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PRICE REGULATIONS AND PRICE ADJUSTMENT DYNAMICS: EVIDENCE FROM THE AUSTRAIN RETAIL FUEL MARKET

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Abstract

After controversial public debates, fuel price regulations were implemented in Austria prohibiting fuel retailers from raising their prices more than once per day. This paper investigates whether these policy measures affected the price transmission dynamics from crude oil prices to retail fuel prices. We estimate different specifications of nonlinear error correction models to quantify a potentially asymmetric adjustment behaviour and compare the results over three subsamples. Particularly, we estimate our models for a pre-regulation period, a between-regulations and a post-regulation period. At first glance, we obtain conflicting results on the efficacy of this policy measure. While the adjustment to the long-run equilibrium seems to be faster if crude oil prices are relatively low, transitory crude oil price decreases are passed through faster than price increases. Only if we consider the combined effect of a crude oil price shock, we can reveal that crude oil price changes are generally passed through faster in the post-regulation period. Further, we find that crude oil price decreases are now passed through slightly faster than crude oil price increases. Hence, we conclude that the Austrian fuel price regulation seems to have fostered competition between fuel retailers.

Keywords: Asymmetric price transmission, price regulation, nonlinear error correction model, retail fuel prices, crude oil prices
JEL Classification: C22, D40, Q41

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1 Introduction

Road transportation is the predominant transportation mode in OECD countries and relies almost completely on oil. This leaves no doubt that fuel prices are of particular interest for consumers, regulatory authorities and policy-makers in general. Regulatory authorities are concerned with the question whether large oil companies exercise their market power to charge consumers with higher prices than necessary. One particular issue that is discussed in the public and political debate is the presumption that oil companies delay input price decreases but pass through input price increases to the retail market immediately. This behaviour is generally known as asymmetric price transmission (APT). From the perspective of standard economic theory such asymmetric price transmissions lead to consumer welfare losses which should be avoided. Consequently, asymmetric fuel pricing has attracted a lot of attention in the economic literature. Several empirical studies have been conducted in the last three decades to evaluate the fuel price transmission in different countries with mixed results (see, for example, Bacon (1991); Manning (1991); Kirchgässner and Kübler (1992); Galeotti et al. (2003); Grasso and Manera (2007); Balaguer and Ripollés (2012); Asane-Otoo and Schneider (2015); Schweikert (2017)). A recent meta-analysis by Perdiguero (2013) found that, among others, the segment of the industry analyzed, different research designs and the time span of the analysis might explain the heterogeneity in results. Also, the level of competition seems to be a key factor for the existence of asymmetries in a specific fuel market.

In this paper, we focus on vertical price transmission along the fuel distribution chain. Vertical APT can be classified into short-run APT and long-run APT according to Meyer and Cramon-Taubadel (2004). Short-run APT describes the asymmetric effects of positive or negative input price changes on output prices. Hence, the pass through depends on the sign and magnitude of the price change. In contrast, long-run APT evaluates reaction times, length of fluctuations and the speed of adjustment towards a long-run equilibrium between input and output prices. Consequently, long-run APT leads to a temporary transfer until the prices have adjusted to a new equilibrium state. Short-run APT, on the other hand, leads to a permanent transfer considering that prices are integrated and shocks have a persistent effect. A combination of short-run and long-run APT is possible and in empirical cases it usually cannot be determined a priori which type of asymmetry is stronger. Further, we distinguish between positive and negative APT. Positive (negative) asymmetry implies that output prices tend to respond faster to an increase (a decrease) in input prices than to a decrease (an increase). However, these terms have an opposite normative interpretation. From the perspective of the consumer, positive asymmetry constitutes a welfare loss while negative asymmetry constitutes a welfare gain. Positive APT has also been referred to as the 'rockets and feathers' phenomenon since it implies that prices rise like rockets but fall like feathers.[1]

A considerable amount of literature has emerged raising explanations for the existence of asymmetric pricing behaviour (see, among others, Johnson (2002), Verlinda (2008), Deltas (2008) and Lewis and Noel (2011)). One of the primary causes could be the exploitation of market power by large oil companies in non-competitive market structures. Fuel markets are often highly concentrated and show different features of oligopolistic markets. Balke et al. (1998) consider oligopolistic firms that engage in tacit collusion. If the input prices increase in such a market, the colluding firms will adjust their prices upwards in order to signal their competitors that they stick to their collusive agreements. In contrast, if input prices decrease, firms will adjust their prices slowly to avoid signalling to their competitors that they are not retaining the tacit collusion equilibrium. The old retail price also offers a natural focal point for oligopolistic sellers so that high prices can be maintained in the short run. Following the tacit collusion or focal point hypotheses, we would expect to find evidence for positive APT in our retail fuel price data.

Another cause for asymmetric pricing is known as the search cost hypothesis (Johnson, 2002). It implies that consumers have a stronger incentive to search for lower prices during periods of increasing prices while their incentive to search is inhibited in periods of falling prices. When wholesale prices increase and the retail filling station decides to adjust its prices, consumers are motivated to search for a better price. In contrast, falling wholesale prices lead to a lower search intensity of consumers. Consequently, retailers have an incentive to delay input price increases to maintain their market share. Following the search cost hypothesis, we would expect to find evidence for negative APT in our retail fuel price data. Also, retailers generally pass through price decreases faster in markets where search costs are low. Menu costs (Ball and Mankiw, 1994) and inventory adjustment (Borenstein and Shepard, 2002) are added as further explanations for positive APT.

The existence of asymmetric pricing in a fuel market has substantial consumer

[1]The name originates from the Bacon (1991) paper entitled: 'Rockets and feathers: the asymmetric speed of adjustment of UK retail gasoline prices to cost changes'

welfare implications since under positive APT consumers do not benefit from price reductions in the same extent as they would under symmetric adjustment. This welfare misallocation could be resolved by appropriate policy measures. Several regulatory measures to fight collusive behaviour have been implemented, for example, in parts of Australia, Belgium, Luxembourg and parts of Canada. Recently, a fuel price regulation was established in Austria which prohibited retail filling stations to increase prices more than once per day. The measure is expected to increase price transparency for consumers by limiting the number of price changes per day and intends to foster competition. However, the real effects of the proposed policy measures are unclear a priori and have to be evaluated empirically.

This paper investigates the price transmission from crude oil to retail fuel prices in Austria using nonlinear error correction models. In particular, we investigate the pass through of Brent spot price changes to retail gasoline and diesel prices excluding tax and duty. For that matter, we use daily observations obtained for a sample ranging from August 2004 to March 2016. The main focus of this study lies in the evaluation of two fuel price regulations introduced in July 2009 and January 2011 in Austria and their effects on the price adjustment dynamics. While the price regulation in July 2009 limited the retailers to one price increase per day at a time depending on the type of retail filling station, the second price regulation standardized the procedure and restricted all retailers to increase their prices at noon. However, retailers were still allowed to decrease prices at any time.

The effects of this policy on the fuel price transmission might be twofold. On the one hand, fixing a date for price increases could have brought greater price transparency which might have lowered the search costs for consumers. Consequently, it would be easier for consumers to select the retail filling station with the best prices and in turn it would be more difficult for oil companies to delay price decreases. On the other hand, it could be easier for oil companies to maintain a tacit collusion equilibrium since they would only have to coordinate a daily maximum price and the subsequent (potential) price decreases. Moreover, it can be expected that firms anticipate that they will not have the possibility to increase their prices and might charge higher initial prices (Obradovits, 2014). The retailers might lower this initial price level in smaller steps throughout the day which would in turn appear as positive short-run APT. We evaluate the efficacy of the Austrian fuel price regulation by comparing the findings on short-run and long-run asymmetry for subsamples comprising of observations before and after the introduction of both regulations.

Our main contributions to the discussion about asymmetric price transmission are the following: First, whereas the majority of studies on asymmetric pricing transmission provide empirical evidence for one or more countries over a specific period of time, we investigate whether the fuel price adjustment dynamics have changed after the introduction of a specific policy measure regulating the retail fuel market. This could be relevant for policy makers intending to implement similar policy measures.[2] Second, to the best of our knowledge, this paper is the first study to analyze the Austrian retail fuel market in terms of asymmetric pricing behaviour using data obtained after the second fuel price regulation was implemented in 2011. The Austrian federal competition authority (BWB) conducted studies in order to test for asymmetries in the Austrian fuel market with mixed results using data until 2012 but did not evaluate a sufficiently long period to study any long-term effects. Third, we use daily data instead of weekly or monthly data to account for the generally fast adjustment of fuel prices to crude oil price changes which has been reported in previous studies (see, for example, Bachmeier and Griffin (2003), Balaguer and Ripollés (2012) and Perdiguero (2013)). Thereby, we can reveal asymmetries which would be invisible in aggregated data.

The remaining part of the paper is structured as follows: Section 2 compares the Austrian fuel price regulation to other regulatory measures found in Australia and Canada. Section 3 contains a review of relevant empirical literature and discusses the economic implications of price regulation policies. Section 4 outlines the methodology used to model potentially asymmetric price transmissions. In Section 5, we apply these techniques to Austrian fuel price data and Section 6 offers a conclusion.

2 A comparison of different fuel price regulations

According to the IEA (2014), the Austrian fuel market is one of the most regulated in the European Union. The current price regulations were implemented in two steps. The first regulation, Austrian Fuel Price Fixing Act ('Spritpreisverordnung') introduced in 2009, restricted retail filling stations to increase prices only once per day. The

[2]The discussion on fuel price regulations is prevalent in many European countries, e.g. Germany (*Die Welt*: 'Regulierung macht das Tanken nur noch teurer' on March 03, 2012) and the UK (*Bloomberg*: 'Price caps are the wrong solution for UK energy' on October 18, 2017). While the competition authority in Germany (Bundeskartellamt) has advised against Austrian-type fuel price regulations (Bundeskartellamt, 2011), the issue is still regularly brought up in public debates. In 2013, the Market Transparency Unit for fuels was established to provide consumers with information about current fuel prices in Germany.

designated time for price increases in case of 24-hour stations was midnight, while self-service stations were allowed to increase prices at 8:30 and stations with regular opening hours could increase prices at the opening hour. Disregarding the regulation led to a 2,000 Euro fine. Price decreases were allowed at any time and without limit. In 2011, the regulation was tightened and all retail filling stations in Austria were only allowed to increase their prices at noon (BWB 2009). Additionally, in July 2011 a law on transparency of prices ('Preistransparenzverordnung Treibstoffpreise 2011') was approved committing all retail filling stations to send their price changes within 30 minutes to E-Control, the Austrian Energy Regulator, which makes the prices available to drivers via a free internet tool (OECD, 2013).

Similar fuel price regulations have been implemented in different countries during the last decade. Starting in 2001[3], Western Australia imposed a regulatory measure restricting the fuel price for 24 hour intervals. Retailers have to submit the next day's fuel prices by 14:00 to the state government and are obliged to follow the notified price from 6:00 for the next 24 hours. The major difference to the Austrian-type regulation is that neither increases nor decreases are allowed in Western Australia. The fuel prices of each day and the following day are posted on the official FuelWatch web page. This form of fuel price regulation reduces daily price volatility by construction but the regulatory authorities intent to provide transparency and reduce consumer search costs as well (Byrne, 2014).

Most Canadian provinces and territories have regulated and deregulated their fuel markets at some point in time. Currently, five of thirteen provinces and territories are regulating their fuel prices. Nova Scotia and New Brunswick enacted the Petroleum Product Pricing Act in July 2006 defining so-called price ceilings or caps, but this policy measure differs between provinces. In Novia Scotia, it sets minimum and maximum prices for each week (before 2009 the prices were set for the following two weeks). The designated fuel prices at the pump are determined from the benchmark spot fuel price (New York Harbor Spot), applicable taxes, wholesale and retail margins and transportation cost. Wholesale prices are set 6 cents per litre above the benchmark price, the markup for transportation cost varies between 0.2 - 2.0 cents per litre and the profit margin is determined to be between 4.0 - 5.5 cents per litre. In New Brunswick, only the maximum price is set by the regulatory authority (Suvankulov et al., 2012). Luxembourg maintains a similar price ceiling mechanism.

[3]Petroleum Products Pricing Regulations 2000 extending the Petroleum Products Pricing Act 1983.

Several studies are concerned with the economic implications of fuel price regulations: Berninghaus et al. (2012) use a game theoretical lab experiment to test the hypothesis that companies would set higher prices under conditions defined by the Austrian-type fuel price regulation. Their results show that prices in the regulated market are indeed at a higher level than in the non-regulated market. In particular 76% of the participants chose the highest prices in the non-regulated market whereas 91% chose the highest prices in the regulated market. Consequently, the hypothesis of higher prices under the Austrian fuel price regulation seems to hold in this specific lab experiment. The authors also measure the price volatility in both synthetic markets which turns out to be lower in the regulated market.

Obradovits (2014) conducts a theoretical analysis in a two-period duopoly model with consumer search and finds that the Austrian-type fuel price regulation has detrimental effects on consumer welfare. Retailers intertemporally distort prices in a way so that their profits remain unchanged compared to the unregulated market. This behaviour is in line with the results obtained by Berninghaus et al. (2012). A delay of purchases would be beneficial for consumers, but most consumers are not flexible enough in their purchasing time and experience a welfare loss.

In a further lab experiment, Haucap and Müller (2012) investigate the effects of three different fuel price regulations (Luxembourg, Western Australia, Austria). The results suggest that the Austrian-type and the Luxembourg-type regulation decrease consumer welfare while the Western Australian regulation does not. In addition, none of them lowers retail prices.

3 Literature Review

Although an extensive empirical literature on asymmetric fuel pricing exists, only a few studies can be found that analyze asymmetric behaviour in connection with fuel price regulations. This can in parts be explained by the small number of developed countries with regulated fuel markets. Therefore, the focus in this section is on studies related to fuel price regulations in the countries mentioned above and on studies specifically investigating asymmetric pricing in the Austrian retail fuel market.

The BWB conducts studies on the Austrian fuel market since 2004, motivated by a public concern that asymmetric pricing could harm consumer welfare (OECD, 2013). They conducted two studies in 2008 and 2010 in order to estimate a possible asymmetric behaviour in response to wholesale prices using an EGARCH error correction model. Daily data were obtained from the Austrian Automobile, Motorcycle and Touring Club (ÖAMTC) for the period September 2003 to March 2008. The result of the first study suggest that gasoline price increases were passed through on the first or second day while price decreases where passed through on the fourth day. A similar pattern could be observed for diesel prices with pass through durations of one day and three days, respectively (BWB 2008). In 2010, the BWB conducted another study to ascertain whether the results were still valid after the implementation of the first Fuel Price Fixing Act. They split their full sample from 2004 to 2010 into two subsample periods. For the first period, from September 2004 to November 2007 they found similar evidence as in the previous study. For the second period, from November 2007 to February 2010, no evidence for asymmetry could be found (OECD, 2013).

Apart from the official investigations by the BWB, only few studies specifically focus on the Austrian retail fuel market. Meyler (2009) analyzes the pass through of oil prices into consumer liquid prices, including gasoline, diesel and heating fuel oil, for the euro area using a bivariate error correction model. For the case of Austria, the author does not identify significant asymmetries in pass through. 50% of a wholesale price shock to the gasoline market is adjusted in approximately three weeks. Wlazlowski et al. (2009) consider single country and cross-country asymmetries in the euro area. They report some evidence of cross-country effects for Austria but cannot find any vertical asymmetry. Similarly, Arpa et al. (2006) cannot detect asymmetric price reactions to changes in oil prices. However, the adjustment speed of the Austrian fuel market seems to be among the lowest in the euro area.

Remarkably, the latter three studies do not find asymmetries in Austria whereas the first study of the BWB does. One explanation for the conflicting results could be the frequency of the data employed by the specific studies. The latter three studies use weekly data instead of daily data. It would not be surprising that the price adjustment appears to be symmetric when data is sampled weekly but the true speed of adjustment could only be measured at a daily frequency. The speed of pass through estimated by the BWB is completed within three days. Therefore, one could imagine that possible asymmetries are smoothed out in weekly intervals. In this context, Meyler (2009) cautions against the use of low frequency data for the analysis of potentially asymmetric price transmissions and suggest to use a weekly frequency at the minimum.

Dewenter and Heimeshoff (2012) analyze the effects of the first Austrian fuel price regulation in 2009 on the price levels using a difference-in-differences method and weekly data from 2005 to 2012. They find empirical evidence for a negative price effect of the regulation which implies that fuel price levels decreased after the implementation. In this regard, the policy measure seems to meet the positive expectations of the regulatory authorities. Concerning the Western Australian regulation, Dewenter and Heimeshoff (2012) find no statistically significant effects of the regulation on price levels. Consequently, the regulation reduces the price volatility but does not necessarily foster competition.

Wang (2009) investigates the dynamic pricing strategies in fuel markets before and after the implementation of price regulations. Empirical evidence is found that retailers in Western Australia have engaged in tacit pricing coordination. Also, the price regulation reduced the price levels only for the first four month after the implementation but did not significantly lower average prices in the long-run. This might imply some learning effects of fuel retailers who had to adjust to a new market structure.

Suvankulov et al. (2012) focus on the price regulations and price convergences of retail fuel markets in Nova Scotia and New Brunswick. For this purpose, they study the pricing behaviour of retailers in 60 Canadian cities which include 6 cities from Nova Scotia and 9 cities from New Brunswick using a monthly dataset from January 2000 to October 2010. They find conflicting results for Nova Scotia and New Brunswick. Since the regulation was implemented in 2006, the prices of the 9 cities in New Brunswick converged to the national mean and the volatility reduced significantly. In contrast, there is no significant convergence in Nova Scotia with increased volatility and overall higher price levels. Sen et al. (2011) also investigate the price regulations implemented in Atlantic Canadian provinces and report that prices in the post-regulation period are generally higher.

4 Empirical Methodology

In the empirical analysis of the fuel price data, we study both short-run and long-run asymmetry. Asymmetric adjustment to a long-run equilibrium can only be modelled in a meaningful way if input and output prices are cointegrated. Since the long-run relationship is not exactly known, we estimate the linear cointegrating regression

$$y_t = \beta_0 + \beta_1 x_t + u_t, \tag{1}$$

where y_t is the output price, x_t is the input price and u_t is the disturbance term that may be serially correlated. We assume that both prices are integrated of order one. The variables are cointegrated if the deviations from the long-run equilibrium, u_t, are only temporary. Following the two-step procedure developed by Engle and Granger (1987), we use an auxilliary ADF regression to test if the cointegration residuals are stationary,

$$\Delta z_t = \varrho\, z_{t-1} + \sum_{j=1}^{k} \gamma_j\, \Delta z_{t-j} + \varepsilon_t. \tag{2}$$

The null hypothesis of no cointegration is given by $\varrho = 0$ while the alternative is given by $-2 < \varrho < 0$. Alternatively, we use a cointegration test described in Enders and Siklos (2001) which accounts for asymmetries in the form of threshold adjustment. After confirming the existence of a cointegration relationship, we can use the cointegrating residuals to specify an asymmetric error correction model (AECM). Similar to Granger and Lee (1989) and Grasso and Manera (2007), we express the change in the output price as

$$\Delta y_t = \alpha^+\, \hat{u}^+_{t-1} + \alpha^-\, \hat{u}^-_{t-1} + \sum_{i=0}^{p} \left(\gamma^+_i\, \Delta x^+_{t-i} + \gamma^-_i\, \Delta x^-_{t-i}\right) + \sum_{i=1}^{q} \delta_i\, \Delta y_{t-i} + \varepsilon_t, \tag{3}$$

where $\hat{u}^+_t$ is equal to $\hat{u}_t$ if $\hat{u}_t > 0$ and to zero if $\hat{u}_t < 0$. Correspondingly, $\hat{u}^-_t$ is equal to $\hat{u}_t$ if $\hat{u}_t < 0$ and to zero if $\hat{u}_t > 0$. The coefficients α^+ and α^- measure the speed of adjustment of the output price after positive or negative deviations from the long-run equilibrium, respectively. Consumer welfare losses due to long-run asymmetry are found if $\alpha^+ > \alpha^-$. The null hypothesis of symmetric long-run adjustment,

$$H_{01} : \alpha^+ = \alpha^-, \tag{4}$$

can be tested using an F-statistic (F_{LR}).

Short-run asymmetries are captured by decomposing the first differences into Δx^+_{t-i} if $\Delta x_{t-i} > 0$ and Δx^-_{t-i} if $\Delta x_{t-i} < 0$. We investigate the short-run impact of input price changes through three different channels. First, we test whether the short-run transmission is symmetric at any lag. The null hypothesis is given by

$$H_{02} : \gamma^+_1 = \gamma^-_1 \wedge \gamma^+_2 = \gamma^-_2 \wedge \cdots \wedge \gamma^+_p = \gamma^-_p\,. \tag{5}$$

and tested using an F-statistic which is denoted by F_{SR}. Second, we test for equality of the cumulated effects. We denote the F-statistic computed for the null hypothesis of symmetric cumulated effects,

$$H_{03} : \sum_{i=1}^{p} \gamma_i^+ = \sum_{i=1}^{p} \gamma_i^-, \tag{6}$$

as F_{CE}. Finally, we test whether positive and negative input price shocks are transmitted fully. The null hypothesis is given by

$$H_{04a} : \sum_{i=1}^{p} \gamma_i^+ = 1, \tag{7}$$

for input price increases and

$$H_{04b} : \sum_{i=1}^{p} \gamma_i^- = 1, \tag{8}$$

for input price decreases. The corresponding F-statistics are denoted by F_{FA^+} and F_{FA^-}, respectively.

Additionally, we generalize the AECM and decompose the error correction term $\hat{u}_t$ using an optimally chosen threshold value. The threshold error correction model (TECM) takes the form of

$$\Delta y_t = \alpha_1 \, I_t \, \hat{u}_{t-1} + \alpha_2 \, (1 - I_t) \, \hat{u}_{t-1} + \sum_{i=0}^{p} \left(\gamma_i^+ \, \Delta x_{t-i}^+ + \gamma_i^- \, \Delta x_{t-i}^- \right) + \sum_{i=1}^{q} \delta_i \, \Delta y_{t-i} + \varepsilon_t, \tag{9}$$

where the Heaviside indicator variable I_t is defined by

$$I_t := \begin{cases} 1 & \hat{u}_{t-1} \geq \tau \\ 0 & \hat{u}_{t-1} < \tau \end{cases} \tag{10}$$

using the threshold parameter τ. Enders and Siklos (2001) suggest to follow Chan (1993) and select the threshold τ by minimizing the sum of squared errors of the regression to obtain a superconsistent estimator for τ. In practice, this is achieved by estimating many regressions with fixed τ, where τ runs through all values of $\hat{u}_t$. We apply a 15% lateral trimming to avoid τ^* becoming too close to the boundary and hence ensuring a sufficient number of observations in each regime. Consequently, the AECM can be interpreted as a special case of TECM with fixed threshold value $\tau = 0$.

The statistical properties of asymmetry tests in residual-based threshold cointegration models are discussed in Schild and Schweikert (2017). While we could in principle test for short-run asymmetry using standard F-tests[4], we have to employ bootstrap tests for long-run asymmetry to accommodate the fact that we use cointegration residuals as proxies for the deviations from long-run equilibrium. A bootstrap algorithm for all symmetry tests in the AECM is given as follows:

(1) Estimate the long-run equilibrium equation to obtain $\hat{\beta}_0$, $\hat{\beta}_1$ and the cointegration residuals $\hat{u}_t$. Estimate the AECM and conduct the short-run and long-run asymmetry tests, i.e. compute F_{LR}, F_{SR}, F_{CE}, F_{FA^+} and F_{FA^-}.

(2) Estimate the auxilliary ADF regression model in (2) to obtain $\hat{\varrho}$, $\hat{\gamma}_1, \ldots, \hat{\gamma}_p$ and save the residuals $\hat{\varepsilon}_t$.

(3) Draw randomly from the residuals $\hat{\varepsilon}_t$ to obtain a bootstrap sample ε_t^b.

(4) Generate the bootstrap cointegration residuals series as $\Delta u_t^b = \hat{\varrho} u_{t-1}^b + \sum_{j=1}^{k} \hat{\gamma}_j \, \Delta u_{t-j}^b + \varepsilon_t^b$ and use $(u_1^b, \ldots, u_p^b) = (\hat{u}_1, \ldots, \hat{u}_p)$ as initial observations.

(5) Generate the bootstrap variable $y_t^b = \hat{\beta}_0 + \hat{\beta}_1 x_t + u_t^b$.

(6) Estimate the long-run equilibrium equation for y_t^b and x_t and re-estimate the AECM to compute the bootstrap F-statistics, F_{LR}^b, F_{SR}^b, F_{CE}^b, $F_{FA^+}^b$ and $F_{FA^-}^b$.

(7) Repeat (2) to (6) sufficiently often to obtain the empirical distributions of the bootstrap F-statistics. Compute the p-values for F_{LR}, F_{SR}, F_{CE}, F_{FA^+} and F_{FA^-} based on the bootstrap distribution.

The algorithm can easily be adapted for the TECM using an optimizing threshold value.

5 Empirical Analysis

We start our analysis by testing all price series for their order of integration. For this matter, we apply Dickey-Fuller and KPSS tests to the prices and to the returns. If each series is determined to be integrated of order one, we proceed with our cointegration analysis and estimate the error correction models.

[4]Bootstrap tests and standard F-tests for short-run asymmetry lead to the same test decision in our empirical application.

5.1 Data, unit root and cointegration tests

The transmission of crude oil price changes to pre-tax retail gasoline and diesel prices is analyzed using daily data from August 2004 to March 2016. Since crude oil is not traded on weekends, we delete the fuel price observations for weekends to obtain a matching sample. Austrian retail gasoline and diesel prices, expressed in Euro per 1000 litres, are obtained from two sources. Prices from August 2004 to December 2011 are collected by the ÖAMTC as part of voluntary disclosures by retailers.[5] The prices from January 2012 to March 2016 are obtained from the E-Control database. Following the law on transparency of prices, retailers were required to submit their prices beginning in July 2011. The European benchmark price for crude oil is Brent which is obtained from Thomson Reuters Datastream. Their source is the Energy Information Administration (EIA) of the U.S. Department of Energy. The price is reported in US Dollars per barrel. Therefore, the Brent crude oil price is converted to Euro per 1000 litres using the USD/EUR exchange rate and the corresponding conversion rate to ensure an ordinary comparison of the prices. Retail gasoline and diesel prices without tax and duty are used to exclude the possibility that the taxation structure affects our results.

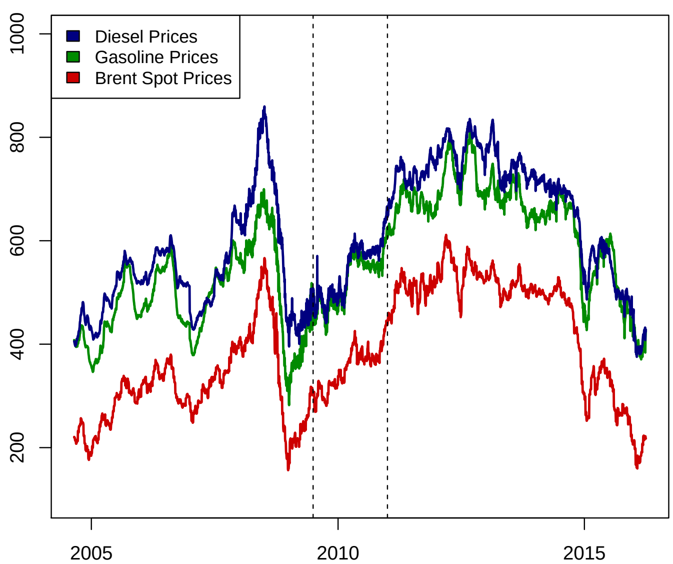

Figure 1: Fuel prices in Euro/1000l. The vertical dashed lines mark the timing of the first and second fuel price regulation in Austria.

[5]On an average day in the period 2004-2008, approximately 61% of all retailers submit their prices which allows for a good approximation of Austrian retail fuel prices (BWB 2008). However, since the average prices are computed based on voluntarily submitted prices by retailers, we have to approach the quality of our data from 2004 to 2011 with some skepticism.

The model specification is based on raw price levels instead of log prices following the study of Meyler (2009) in which the author argues that the long-run relationship between price levels is relatively stable while the relationship between log prices fluctuates substantially over a long sample. We estimate our models for a pre-regulation period, a between-regulations and a post-regulation period to investigate whether the characteristics of the price transmission process have changed after the implementation of two fuel price regulations. The first sample period (pre-regulation) runs from August 2004 to June 2009, the second subsample (between-regulations) spans from July 2009 to December 2010 and the last subsample (post-regulation) spans from January 2011 to March 2016.

Figure 1 shows the evolution of the Brent price series and the Austrian retail gasoline and diesel prices for the full sample period. It is visible that the trajectories move closely together. Moreover, extreme fluctuations can be observed during the Financial Crisis. Pre-tax and duty retail diesel prices are on average higher than gasoline prices. Nonetheless, we observe prices at the pump in Austria which are higher for gasoline than diesel prices. This can be explained by a higher mineral oil tax for gasoline than for diesel (48.2 vs. 39.7 cents per litre in 2016).

Table 1: Unit root tests for crude oil and fuel prices

	ADF				KPSS			
	drift	lags	*trend*	lags	*drift*	lags	*trend*	lags
p^c	-1.826	5	-1.324	5	8.836^{***}	9	2.897^{***}	9
p^g	-2.313	10	-2.086	10	10.140^{***}	9	2.379^{***}	9
p^d	-1.971	10	-1.517	10	8.015^{***}	9	2.455^{***}	9
Δp^c	-21.433^{***}	4	-	-	0.296	9	-	-
Δp^g	-13.280^{***}	9	-	-	0.220	9	-	-
Δp^d	-13.638^{***}	9	-	-	0.374^{*}	9	-	-

Note: p^c denotes the daily crude oil price, while p^g and p^d denote the daily gasoline and diesel price, respectively. Including an intercept in the ADF/KPSS test equation is indicated with *drift*, including an additional linear trend term with *trend*. The lag selection for the ADF test was achieved via Akaike Information Criterion (AIC). Since it is highly unrealistic that first differences of prices series follow a linear trend, we do not conduct the unit root test using the trend specification for the first differences.

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

For the empirical analysis, we first test the price series for their order of integration. For this purpose, we employ the ADF test as well as a KPSS test. The null hypothesis of the ADF test is nonstationarity in form of a unit root while the alternative hypothesis is (trend-)stationarity. The optimal lag length selection is achieved by using the AIC. As shown in Table 1 all values for the ADF t-statistic are above the critical 5% value. This means that the null hypothesis cannot be rejected. In contrast, the null hypothesis is rejected for the first differenced series. The KPSS test assumes (trend-)stationarity under the null hypothesis and nonstationarity under the alternative. Here, we reject the stationarity hypothesis at the 5% significance level and cannot reject the null for the first differenced series. This is the case for all variables in all sample periods. Hence, the variables are assumed to be integrated of order one and we can continue our cointegration analysis.

Table 2: Long-run equilibria and cointegration tests

Sample:	08/04 - 06/09		07/09 - 12/10		01/11 - 03/16	
Obs.	1263		393		1367	
	p^g	p^d	p^g	p^d	p^g	p^d
β_0	167.47	172.06	134.80	103.06	248.94	228.57
β_1	1.01	1.17	1.10	1.25	0.856	0.996
EG	-3.698^{**}	-3.348^{**}	-3.887^{**}	-3.524^{**}	-3.427^{**}	-4.428^{***}
Φ_{TAR}	11.07^{***}	8.233^{***}	17.36^{***}	17.70^{***}	9.679^{***}	22.90^{***}

Note: EG denotes the Engle-Granger cointegration test. Critical values for the EG test are 10%: −3.044, 5%: −3.336, 1%: −3.896. Φ_{TAR} denotes the Enders-Siklos test with TAR adjustment. Critical values for the Enders-Siklos test are: 10%: 4.88, 5%: 5.79, 1%: 7.81.

Given that our variables are all integrated of order one, we can use the Engle-Granger two step approach to test for cointegration. As the first step, we estimate the long-run equilibrium relationship based on the linear model (1) to obtain the cointegrating vector. In this study, we treat crude oil prices (p^c) as an exogenous variable and use either gasoline prices (p^g) or diesel prices (p^d) as the dependent variable. The Brent crude oil market is considerably larger than the Austrian retail fuel market so that price shocks coming from an unelastic local fuel demand should not influence the Brent price significantly. This assumption is supported by results from additional error correction models for Brent adjustment which are not reported. The second step of the Engle-

Granger approach involves conducting a unit root test on the least squared residual series to ascertain whether $\hat{u}_t$ indeed constitutes a stationary equilibrium error. For this purpose, we apply the EG test based on an ADF regression where the AIC is used to select the lag truncation parameter. Since the EG test does not account for the possibility of asymmetric adjustment, we also apply the Enders-Siklos threshold cointegration test (Φ_{TAR}). The results are reported in Table 2. We observe that the null hypothesis of no cointegration is rejected in all cases. We can therefore analyze the particular retail fuel adjustment in response to input price shocks.

5.2 Results

According to the Federal Minister of Science, Research and Economy of Austria, the aim of the fuel price regulation in 2011 was to improve competitive conditions for retailers and provide price transparency for consumers (BMDW, 2010). Hence, the policy measure should positively influence the competitive structure of the Austrian fuel market and reduce the search costs of consumers through increased transparency. If the Austrian fuel market was indeed uncompetitive before the price regulation, we should find evidence of asymmetric price transmission in the first subsample. The price regulation should then lead to faster responses in case of falling crude oil prices. We would expect generally faster adjustment rates and less evidence for APT in the second and third subsamples. However, if the retail fuel prices are distorted because retailers anticipate that they cannot raise their prices after noon and charge generally higher prices, this could lead to positive APT in the post-regulation period.

Hence, in order to evaluate the fuel price regulations we first estimate an asymmetric error correction model according to (3) and a threshold error correction model according to (9) for the pre-regulation subsample. The main results are reported in Table 3 and additional estimates are presented in Table 4.

5.2.1 Pre-regulation subsample

Before the introduction of the first regulation, we find some evidence of positive long-run asymmetry for gasoline, but do not reject the null hypothesis of symmetric adjustment for diesel. In contrast to the finding of positive long-run asymmetry, the short-run coefficient estimates reveal that crude oil price decreases are passed through faster than crude oil price increases in the gasoline market. While the gasoline price needs two days to react to crude oil price increases, it responds to decreases on the first day. Further, we observe that crude oil price increases are not completely passed through. Although the diesel price reacts symmetrically to crude oil price changes, it does not completely pass through crude oil price changes (see Table 4). Overall, the diesel adjustment to short-run changes in crude oil appears to be sluggish. We obtain very similar results if the optimal threshold value is estimated from the data (see the TECM results in panel (d) of Table 3).

To better understand the complete dynamics of the adjustment process and to evaluate whether short-run or long-run asymmetries dominate the adjustment process, we also investigate the impulse-response curves of the retail prices. For this matter, we consider a combined long-run and short-run impulse. We model an exogenous input price increase by one cent which also leads to a negative deviation from the long-run equilibrium by one cent leaving the retail fuel price fixed. Since the input price is assumed to be exogenous, the retail fuel price has to react to maintain the long-run equilibrium. Similarly, an input price decrease leads to a positive deviation from the long-run equilibrium. The pass through computations reported in panel (b) and panel (c) of Table 3 are based on the AECM specification.

The pass through of input price shocks is relatively slow for both retail markets in the first subsample. In case of gasoline, the pass through of crude oil price decreases is slower (by six days) than the pass through of increases. Diesel passes through crude oil price decreases slightly faster (by two days) than decreases. In contrast to our results, the first BWB study reports faster adjustment of diesel prices. This might, however, be attributed to the different impulse specifications. It also appears difficult to compare our results against those obtained by other studies on the Austrian fuel market using different model specifications. The results of Meyler (2009) are based on weekly data and can therefore hardly be used as a comparison since the minimal speed of adjustment is one week per construction. Arpa et al. (2006) use a log-specification which makes it difficult to compare the speed of adjustment.

In summary, we can only link some empirical evidence of positive long-run asymmetric adjustment in the Austrian retail gasoline market to the public debate on potential asymmetric fuel price transmission prior to the first fuel price regulation. The corresponding price adjustment in the diesel market seems to be symmetric during this period. Our results on the pre-regulation period are largely in line with the first study of the BWB where Rotterdam fuel spot prices were used instead of crude oil prices. They do not find significant evidence for long-run asymmetry in both retail markets but report substantial short-run asymmetries in the gasoline market (BWB 2008).

Table 3: Results of the error correction models

Sample:	08/04 - 06/09		07/09 - 12/10		01/11 - 03/16	
Obs.	1263		393		1367	
	p^g	p^d	p^g	p^d	p^g	p^d
	Panel (a): AECM					
α^+	-0.010	-0.044	-0.030	-0.185	-0.010	-0.002
α^-	-0.056	-0.029	-0.156	-0.013	-0.048	-0.067
F_{LR}	4.743**	0.780	3.373*	5.704**	4.914**	6.575**
F_{SR}	2.678**	0.893	5.971***	5.105***	10.41***	10.27***
F_{CE}	6.425**	0.237	0.884	4.733**	2.316*	2.399*
F_{FA^+}	11.65***	9.552***	3.998**	1.444	28.00***	14.56***
F_{FA^-}	0.091	10.51***	0.774	2.449	11.91***	4.278**
	Panel (b): 50% pass through of input price increases reached					
t_{days}	12	16	4	2	8	5
	Panel (c): 50% pass through of input price decreases reached					
t_{days}	18	14	4	3	6	4
	Panel (d): TECM					
$\hat{\tau}$	23.94	-28.02	-17.44	15.04	-25.18	16.44
α_1	-0.011	-0.053	-0.065	-0.205	-0.013	0.004
α_2	-0.047	-0.019	-0.149	-0.029	-0.051	-0.065
F_{LR}	3.508	4.745	2.130	7.031**	5.725*	9.310***
F_{SR}	2.448*	1.105	5.834***	8.196***	10.19***	10.22***
F_{CE}	5.374***	1.133	0.063	5.336**	1.559*	1.759*
F_{FA^+}	10.45***	10.51***	3.038	3.064*	28.60***	14.47***
F_{FA^-}	0.521	15.75***	2.238	0.682	16.54***	6.401**

Note: F_{LR} denotes the F-statistic computed for the null hypothesis of symmetric long-run adjustment, $H_{01} : \alpha^+ = \alpha^-$. F_{SR} denotes the F-statistic computed for the null hypothesis of symmetric short-run effects, $H_{02} : \gamma_1^+ = \gamma_1^- \wedge \dots \wedge \gamma_p^+ = \gamma_p^-$. F_{CE} denotes the F-statistic computed for the null hypothesis of symmetric cumulated effect, $H_{03} : \gamma_0^+ + \dots + \gamma_p^+ = \gamma_0^- + \dots + \gamma_p^-$. F_{FA^+} and F_{FA^-} denote the F-statistics computed for the null hypothesis, $H_{04a} : \gamma_0^+ + \dots + \gamma_p^+ = 1$ and $H_{04b} : \gamma_0^- + \dots + \gamma_p^- = 1$, respectively.
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

5.2.2 Between-regulation subsample

After the introduction of the first price regulation, we find very different adjustment characteristics compared to the first subsample. Deviations from the long-run equilibrium seem to be adjusted with greater asymmetry in the gasoline market if we consider the numerical value of the coefficients (see panel (a) of Table 3). However, the difference between the adjustment coefficients of the AECM is only significant at the 10% level. This provides only slight evidence for positive APT in the gasoline market. Also, we find some evidence for asymmetric short-run effects. A closer look at the short-run adjustment estimates reveals that retail prices respond faster to input price decreases than input price increases at the first lag (see Table 4). If the input price changes date back longer, increases are adjusted faster than decreases. The difference of the cumulative effects is insignificant which means that the differences at specific lags average out over the full adjustment process.

The diesel price seems to respond faster to positive deviations from the long-run equilibrium in the between-regulation subsample. We find significant evidence for negative APT in the AECM and the TECM. The short-run transmission of crude oil price changes is asymmetric and the difference of cumulative effects is significant. The diesel price seems to react to transitory input price changes with a similar pattern as the gasoline price. We observe a faster response to crude oil price decreases than to price increases at the first lag but a more rapid pass through of price increases at the remaining lags. In total, the faster response to price increases dominates the cumulative response time which leads to a rejection of H_{03}. The speed of pass through after the first price regulation is now considerably faster for both retail markets and appears to be largely symmetric. Gasoline passes through 50% of a crude oil price shock in four days while diesel needs two to three days.

In summary, the results of the first regulation indicate slight evidence for positive long-run APT in case of gasoline and negative long-run APT in case of diesel as well as asymmetries in the short-run. Remarkably, the speed of pass through from input prices to output prices seems to increase for both retail markets after the implementation of fuel price regulations. However, the results of the between-regulations subsample have to be interpreted cautiously since the sample size is relatively small.

5.2.3 Post-regulation subsample

The longer post-regulation subsample surprisingly reveals significant positive long-run asymmetry for diesel. These results can be attributed to the non-adjustment of very large positive deviations found in the TECM. Still, we observe a faster pass through of transitory input price decreases and a general underreaction to crude oil price changes. Long-run adjustment rates for gasoline are still significantly different and again we find short-run asymmetry following the pattern of previous subsamples. Price decreases from the last period are passed through faster than price increases but the opposite holds if the price changes date back more than two days. In the post-regulation period, both crude oil price increases and decreases are not completely transmitted. The impulse-response analysis for this subsample shows that the short-run effect dominates the long-run effect as input price decreases are passed through slightly faster than input price increases.

Following the implementation of both fuel price regulations, we find statistical evidence for long-run asymmetry in the diesel market. While the numerical values of the adjustment coefficients in the between-regulations subsample indicates negative APT, we report positive APT in the longer post-regulation period. The long-run adjustment of gasoline prices tends to indicate positive APT in all subsamples. However, we reveal a similar asymmetric short-run response pattern in both retail markets. Both retail fuel prices seem to show a substantial underreaction to input price changes after the second price regulation. Considering the numerical estimates, it seems that retailers are particularly cautious in the process of increasing their prices. This might be interpreted as a strategic response caused by the added competitive pressure from fixing the price structure for the next 24 hours. Retailers are forced to position themselves competitively in the market by choosing a maximum price for one day. Hence, they have to pass through crude oil price decreases immediately while they delay crude oil price increases. These adjustments are decomposed into smaller steps and are subsequently passed through over the following days. Unfortunately, our methodology is not able to account for confounding factors such as technological progress or further organizational changes of the Austrian retail fuel market which might coincide with the implementation of fuel price regulations. However, we identify an immediate change of the speed of pass through after the first Fuel Price Fixing Act which remains quite stable in the longer post-regulation period.

6 Conclusion

Our findings in this study shed some light on the implications of price regulations in the Austrian retail fuel market. Our results indicate that the decision to implement fuel price regulations could have been motivated by previously existing asymmetric price transmissions in the gasoline market, but we do not find any statistical evidence for positive APT in the diesel market. Public concerns about possibly collusive behaviour of the Austrian retail fuel industry were noted before the first Fuel Price Fixing Act was implemented and the policy measure was expected to improve competitive conditions for retailers and provide price transparency for consumers. Nevertheless, we still find evidence for asymmetric adjustment after the revision of the Fuel Price Fixing Act in 2011. Retail fuel prices seem to adjust more slowly if crude oil prices are relatively low. In contrast, we mostly find negative short-run APT which indicates competitive pressure to delay input price increases. Only if we consider the combined effect of input price shocks, we observe that the speed of pass through has generally become faster for both retail fuel markets. Crude oil price decreases are now passed through slightly faster than crude oil price increases. Overall, the Fuel Price Fixing Act can be considered a partial success as it seems to have fostered competition between retail filling stations and thereby increased consumer welfare. However, it remains difficult to predict how fuel price regulations would affect other retail fuel markets as their market structures are not identical to the Austrian retail fuel market. Future research might consider the identification of causal effects of fuel price regulations which unfortunately cannot be measured under the present research design and data availability.

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8 Appendix

Table 4: Additional estimates for the AECM

Sample:	08/04 - 06/09		07/09 - 12/10		01/11 - 03/16	
Obs.	1263		393		1367	
	p^g	p^d	p^g	p^d	p^g	p^d
α^+	-0.010	-0.044	-0.030	-0.185	-0.010	-0.002
α^-	-0.056	-0.029	-0.156	-0.013	-0.048	-0.067
γ_0^+	0.264^{***}	0.294^{***}	0.288^{***}	0.485^{***}	0.139^{***}	0.162^{***}
γ_1^+	-0.083	0.014	-0.247^{**}	-0.124	-0.072	-0.044
γ_2^+	0.006	0.001	0.159	0.356^{***}	0.169^{***}	0.187^{***}
γ_3^+	0.114^{**}	0.114^{**}	0.174^{*}	0.266^{**}	0.066	0.095^{**}
γ_4^+	0.105^{**}	0.111^{**}	0.111	0.328^{***}	0.059	0.077^{*}
γ_5^+	0.144^{**}	0.095	0.005	-	0.063	0.099^{**}
γ_0^-	0.363^{***}	0.314^{***}	0.101	0.082	0.086^{*}	0.110^{**}
γ_1^-	0.152^{***}	0.013	0.536^{***}	0.484^{***}	0.392^{***}	0.417^{***}
γ_2^-	0.101^{**}	0.062	-0.110	-0.048	0.013	0.034
γ_3^-	0.066	0.004	0.047	0.025	0.071	0.081^{*}
γ_4^-	0.044	0.026	0.176^{*}	0.018	0.106^{**}	0.108^{**}
γ_5^-	0.237^{***}	0.124^{**}	0.025	-	0.040	0.027
δ_1	-0.407^{***}	-0.424^{***}	-0.253^{***}	-0.427^{***}	-0.195^{***}	-0.279^{***}
δ_2	-0.197^{***}	-0.141^{***}	-0.106^{**}	-0.268^{***}	-0.103^{***}	-0.161^{***}
δ_3	-0.009	-0.049	-0.138^{***}	-0.203^{***}	-0.011	-0.022
δ_4	-0.091^{***}	-0.021	-0.084	-0.140^{**}	0.025	-0.020
δ_5	0.132^{***}	0.190^{***}	0.335^{***}	0.193^{***}	0.153^{***}	0.077^{***}

Note: Results of individual significance tests for the long-run adjustment coefficients are not indicated, since their significance is evaluated using an F-test (see Table 3). The number of lags is based on the AIC and on tests for residual autocorrelation.
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

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