results, this section uses several possibilities to compute the European average retail fuel prices: the simple average, median, average weighted by fuel consumption, and average weighted by GDP in constant prices.

The yearly fuel consumption and the quarterly GDP at constant market prices are gathered from the Eurostat database. In the weights' data series, the value for each week is the value for a corresponding year or quarter. Using the data, Brent crude oil prices, and the system GMM, the estimates of (20) follow. The results are summarised in Table 11.

Table 11
Results of the Linex Adjustment Cost Function Approach using Different Computations of the Average Retail Fuel Prices in the European Union – system GMM

EU average	k	γ	**Bias**	**EU median**	k	γ	**Bias**
gasoline	1.494***	–0.001***	**0.3666**	gasoline	1.734***	–0.001***	**0.5913**
diesel	1.497***	–0.001***	**0.4076**	diesel	1.638***	–0.001***	**0.5268**
EU average weighted by consumption	k	γ	**Bias**	**average weighted by GDP**	k	γ	**Bias**
gasoline	1.446***	–0.001***	**0.3174**	gasoline	1.549***	–0.001***	**0.3757**
diesel	1.469***	–0.001***	**0.3869**	diesel	1.631***	–0.001***	**0.4644**

Note: Three asterisks indicate statistical significance at the 1% significance level.
Source: Authors' calculations

The table is divided into four parts according to the computations of the European averages of the retail fuel prices. The reaction function specification coefficients k and γ are statistically significant in all cases. As the hypothesis $\gamma < 0$ is not rejected, an asymmetry in the retail fuel price adjustment is indicated. The estimated price biases (21) are in the last columns.

Discussion

Deltas et al. (2020) noted that different methods, data samples, and frequencies generate different results when examining the crude oil price pass-through to retail fuel prices. The same is true of this analysis.

The result of simple static models implies that gasoline pricing is asymmetric, whereas diesel pricing is not. However, according to the simple dynamic model, both commodity prices change symmetrically. Only in one ECM model is the short-run symmetric adjustment of retail diesel prices rejected (not the long-run) in the specification with a trend in which the gasoline price does not appear as another explanatory variable. On the contrary, vector error correction models reject the short-run asymmetric price adjustment of retail gasoline prices. In the estimations of M-TAR models, in which the price of diesel appears as the second explanatory variable (Cross M-TAR), the symmetrical adjustment of gasoline prices is rejected (in the other TAR and M-TAR models, the symmetrical adjustment of fuel prices is not rejected).

According to the Linex adjustment cost function approach, the retail fuel prices adjust asymmetrically. The corresponding price bias is slightly higher for gasoline. Price decreases are more sluggish than price increases. There are several theoretical explanations for this behaviour.

The first theoretical rationale, known as oligopolistic coordination theory, posits that firms in an oligopoly market structure attempt to maintain a tacit agreement with their competitors. They do this by responding asymmetrically to changes in oil prices, thereby signalling their commitment to the agreement.

The second explanation, rooted in the production and inventory adjustment cost, reveals a strategic decision-making process. Aware of the costliness of adjusting production and inventory levels, firms spread the adjustment over time.

The search theory, the third theoretical reason, sheds light on how firms exploit the high oil price volatility. This volatility allows them to leverage consumers' high search costs, temporarily increasing their margin after an oil price hike.

The last explanation is the theory of strategic interactions. Price makers try not to pit against themselves a minority of consumers who tend to "punish" firms for unjustified or insufficiently explained price increases. There is a tendency to use the crude oil price increase to adjust retail fuel prices by additional unexpected costs or the effects of an increase in demand.

Due to frequent changes in trends and volatilities, some authors using traditional approaches to estimating the asymmetries in pricing tend to split data samples (Bagnai et al., 2018; Bumpass et al., 2019; Cipcic, 2021). On the other hand, the Linex approach uses the GMM estimator, the advantages of which will be demonstrated with sufficiently large numbers of observations. Therefore, splitting the data samples is not recommended when applying the approach.

Cipcic (2021) was one of the few authors investigating the asymmetric transition of crude oil prices to retail gasoline prices in Croatia. The author used the ECM model with a specification in which the oil price appeared as the only explanatory variable. She did not reject the symmetrical adjustment of gasoline prices in Croatia, even when she split the data sample into two subsamples. The same method reached the same conclusion (Table 4, model (8)).

The study estimates a weekly bias of 29 cents per 1000 litre (0.029 cents per litre) for gasoline and 17 cents per 1000 litre (0.017 cents per litre) for diesel (based on the two-equation system estimate). It should be noted that the observations include a highly volatile period since the outbreak of the COVID-19 pandemic.

The paper's authors have so far applied the Linex approach of verifying asymmetric pricing of retail fuels on weekly data from Slovakia and the USA. Both studies confirmed the asymmetric reactions of gasoline and diesel prices to changes in oil prices. Using Slovak data on fuel prices in the period 2009 to 2019 and the Linex approach, Szomolanyi et al. (2020) estimated the average weekly bias of gasoline prices in the value of approximately 0.13 – 0.16 euros per 1000 litres and the average weekly bias of diesel prices in the value of approximately 0.20 - 0.22 euros per 1000 litres.

Using the same approach, Szomolanyi et al. (2022b) note the asymmetric formation of retail fuel prices in the US in US regions and selected US states and cities. Interestingly, the asymmetric response of wholesale fuel prices traded in the three major US ports is not confirmed. However, on the other hand, research finds asymmetric responses of retail fuel prices to changes in New York and Gulf port fuel prices. These findings are similar to those of Gosinski et al. (2020). The authors found no asymmetry between crude oil and Polish wholesale fuel prices. In contrast, Polish retail fuel prices' reactions to wholesale price changes have been asymmetric between 2000 and 2016. The subject of the authors' further research is to examine the asymmetric reactions of retail fuel prices in all EU states.

As expected, the estimated price biases in the EU differ according to how retail fuel prices are averaged. Unlike in Croatia, in the European Union, average diesel price bias estimates are higher than gasoline in all cases. Comparing the GMM system estimates, the average price biases are higher in the EU average.

The theoretical reasons for asymmetric pricing (oligopolistic coordination, production and inventory adjustment cost, search theory, and strategic interactions) are described in more detail in the Introduction. The theory of strategic interactions between a firm and its consumers is one of three theories explaining the price stickiness hypothesis used in New Keynesian monetary models.

Douglas et al. (2010) argue that the evidence of asymmetric reactions of retail fuel prices predicts that this theory is the main reason for the price stickiness hypothesis. The hypothesis can be formulated mathematically by adjusting the profit functions of companies by adjusting costs in the form of Linex. Adjustment and estimated parameter values can be applied in dynamic stochastic general equilibrium models (DSGE), which analyse the effects of monetary and fiscal policy and other shocks in Croatia.

Conclusion

The paper focused on the possibilities of quantifying the asymmetric transition of crude oil prices to Croatian retail fuel prices. The traditional, well-known procedures and the Linex approach formulating a non-linear adjustment cost function were used. This approach using the GMM estimator demands a large data set, so the data sample was not divided into more subsamples. Although the studied time series are subject to frequent changes in trends and volatilities, the Linex approach on Croatian data detects systematic price bias caused by asymmetric price reactions in the sense of rockets' and feathers' effects.

Among all the approaches, Linex is preferred because, as the methodological part demonstrated, the theoretical justifications for the asymmetric price transition correspond to the non-linear formulation of the adjustment cost function. Another advantage of the Linex approach is that it can be used to directly estimate the average price bias, which allows comparison across different periods and regions.

As Deltas et al. (2020) point out, traditional methods of investigating asymmetric pricing depend on the choice of frequency and range of data and the method used. According to these results and the results of the authors' studies (Szomolanyi et al., 2020, 2022a, 2022b), the Linex approach always confirmed price asymmetries. The differences were only in the size of the bias.

A comparison with the results from Cipcic (2021) also corresponds to this. One of the approaches applied in this study uses the same methodology as the author. It also does not reject the symmetrical adjustment of fuel prices in Croatia. The result obtained by the Linex approach is the opposite.

Asymmetric price reactions do not have to occur only in the fuel market but also in other markets, such as agricultural, energy, and others, for the analysis of which the presented approaches, including Linex, can be used. Considering the advantages of the Linex approach, the price asymmetries for different European fuel prices will be compared in future research.

Limitations of the research

The transition from crude oil prices to retail fuel prices is complex. Several entities are involved in the supplier-customer chain. However, the available data reflect prices only at the input and output of the entire chain. The study's limitation is that the originator of the asymmetries cannot be identified.

In addition, the econometric test specification is non-linear, which is associated with other limitations of the paper. The results of software iterative methods for calculating parameter estimates are sensitive to the choice of starting values. The results of the instrumental estimation of the specification depend on the choice of instruments. The choice is applied by testing the instruments' weakness, orthogonality, and regressors' endogeneity. Tests of weakness of instruments and endogeneity of regressors are commonly applied only in linear econometric models.

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