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Who is responsible for asymmetric fuel price adjustments? An application of the threshold cointegrated VAR model

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Who is responsible for asymmetric fuel price adjustments? An application of the threshold cointegrated VAR model

Emilia Gosińska, Katarzyna Leszkiewicz-Kędzior and Aleksander Welfe

Chair of Econometric Models and Forecasts, Faculty of Economics and Sociology, University of Lodz, Lodz, Poland

ABSTRACT
The purpose of the paper is to test the hypothesis about asymmetric price transmission between the fuel markets. The distribution chain is considered at three levels: the European wholesale market, the domestic wholesale market, and the domestic retail market. It is shown that between the European and domestic wholesale markets fuel prices adjust symmetrically and asymmetrically between the domestic wholesale market and the retail market. This finding confirms that the most probable cause of asymmetric price adjustments (especially in new EU member states) is the behaviour of petrol stations and not of oil companies. The empirical analysis is conducted using an appropriately modified Hansen-Seo method. The procedure, which has until recently been used to estimate bivariate threshold models, prevents the presence of the constant in the cointegrating vector entailing the risk of severe distortion of the estimation results. Moreover, the interpretation of a dummy variable present in the CVAR equation as a result of a data generating process distortion is limited by its presence in the cointegration space.

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1. Introduction

Asymmetric adjustments to the equilibrium paths are revealed in many economic processes. Earlier studies have shown that asymmetries especially frequently occur in price transmissions in the fuel market (see for instance, Al-Gudhea et al., 2007; Bacon, 1991; Chen et al., 2005; Galeotti et al., 2003; Grasso & Manera, 2007; Rahman, 2016) and in the market for agricultural products such as vegetables, meat, and dairy products (see, for instance, Abdulai, 2002; Chavas & Mehta, 2004; Griffith & Piggott, 1994; Miller & Hayenga, 2001; Mohanty et al., 1995), as well as in the transmission of official interest rates in interest rates on bank deposits and loans in the financial market (see, for instance, Gambacorta & Iannotti, 2007; Kovanen, 2011; Liu et al., 2009; Mangwengwende et al., 2011; Sznajderska, 2013).

The mechanism of asymmetry that is commonly observed by final consumers in the fuel market is the following: with the crude oil price rising in the global market, the

CONTACT Katarzyna Leszkiewicz-Kędzior katarzyna.kedzior@uni.lodz.pl Chair of Econometric Models and Forecasts, Faculty of Economics and Sociology, 3/5 POW Street, University of Lodz, 90-214 Lodz, Poland

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prices of oil-based products go up rapidly at petrol stations, or fall slowly when the former enter a downward trend (so-called rockets and feathers, see Bacon, 1991). The key factors that may contribute to the asymmetric fuel price adjustments include the strong market position of oil concerns operating under imperfect competition (usually in an oligopolistic market), refinery inventory management (the type of inventory costing method: FIFO or LIFO), fuel traders entering into tacit collusion at the wholesale and retail levels and consumers' high costs of acquiring information about fuel prices (see Borenstein et al., 1997; L'Oeillet & Lantz, 2009; Meyer & Von Cramon-Taubadel, 2004).

The Polish fuel market comprises two segments: the wholesale market where the petrol stations (intermediaries) purchase oil-based products from producers and the retail market where they deliver fuel products to final customers. The wholesale market has a duopolistic structure determined by the domination of two refineries, namely, PKN Orlen SA and Lotos SA Group. The retail market is represented by nearly 8000 petrol stations owned by Polish and foreign oil companies, independent operators and hypermarkets that act under monopolistic competition.

Empirical studies analyzing asymmetry effects based on high-frequency data concentrate on fuel price volatility, a consequence of which is the use of the GARCH-type models (see i.a. Balaguer & Ripollés, 2012; Bettendorf et al., 2003; De Salles, 2014; Zlatcu et al., 2015). Lower frequency data, e.g. monthly, are usually employed in studies that use cointegration techniques to test for long-run causal relationships (see i.a. Bermingham & O' Brien, 2011; Bumpass et al., 2015; Galeotti et al., 2003; Grasso & Manera, 2007; Honarvar, 2009; Kaufmann & Laskowski, 2005; Reilly & Witt, 1998). The research approach presented in this paper falls into the second group.

Asymmetric price adjustments can be relatively easily studied within the framework of the threshold cointegrated vector autoregressive model (TCVAR) and using the Hansen-Seo method (see Hansen & Seo, 2002). There are two reasons why this approach seems to outperform analyses based on the univariate threshold error correction models. Firstly, it is more efficient because the maximum likelihood method is applied jointly with a grid search over the threshold and the cointegrating vector. Secondly, it allows testing for the presence of a threshold cointegration effect and verifying if a two-regime TCVAR is more appropriate than a conventional linear CVAR. However, the Hansen-Seo method also has a major flaw: its assumption about deterministic variables (the constant) lying outside the cointegration space (i.e. not being present in the long-run relationships) can severely distort the estimation results. Given that, an appropriate augmentation of the method is proposed in the paper.

The empirical model focuses on the fuel price transmission along the distribution chain without analysing fundamental causes of fuel price movements (such as unit costs, market pressures, etc.). In the course of research, the following problems had to be solved. The econometric side required firstly, the development of a nonlinear method for simultaneous estimation of the TCVAR parameters and the threshold value in the presence of the constant restricted to the cointegration space; secondly, the approximation of the critical values of the *LM* statistic used in testing for a threshold cointegration by means of the fixed regressor bootstrap because of its non-standard distribution in the context of the analyzed model. On the economic side, the research sought to tackle two key problems within the fuel price-setting mechanism. Firstly, to test the hypothesis about asymmetric price transmission in the fuel market. Secondly, to determine which level of distribution

(European wholesalers, domestic wholesalers or retailers) is responsible for asymmetric price adjustments. Contrary to popular belief, oil companies do not execute an asymmetric pricing strategy despite their strong market position. The following empirical analysis proves that the cause of the 'rockets and feathers' effect in the fuel market is the pricing behaviour of petrol stations.

The multivariate approach we advocate offers far more advanced estimation and testing possibilities than a univariate analysis that has been used in the previous studies (the asymmetric ECM was applied, inter alia, by Bettendorf et al., 2003; Clerides, 2010; Galeotti et al., 2003; Oladunjoye, 2008, and the threshold autoregressive ECM by Bermingham & O' Brien, 2011; Douglas, 2010; Godby et al., 2000). The most sophisticated so far, however still univariate, ECM with the threshold cointegration has been successfully employed, for example, by Chen et al., 2005; Leszkiewicz-Kędzior & Welfe, 2014; Martin-Moreno et al., 2019.

The paper is organized as follows. Section 2 presents the derivation of a threshold cointegrated VAR model for the case of a data generating process with deterministic terms. In Section 3, an estimation method for a TCVAR with a constant restricted to the cointegration space and a threshold cointegration test are proposed. Section 4 opens with a brief description of the price transmission process in the fuel market followed by the presentation of empirical results obtained for different levels of distribution. The last section provides a summary of the main conclusions.

2. The threshold cointegrated VAR with deterministic terms

The implications of the presence of deterministic terms in the VAR process with a unit root can be explored by defining the data generating process as (the advantages of this approach are discussed in Lütkepohl, 2005, pp. 256–258):

$$v_t = y_t + Hd_t, \tag{1}$$

where $y_t = [y_{1t} \dots y_{Mt}]^T$ denotes $M \times 1$ vector of M variables integrated of order one, $d_t = [d_{1t} \dots d_{Zt}]^T$ – $Z \times 1$ vector of Z deterministic variables, $H = \begin{bmatrix} h_1 & \vdots & \dots & \vdots & h_Z \end{bmatrix}$ – $M \times Z$ matrix of parameters, $h_z = \begin{bmatrix} h_{z1} & \dots & h_{zM} \end{bmatrix}^T$, $m = 1, \dots, M$, $z = 1, \dots, Z$ $m = 1, \dots, M$. If the deterministic part consists of a constant only, that is $v_t = y_t + h_1$, the VAR model allowing for the possibility of two regimes being present can be written as:

$$\begin{bmatrix} (1 - u_{t-p}) & u_{t-p} \end{bmatrix} \begin{bmatrix} \Pi_1(L)(v_t - h_1) \\ \Pi_2(L)(v_t - h_1) \end{bmatrix} = \xi_t, \tag{2}$$

where $\xi_t = [\xi_{1t} \dots \xi_{Mt}]^T$ is $M \times 1$ vector of white-noise error terms, $u_{t-p} = \begin{cases} 1 \text{for } q_{t-p} \geq \kappa \\ 0 \text{for } q_{t-p} < \kappa \end{cases}$, $\Pi_i(L) = I - L\Pi_{1,i} - L^2\Pi_{2,i} - \dots - L^S\Pi_{S,i}$ and $i = 1, 2$ represents the regime. The threshold variable, q_{t-p}, can be defined by lagged endogenous or exogenous variables, p=1, … ,P. Multiplying the components by the appropriate lag polynomials gives:

$$(1 - u_{t-p})\Pi_1(L)v_t + u_{t-p}\Pi_2(L)v_t + \xi_t = (1 - u_{t-p})\Pi_1(L)h_1 + u_{t-p}\Pi_2(L)h_1. \tag{3}$$

The left-hand side of the above equation equals (standard CVAR is derived in Johansen,

1995):

$$\Delta v_t = (1 - u_{t-p})[\Pi_1 v_{t-1} + \sum_{s=1}^{S-1} \Gamma_{s,1} \Delta v_{t-s}] + u_{t-p}[\Pi_2 v_{t-1} + \sum_{s=1}^{S-1} \Gamma_{s,2} \Delta v_{t-s}] + \xi_t, \tag{4}$$

where Π_1 and Π_2 represent standard Π matrix in each regime i, $\Pi_i = \sum_{s=1}^{S} \Pi_{s,i} - I$, $\Gamma_{j,i} = -\sum_{s=j+1}^{S} \Pi_{s,i} = -(\Pi_{j+1,i} + \Pi_{j+2,i} + \ldots + \Pi_{S,i})$, whereas the deterministic components on the right-hand side of equation (3) are equal to:

$$\Pi_i(L)h_1 = (I - L\Pi_{1,i} - L^2\Pi_{2,i} - \ldots. - L^S\Pi_{S,i})h_1 = -\Pi_i h_1. \tag{5}$$

Provided the dimension of the cointegration space is $0 < R < M$ substituting (5) in (3) leads to the TCVAR:

$$\begin{aligned} \Delta v_t = {} & (1 - u_{t-p})A_1[B_1^T v_{t-1} + g_{1,1}] + u_{t-p}A_2[B_2^T v_{t-1} + g_{1,2}] + (1 - u_{t-p}) \sum_{s=1}^{S-1} \Gamma_{s,1} \Delta v_{t-s} + \\ & + u_{t-p} \sum_{s=1}^{S-1} \Gamma_{s,2} \Delta v_{t-s} + \xi_t \end{aligned} \tag{6}$$

for $t = S + 1, S + 2, \ldots$, where B_i, A_i are $M \times R_i$ matrices of the standard interpretation of cointegrating vectors and weights, $\Pi_i = A_i B_i^T$, $g_{1,i} = -B_i^T h_1$ or by defining the appropriate matrices:

$$\begin{aligned} \Delta v_t = {} & (1 - u_{t-p})A_1 B_1^{*T} v_{t-1}^* + u_{t-p}A_2 B_2^{*T} v_{t-1}^* + (1 - u_{t-p}) \sum_{s=1}^{S-1} \Gamma_{s,1} \Delta v_{t-s} \\ & + u_{t-p} \sum_{s=1}^{S-1} \Gamma_{s,2} \Delta v_{t-s} + \xi_t, \end{aligned} \tag{7}$$

where $G_i = g_{1,i}$, $v_{t-1}^* = \begin{bmatrix} v_{t-1} \\ \cdots \\ d_{t-1} \end{bmatrix}$ is $N \times 1$ vector of variables (in presented case $d_{t-1} = 1$), $B_i^{*T} = \begin{bmatrix} B_i^T & \vdots & G_i \end{bmatrix}$ – $R_i \times N$ matrix of the cointegrating vectors.

It follows from the above that introduction of a constant into the data generating process results in a TCVAR with deterministic variables restricted to the cointegration space. It also means that the inclusion of a dummy variable into any CVAR equation (also the TCVAR) must be reflected in the cointegration space, otherwise, the inclusion cannot be interpreted as the change of DGP. Moreover, if the stochastic component of the data generating process has the threshold VAR representation, adjustment should be made to the parameters of the cointegrating vectors, their weights, the short-run adjustment parameters, as well as to the parameters of the deterministic component.

In the special case of a unit-dimensional cointegration space, and the threshold variable being equal to the error correction term, $ect_{t-1} = \beta^{*T} v_{t-1}^*$ (see Balke & Fomby, 1997), the cointegrating vector $\beta^{*T} = \begin{bmatrix} \beta^T \vdots & G \end{bmatrix}$ is the same in both regimes and model (7) simplifies

to:

$$\Delta v_t = (1 - u_{t-1})\alpha_1 \beta^{*T} v^*_{t-1} + u_{t-1}\alpha_2 \beta^{*T} v^*_{t-1} + (1 - u_{t-1}) \sum_{s=1}^{S-1} \Gamma_{s,1} \Delta v_{t-s}$$
$$+ u_{t-1} \sum_{s=1}^{S-1} \Gamma_{s,2} \Delta v_{t-s} + \xi_t, \tag{8}$$

where $u_{t-1} = \begin{cases} 1 \text{ for } ect_{t-1} \geq \kappa \\ 0 \text{ for } ect_{t-1} < \kappa \end{cases}$, $v^*_{t-1} = \begin{bmatrix} v_{t-1} \\ \cdots \\ 1 \end{bmatrix}$, $\beta^{*T} = \begin{bmatrix} \beta^T \vdots & g_1 \end{bmatrix}$, $g_1 = -\beta^T h_1$.

The above model is based on the following assumptions: the threshold depends on the error correction term and both weights and short-run parameters vary between regimes, while the cointegrating vector is not affected by the structural change.

3. Estimation method and testing for a threshold

The intercept is directly introduced into the cointegrating vector because the DGP structure is defined as the sum of stochastic and deterministic components (see formula (1)). As a consequence, the constant term needs to be restricted to the cointegration space (see derivations presented in Section 2) because otherwise, the estimation procedure results in a biased estimator (a classical misspecification error) and frequently leads to economically unacceptable results.

To prevent such situation from happening, the following five-step algorithm is proposed that allows the constant to be present in the long-run relationship:

1. Estimate parameters of the standard CVAR with the constant restricted to the cointegration space:

$$\Delta v_t = \alpha[\beta^T v_{t-1} + g_1] + \sum_{s=1}^{S-1} \Gamma_s \Delta v_{t-s} + \xi_t, \quad \text{where } v_t = y_t + h_1. \tag{9}$$

2. Construct the interval for the constant $g_1 \in \langle g_{1L}, g_{1U} \rangle$, where $g_{1L} = \hat{g}_1 - se(\hat{g}_1)$, $g_{1U} = \hat{g}_1 + se(\hat{g}_1)$, $\hat{g}_1$ is the constant estimate from the standard CVAR, and $se(\hat{g}_1)$ – the standard error of $\hat{g}_1$.
3. For the interval from step 2 use the grid search method to estimate simultaneously the cointegrating vector and the value of the threshold that maximize the likelihood function for the CVAR given as (8).
4. Choose the value of the constant, corresponding estimates of the cointegrating vector and the threshold obtained in the previous step for which the likelihood function reaches the maximum.
5. Construct the variable u_{t-1}, estimate the remaining parameters of the TCVAR.

The grid search optimization method used in step 3 allows the use of the maximum likelihood estimator, but the standard errors and the *t*-statistics for the long-run equation cannot be calculated since the theory of inference (with its proof of estimators' consistency, distribution theory) is not provided (the same claim is in Hansen & Seo, 2002). The

parameter estimates of the short-run equation (obtained in step 5) have normal asymptotic distributions as if β^{*T} and κ were known, hence standard errors and t-statistics exist.

Testing for a threshold cointegration in TCVAR can be done with the *LM* statistic (the advantages of this approach are discussed in Hansen & Seo, 2002). The following hypotheses can be formulated:

$$H_0{:}\Delta v_t = A^* V_{t-1}(\beta^*) + \xi_t \tag{10}$$

$$H_1{:}\Delta v_t = (1 - u_{t-1}(\beta^*, \kappa))A_1^* V_{t-1}(\beta^*) + u_{t-1}(\beta^*, \kappa)A_2^* V_{t-1}(\beta^*) + \xi_t,$$

where $A^* = \begin{bmatrix} \alpha & \Gamma_1 & \Gamma_2 & \dots & \Gamma_{S-1} \end{bmatrix}$, $A_i^* = \begin{bmatrix} \alpha_i & \Gamma_{1,i} & \Gamma_{2,i} & \dots & \Gamma_{S-1,i} \end{bmatrix}$, $V_{t-1}(\beta^*) = \begin{bmatrix} \beta^{*T} v_{t-1}^* \\ \Delta v_{t-1} \\ \Delta v_{t-2} \\ \vdots \\ \Delta v_{t-S+1} \end{bmatrix}$, $u_{t-1}(\beta^*, \kappa) = \begin{cases} 1 \text{ for } \beta^{*T} v_{t-1}^* \geq \kappa \\ 0 \text{ for } \beta^{*T} v_{t-1}^* < \kappa \end{cases}$ and (β^*, κ) are known and fixed.

It should be noted that in both hypotheses, H_0 and H_1, the constant is restricted to the cointegration space (contrary to Hansen & Seo, 2002). What is more, under the alternative hypothesis both the adjustment coefficient as well as short-run dynamic parameters vary between regimes, so the decomposition of identified asymmetry into long-run and short-run effects is not feasible.

The heteroscedasticity-robust *LM* statistic takes the form:

$$\begin{aligned} LM(\beta^*, \kappa) = {} & vec(\hat{A}_1^*(\beta^*, \kappa) - \hat{A}_2^*(\beta^*, \kappa))^T (\hat{D}^2_{\hat{A}_1^*}(\beta^*, \kappa) + \hat{D}^2_{\hat{A}_2^*}(\beta^*, \kappa))^{-1} \\ & \times vec(\hat{A}_1^*(\beta^*, \kappa) - \hat{A}_2^*(\beta^*, \kappa)), \end{aligned} \tag{11}$$

where $\hat{D}^2_{\hat{A}_1^*}(\beta^*, \kappa)$ and $\hat{D}^2_{\hat{A}_2^*}(\beta^*, \kappa)$ are the Eicker-White covariance matrix estimators for $vec(\hat{A}_1^*(\beta^*, \kappa))$ and $vec(\hat{A}_2^*(\beta^*, \kappa))$, respectively.

The asymptotic distribution of *LM* statistic (11) for a case of known cointegrating vector has been presented by Hansen and Seo (2002). However, proving that the use of the cointegrating vector estimate instead of its true value does not alter the asymptotic distribution of *LM* statistic is not possible (see Hansen & Seo, 2002) because the model is non-linear and vector v_{t-1}^* consists of non-stationary variables. As a result, the critical values are unavailable and the fixed regressor bootstrap method is proposed as an adequate solution.

Since β^*, κ are unknown, the above *LM* test statistic is evaluated at point estimates obtained under H_0. The null estimate of β^* can be calculated from (9), however, because of the lack of κ estimate under H_0, the following modification must be applied (see Davies, 1987; Hansen & Seo, 2002):

$$SupLM = \sup_{\kappa_L \leq \kappa \leq \kappa_U} LM(\beta^*, \kappa), \tag{12}$$

where κ_L and κ_U are defined as κ_0 and $(1 - \kappa_0)$-percentile of $\beta^{*T} v_{t-1}^*$, respectively, κ_0 is the trimming parameter. Maximization is achieved by a grid search over $\langle \kappa_L, \kappa_U \rangle$ and the distribution of *SupLM* can be approximated by the fixed regressor bootstrap or the residual bootstrap. The first method was used in the empirical analysis, because the simulation results presented by Hansen and Seo (2002) confirm that a greater power of *SupLM* is

guaranteed then. It assumes that $e\hat{c}t_{t-1} = \hat{\beta}^{*T} v^*_{t-1}$, $\hat{V}_{t-1} = V_{t-1}(\hat{\beta}^*)$ and $e_t = \Delta v_t - \hat{A}^* V_{t-1}(\hat{\beta}^*)$, i.e. residuals from the standard CVAR model, are held fixed at their sample values. The following steps lead to a single draw from the approximated distribution:

1. Set $v_{bt} = e_t r_{bt}$, where r_{bt} is iid $N(0, 1)$, see Hansen and Seo (2002).
2. Regress v_{bt} on $\hat{V}_{t-1}$ yielding residuals e_{bt}.
3. Regress v_{bt} on $(1 - u_{t-1}(\hat{\beta}^*, \kappa))\hat{V}_{t-1}$ and $u_{t-1}(\hat{\beta}^*, \kappa)\hat{V}_{t-1}$ yielding estimates $\hat{A}^*_1(\kappa)_b$ and $\hat{A}^*_2(\kappa)_b$, as well as residuals $e_{bt}(\kappa)$.
4. Define $\hat{D}^2_{\hat{A}^*_1}(\kappa)_b$ and $\hat{D}^2_{\hat{A}^*_2}(\kappa)_b$ as in (11), setting $\beta^* = \hat{\beta}^*$ and replacing e_t with e_{bt}.
5. Consequently, the *LM* statistic for a single draw is given by:

$$SupLM^* = \sup_{\kappa_L \le \kappa \le \kappa_U} vec(\hat{A}^*_1(\kappa)_b - \hat{A}^*_2(\kappa)_b)^T (\hat{D}^2_{\hat{A}^*_1}(\kappa)_b + \hat{D}^2_{\hat{A}^*_2}(\kappa)_b)^{-1} \times vec(\hat{A}^*_1(\kappa)_b - \hat{A}^*_2(\kappa)_b). \tag{13}$$

The distribution of $SupLM^*$ yields a valid approximation to the asymptotic null distribution of *SupLM* (see Hansen, 1996). The distribution of $SupLM^*$ is unknown, but it can be simulated. With independent draws of r_{bt}, the procedure should be repeated (in our study 500 times) and *p*-value calculated as a percentage of simulated $SupLM^*$ values greater than the *SupLM*.

Because the fixed regressor bootstrap is based on the sample values, in the case of any empirical model the appropriate *p*-values must be simulated individually.

4. Empirical analysis

The behaviour of the prices of two types of fuels with the largest shares in the Polish market (unleaded 95 petrol and diesel) was analysed by using a TCVAR and the foregoing estimation procedure. All monthly time series span the period from January 2000 to July 2016 and are given as indexes expressed in Polish zlotys (to eliminate the effects of the exchange rate fluctuations) and transformed into real terms (adjusted for inflation) by means of the consumer price index (2000.01 = 1.00). They were sourced from the BM Reflex, the Thomson Reuters and the Central Statistical Office of Poland. The results of the unit root tests clearly showed that all variables were integrated of order one (see Table 1 in Appendix 1).

The formation of fuel prices is a three-stage process. Firstly, the Brent crude oil price (a reference price in Europe) is transmitted into the prices of oil-based products in the European wholesale market (the ARA market). Secondly, the domestic wholesale prices are determined by the fuel spot prices (quoted daily at the ARA) converted into domestic currency (zloty, PLN) and enlarged by the taxes (excise tax, fuel fee and inventory fee) and mark-ups. Lastly, the domestic wholesale prices augmented by the value added tax (calculated ad valorem) and the retailer's mark-ups become the fuel prices to be paid by the final customers. This mechanism can be formalized by the following three long-run relationships:

$$ara^j_t = \beta^{ara}_0 + \beta^{ara}_1 brent_t + \varepsilon^{ara}_t, \tag{14a}$$

$$wp^j_t = \beta^{wp}_0 + \beta^{wp}_1 ara^j_t + \beta^{wp}_2 tax^j_t + \varepsilon^{wp}_t, \tag{14b}$$

$$rp^j_t = \beta^{rp}_0 + \beta^{rp}_1 wp^j_t + \varepsilon^{rp}_t, \tag{14c}$$

where *brent* denotes the Brent oil price, *ara* – the ARA price, *wp* – the domestic wholesale price, *tax* – excise tax, fuel fee and inventory fee (summed up), *rp* – the net retail price of fuel; *j* stands for the type of fuel (unleaded 95 petrol or diesel). All variables are in logarithms. Since VAT is calculated ad valorem, it does not enter (14c) and the net retail prices can be used instead.

The presented model assumes time-constant average mark-ups, which are represented by the relevant parameters. In consequence, the dynamic nature of mark-ups is not captured. The main reason for adopting this solution was the limited length of time series and data availability. However, the solution takes account of the diversification of mark-ups depending on the direction of fuel price changes (upward or downward).

To test separately for asymmetric adjustments to the long-run paths at each of three stages of commodity trading (14a)-(14c), three TCVAR models defined respectively by vectors $\left[\, ara_t^j \quad brent_t \,\right]$, $\left[\, wp_t^j \quad ara_t^j \quad tax_t^j \,\right]$ and $\left[\, rp_t^j \quad wp_t^j \,\right]$ were built. A system of four-variable CVAR with 3 cointegrating relations would likely afford more sophisticated hypothesis testing, but at the current stage of the econometric theory of nonlinear processes, the approach is not feasible.

The lag order for each of the three TCVAR models was set to 1 (according to the AIC and BIC information criteria), while the trimming parameter was chosen at the level of 0.1. Because tax_t^j proved to be weakly exogenous, each system can possibly contain at most one cointegrating vector.

The estimation results for all three levels of the fuel price-setting process are presented in the context of the first equation (each system consists of two equations) because, even though both endogenous variables react to the error correction term and adjust to the long-run trajectory, only the first equation can be given a full economic interpretation. The estimates for the first level of distribution of unleaded 95 petrol are as follows (*t*-Student statistics are under parameters in the brackets):

$$\Delta ara_t^{PB95} = \begin{cases} -\underset{(-3.628)}{0.631}\,(ara_{t-1}^{PB95} - 0.871 brent_{t-1} - 0.077) + \text{short-run terms}, ect_{t-1} \leq -0.081, \\ -\underset{(-2.401)}{0.384}\,(ara_{t-1}^{PB95} - 0.871 brent_{t-1} - 0.077) + \text{short-run terms}, ect_{t-1} > -0.081. \end{cases} \tag{15}$$

The shares of observations in the two regimes are 0.254 and 0.746, respectively, $SupLM = 13.901$ ($p - \text{value} = 0.283$).

For the second level of distribution we got:

$$\Delta wp_t^{PB95} = \begin{cases} -\underset{(0.658)}{0.275}\,(wp_{t-1}^{PB95} - 0.408 ara_{t-1}^{PB95} - 0.580 tax_{t-1}^{PB95} + 0.001) + \\ +\text{short-run terms}, ect_{t-1} \leq 0.003, \\ -\underset{(-2.566)}{0.254}\,(wp_{t-1}^{PB95} - 0.408 ara_{t-1}^{PB95} - 0.580 tax_{t-1}^{PB95} + 0.001) + \\ +\text{short-run terms}, ect_{t-1} > 0.003. \end{cases} \tag{16}$$

The shares of observations in the two regimes are 0.406 and 0.594, respectively, $SupLM = 12.829$ ($p - value = 0.348$).

In neither model was the difference between the long-run adjustment parameters in the two regimes statistically significant (as proved by the *SupLM* test), so asymmetry was not detected. This means that, firstly, crude oil price changes bring about a symmetric

response from fuel prices quoted in the European wholesale market. This result is in line with the efficient market hypothesis according to which it is not possible to have extra profits over the long-run. Secondly, the wholesale petrol price goes back to the equilibrium trajectory with the same speed regardless of whether the ARA wholesale price goes up or down. This finding confirms that even though Polish oil companies operate within the oligopolistic market they do not exploit their market position to generate extra profits. More than that, fearing possible imports they align their fuel prices with the European prices (converted into zlotys).

On the contrary, the asymmetric price transmission manifests itself at the third level of the distribution chain where the domestic wholesale price of unleaded 95 petrol is transmitted to the retail price, as shown by the following results:

$$\Delta rp_t^{PB95} = \begin{cases} -\underset{(-3.334)}{0.935}\,(rp_{t-1}^{PB95} - 0.817wp_{t-1}^{PB95} + 0.015) + \text{short-run terms}, ect_{t-1} \leq -0.039, \\ -\underset{(-5.438)}{0.462}\,(rp_{t-1}^{PB95} - 0.817wp_{t-1}^{PB95} + 0.015) + \text{short-run terms}, ect_{t-1} > -0.039. \end{cases} \tag{17}$$

The shares of observations in the two regimes are 0.102 and 0.898, respectively, $SupLM = 44.169$ ($p - \text{value} = 0.000$).

The difference in the loading coefficients in both regimes is statistically significant what means that if the retail fuel price exceeds the equilibrium level or falls below it (but not more than 3.9%), then the speed of adjustment to the long-run trajectory will be two times slower comparing with the deviation beyond the threshold value. This means that trying to compensate for wholesale price increases that exceed the trigger level, retailers rapidly raise their prices, but their response to wholesale price decreases is significantly weaker. Low price elasticity of the demand for fuels is probably the reason why the asymmetric pricing policy can be adopted at this distribution level (similar results are reported by Leszkiewicz-Kędzior & Welfe, 2014).

The main conclusions remain unchanged in the case of diesel. The results for the first two levels of distribution confirm the symmetry of the price-setting behaviour while statistically significant asymmetric price transmission was identified at the retail level:

$$\Delta rp_t^{ON} = \begin{cases} -\underset{(-3.460)}{0.846}\,(rp_{t-1}^{ON} - 0.923wp_{t-1}^{ON} - 0.014) + \text{short-run terms}, ect_{t-1} \leq -0.024, \\ -\underset{(-2.095)}{0.251}\,(rp_{t-1}^{ON} - 0.923wp_{t-1}^{ON} - 0.014) + \text{short-run terms}, ect_{t-1} > -0.024. \end{cases} \tag{18}$$

The shares of observations in the two regimes are 0.147 and 0.853, respectively, $SupLM = 47.173$ ($p - \text{value} = 0.000$).

From the results it follows that if the diesel price falls below the equilibrium level by more than 2.4%, then retailers raise it over three times faster than if its deviation from the equilibrium exceeds −2.4%. This implies that retailers pursue an asymmetric pricing policy and respond to increases in wholesale diesel prices much more strongly than to the falls.

To make the study complete, the hypotheses that prices are fully transmitted between individual distribution levels were also verified. Understanding the reasons for full or incomplete pass-through allows economists to better assess the pricing strategies of

market entities and produce more accurate forecasts of fuel price variations. The cointegrating parameters estimates obtained for the third level of the distribution chain ($\beta_1^{rp} = 0.817$ for unleaded 95 petrol and $\beta_1^{rp} = 0.926$ for diesel) demonstrated that the pass-through of the domestic wholesale prices to the retail fuel prices was incomplete in the sample. There are probably two key reasons for this. Firstly, to remain competitive, retailers kept reducing their fuel price margins while increasing the sales of high-margin products such as food, snacks, alcohol, car accessories and setting up their own fast food chains (e.g. Wild Bean Café, Stop Café, Deli2go) to compensate for lower profits on the core business. Secondly, the downward pressure on the refinery products prices was partly due to the grey market phenomenon affecting the Polish fuel industry in the sample period. Its consequence was a rising number of fuel traders who circumvented all applicable legal and fiscal requirements and regulations (taxes and fees, mandatory reserves, the national biofuels policy) and so they could offer prices below the market rates. It should be added, that after the Polish government introduced new legislation (the so-called fuel package) in 2016, much lower estimates of illegally sold fuels are reported in Poland. It should be noted that while the grey market effect on fuel prices is specific to Poland, the range of products available at petrol stations is expanding in all European countries. Because data on the volume of non-fuel products sold at petrol stations and on the amount of illicit fuel entering the legitimate market are not available, neither of the factors could be included in the relation (14c), which led to the non-unit estimates of β_1^{rp} parameters.

As the wholesale price adjustments in the European market are symmetric, the standard cointegrated VAR model was applied to empirically verify the hypothesis about the full pass-through of crude oil price changes into the ARA fuel prices in the long run:

$$\Delta ara_t^{PB95} = -\underset{(-4.337)}{0.444}\left(ara_{t-1}^{PB95} - \underset{(-24.804)}{0.902}\, brent_{t-1} - \underset{(-5.767)}{0.071}\right) + \text{short-run terms.} \tag{19}$$

The restriction $\beta_1^{ara} = 1$ was rejected by the data ($LR = 5.815$, $p - \text{value} = 0.016$) which supports the hypothesis of the incomplete price transmission. The refining margins in the petroleum industry are determined by the technology utilized by the least efficient refiner that meets the fuel demand. Therefore, the wholesale fuel market is under continuous pressure of technological progress, because to remain profitable, oil companies tend to produce at the lowest cost possible. Not only has this substantially improved the crude oil processing technologies in Europe in the past two decades, but it also explains why the transmission of the crude oil price into the European wholesale prices of oil-based products is also most probably incomplete.

The results for diesel are virtually the same:

$$\Delta ara_t^{ON} = -\underset{(-3.765)}{0.320}\left(ara_{t-1}^{ON} - \underset{(-28.649)}{0.912}\, brent_{t-1} - \underset{(-1.451)}{0.022}\right) + \text{short-run terms.} \tag{20}$$

Also in this case, the hypothesis about the complete transmission of crude oil prices into the ARA diesel price was rejected ($LR = 4.800$, $p - \text{value} = 0.028$).

The domestic wholesale market is the only level of distribution for which the pass-through transmission proved to be complete. The homogeneity restriction given as

$\beta_1^{wp} + \beta_2^{wp} = 1$ was not rejected for either unleaded 95 petrol:

$$\Delta wp_t^{PB95} = -\underset{(-2.398)}{0.299}\left(wp_{t-1}^{PB95} - \underset{(-24.896)}{0.403}\, ara_{t-1}^{PB95} - \underset{(-36.957)}{0.597}\, tax_{t-1}^{PB95} + \underset{(0.388)}{0.002}\right) + \text{short-run terms}, \tag{21}$$

$$LR = 2.115\ (p - \text{value} = 0.146)$$

or diesel:

$$\Delta wp_t^{ON} = -\underset{(-3.293)}{0.334}\left(wp_{t-1}^{ON} - \underset{(-13.736)}{0.449}\, ara_{t-1}^{ON} - \underset{(-16.839)}{0.551}\, tax_{t-1}^{ON} + \underset{(1.930)}{0.018}\right) + \text{short-run terms}, \tag{22}$$

$$LR = 2.456\ (p - \text{value} = 0.117)$$

meaning that both determinants (fuel prices in the ARA market and taxes) are fully transmitted into domestic wholesale prices.

From the above it follows that the fuel price-setting process is quite complex and that the price transmission mechanisms are specific to trading levels represented by the European wholesale market, the domestic wholesale market and the domestic retail market. This means that they should be considered individually. Unlike studies that concentrate on the direct relationship between the crude oil price and the retail prices of fuels and consequently fail to identify entities which contribute to the asymmetric price adjustments, our analysis has clearly demonstrated that the 'rockets and feathers' effect in the Polish fuel market is caused by the petrol stations' pricing policy.

Similar study on the Polish fuel market have been performed so far by Leszkiewicz-Kędzior and Welfe (2014), who have reached analogous conclusions for diesel, but somewhat different for unleaded 95 petrol (identifying weak asymmetry at the ARA market). However, current analysis is based on a longer time series and utilizes more advanced econometric method, so should be regarded as more reliable.

5. Conclusions

The empirical estimates show that fuel price adjustments at the first two levels of the distribution chain (the European wholesale market and the domestic wholesale market) are symmetric, meaning that neither European oil companies nor domestic refiners take advantage of the crude oil price volatility to earn extra profits at the cost of intermediate and final customers. Only the transmission of the wholesale fuel price to the retail prices is asymmetric, indicating that the petrol stations' pricing policy leads to the so-called 'rockets and feathers' effect in the fuel market. The main findings remain the same for both types of fuels considered.

The result for the European wholesale level is general and common for all European countries, and the validity of the conclusion on price adjustments at the domestic wholesale level depends on the organization of the wholesale market in a country. Given the structural similarity of the wholesale markets for oil-based products in the EU member states, this finding can possibly apply to other countries or be at a least working hypothesis. The result for the last level of distribution may, in turn, differ between EU countries depending

on the market behaviour of local retailers. In countries where the petroleum product prices tend to respond asymmetrically to changes in the crude oil prices, the responsibility for this phenomenon usually lies with retailers and their pricing policies. This implies that they coordinate their actions to some extent which restrains effective price competition.

In order to demonstrate asymmetry in fuel price adjustments, the threshold cointegrated VAR model needs to be derived. The paper provides evidence that deterministic variables must be restricted to the cointegration space if the data generating process consists of stochastic and deterministic parts. Accordingly, the estimation method proposed augments the Hansen-Seo procedure (Hansen & Seo, 2002) by allowing the constant to be present in the cointegrating vector. This modification is successful in providing undistorted and thus interpretable results and it can be applied to any economic problem reducible to a bivariate system. A methodology for estimating a threshold cointegrated VAR in higher dimensions is not yet available, and it can be a subject of future research. The introduction of structural breaks in the long-run relationships, too, can set the direction of future research and entail further modifications of the method.

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Notes on contributors

Emilia Gosińska PhD in Economics, works as an Assistant Professor at the Chair of Econometric Models and Forecasts at the University of Lodz. Her main research interests are related to cointegrated VAR models with structural breaks, threshold cointegration and threshold cointegrated VAR models.

Katarzyna Leszkiewicz-Kędzior PhD in Economics, Assistant Professor at the Chair of Econometric Models and Forecasts at the Faculty of Economics and Sociology of the University of Lodz. She has published research papers in the fields of macromodelling, including modelling of nonlinear economic processes, as well as standard and threshold cointegration.

Aleksander Welfe Professor of Economics, member of the Polish Academy of Sciences and of the Academie Europeenne des Sciences, des Arts et des Lettres. Macroeconomist and econometrician, the author of several macromodels of the Polish economy which were regularly used for forecasting and simulation analyses. He published in the journals in the fields of econometric theory, applied econometrics, macromodelling, time series analysis and cointegration. He is the Associate Editor of

Economic Modelling journal and serves as Chief Editor for *Central European Journal of Economic Modelling and Econometrics*.

ORCID

Emilia Gosińska http://orcid.org/0000-0002-5325-6144
Katarzyna Leszkiewicz-Kędzior http://orcid.org/0000-0003-4085-7271
Aleksander Welfe http://orcid.org/0000-0002-9925-3432

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Appendix

Table 1. Inference on the order of integration

Variable	Hypotheses	ADF-GLS test statistic	KPSS test statistic	ADF-GLS test conclusion	KPSS test conclusion
brent	ADF-GLS	−1.522*	0.393*	I(1)	I(1)
ara^{PB95}	I(1) vs. I(0)	−1.589*	0.317*	I(1)	I(1)
ara^{ON}	KPSS	−1.826*	0.391*	I(1)	I(1)
wp^{PB95}	I(0) vs. I(1)	−0.921*	0.329*	I(1)	I(0)
wp^{ON}		−2.292*	0.358*	I(1)	I(1)
tax^{PB95}		−2.023*	0.349*	I(1)	I(0)
tax^{ON}		−2.140*	0.338*	I(1)	I(1)
rp^{PB95}		−1.242*	0.361*	I(1)	I(1)
rp^{ON}		−1.931*	0.403*	I(1)	I(1)
brent	ADF-GLS	−3.007	0.111	I(1)	I(1)
ara^{PB95}	I(2) vs. I(1)	−2.451	0.086	I(1)	I(1)
ara^{ON}	KPSS	−2.220	0.127	I(1)	I(1)
wp^{PB95}	I(1) vs. I(2)	−2.482	0.079	I(1)	I(1)
wp^{ON}		−2.215	0.113	I(1)	I(1)
tax^{PB95}		−2.164	0.103	I(1)	I(1)
tax^{ON}		−2.247	0.089	I(1)	I(1)
rp^{PB95}		−8.567	0.080	I(1)	I(1)
rp^{ON}		−2.601	0.148	I(1)	I(1)

Notes: Inference performed at a 0.05 level of significance.

* means that the test equations included constant and trend, in the remaining equations only the constant was included. The Schwert criterion was used to determine the maximum lag length.