Price Regulations and Price Adjustment Dynamics: Evidence from the Austrian Retail Fuel Market

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Abstract

This paper investigates fuel price regulations implemented in Austria prohibiting retailers from raising their prices more than once per day. We analyze price transmission dynamics over three subsamples and still find evidence for asymmetric adjustment in the post-regulation period. Considering the combined effect of input price changes reveals that gasoline now passes through input price changes faster in the post-regulation period. However, we do not obtain the same finding for diesel where only input price increases are transmitted significantly faster. Hence, we conclude that the Austrian fuel price regulation seems to have been a partial success in terms of efficiency in price transmission.

Keywords: Asymmetric price transmission; diesel; gasoline; nonlinear error correction model; price regulation; spot prices; retail fuel prices

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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 often 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 behavior is generally known as asymmetric price transmission (APT). From the perspective of standard economic theory such APTs 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 fuel price transmissions 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 (2019)). 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. Short-run APT describes the asymmetric effects of positive or negative input price changes on output prices so that pass through depends on the sign 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. A combination of short-run and long-run APT is possible and it is usually difficult to determine 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 consumer's perspective, 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 behavior (see, among others, Deltas (2008), Verlinda (2008), Tappata (2009), Lewis and Noel (2011) and Remer (2015)). One of the primary causes could be the exploitation of market power by large oil companies (Borenstein, 1991; Shepard, 1991). 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 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 signaling 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 enabling them to maintain high prices in the short run.

Another cause for asymmetric pricing is known as the search cost hypothesis (Johnson, 2002; Yang and Ye, 2008; Lewis, 2011). It states that consumers have a greater incentive to search for lower prices during periods of increasing prices while their incentive to search is inhibited in periods of falling prices. Since less search effort by consumers leads to higher margins and more price dispersion in the market, retailers maximize profits by responding slower to input price decreases. Put differently, retailers face a different competitive environment depending on the consumer search intensity and react strategically to these situations. This also implies that retailers should pass through price decreases faster in markets where search

[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'.

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 behavior in a fuel market has substantial consumer welfare implications since under positive APT consumers do not benefit from price reductions to the same extent as they would under symmetric adjustment. This welfare misallocation could be resolved by appropriate policy measures. Several regulatory measures 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 was expected to increase price transparency for consumers by limiting the number of price changes per day and intended to foster competition. However, the real effects of the proposed policy measure were unclear a priori and have to be evaluated empirically.

This paper investigates price transmissions from input prices to retail fuel prices in Austria using nonlinear error correction models. In particular, we investigate the pass through of fuel 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 price adjustment dynamics. While the first price regulation in July 2009 limited 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 fuel price transmissions might be twofold. On the one hand, fixing a date for price increases could have brought greater price transparency which might have lowered 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. Hence, the speed of adjustment after negative cost shocks should increase in our post-regulation subsample. 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. This would in turn allow them to delay price decreases more effectively compared to the pre-regulation period. The speed of adjustment after input price increases would then be slower for the post-regulation subsample. 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 if and how the Austrian fuel price regulation has influenced efficiency in price transmission by distinguishing between 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 APT are the following: First, whereas the majority of studies on APT provide empirical evidence for one or more countries over a specific period of time, we investigate whether 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 behavior using data obtained after the second fuel price regulation was implemented in 2011. 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.

[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.

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 and discusses economic implications of price regulation policies. Section 3 outlines the methodology used to model potentially APTs. In Section 4, we apply these models to Austrian fuel price data and Section 5 offers a conclusion.

2 Economic implications of 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 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 submit their price changes within 30 minutes to 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. This form of fuel price regulation reduces daily

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

price volatility by construction but the regulatory authorities intent to provide transparency and reduce consumer search costs as well (Byrne, 2014).

Currently, five of thirteen Canadian 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. The designated fuel prices at the pump are determined from the benchmark spot fuel price, applicable taxes, wholesale and retail margins and transportation costs. In New Brunswick, only the maximum price is set by the regulatory authority (Suvankulov et al., 2012). Luxembourg maintains a similar price ceiling mechanism.

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 suggest that prices in regulated markets are higher and less volatile compared to non-regulated markets. In a further lab experiment, Haucap and Müller (2012) investigate the effects of three different fuel price regulations (Luxembourg, Western Australia, Austria). They find that Austrian-type and Luxembourg-type regulations decrease consumer welfare while the Western Australian regulation does not. In addition, none of them lowers retail prices. 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. He argues that retailers intertemporally distort prices in a way so that their profits remain unchanged compared to the unregulated market.

One of the first empirical studies on fuel price regulations is conducted by Wang (2009) who investigates the dynamic pricing strategies before and after the implementation of price regulations. The Western Australian 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. Suvankulov et al. (2012) focus on price regulations and price convergence of retail fuel markets in Nova Scotia and New Brunswick. After the price regulation was implemented, the prices of nine 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 effects of price regulations implemented in Atlantic Canadian provinces and report that prices in the post-regulation period are generally higher. Finally, Dewenter et al. (2017) analyze the effects of the Austrian and Western Australian fuel price regulations on fuel price levels using a difference-in-differences method. They find empirical evidence for a negative price effect of the Austrian-type regulation. In this regard, the policy measure seems to have met the positive expectations of regulatory authorities. Concerning the Western Australian regulation, Dewenter et al. (2017) find no statistically significant effects of the regulation on price levels. Consequently, the regulation seems to reduce price volatility but does not necessarily foster competition.

Although an extensive empirical literature on asymmetric fuel pricing exists (see Perdiguero (2013) for a recent overview), only a few studies can be found that analyze asymmetric behavior in connection with fuel price regulations. This can in parts be explained by the small number of developed countries with regulated fuel markets. The BWB (Bundeswettbewerbsbehörde) has conducted several 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 behavior in response to wholesale price changes. The results of the first study suggest that gasoline price increases were passed through on the first or second day while price decreases were 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).

Meyler (2009) analyzes the pass through of oil prices into consumer liquid prices, including gasoline, diesel and heating fuel oil, for the euro area. 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) and Wlazlowski et al. (2012) caution against the use of low frequency data for the analysis of potentially APTs and suggest to use at least a weekly sampling frequency.

3 Empirical Methodology

In the empirical analysis of our fuel price data, we study both short-run and long-run asymmetry. Asymmetric adjustment to a long-run equilibrium can only be modeled 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 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,

$$\Delta u_t = \varrho u_{t-1} + \sum_{j=1}^{p} \gamma_j \Delta u_{t-j} + \varepsilon_t, \tag{2}$$

to test if the cointegration residuals are stationary. 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). In order to keep notation as simple as possible, we describe the baseline specification without regime-specific dummy variables which are used to indicate subsamples. 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 p, q denote the respective lag length, $\hat{u}^+_t = \max\{0, \hat{u}_t\}$ and $\hat{u}^-_t = \min\{\hat{u}_t, 0\}$. 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 $\Delta x_{t-i}^{+} = \max\{0, \Delta x_{t-i}\}$ and $\Delta x_{t-i}^{-} = \min\{\Delta x_{t-i}, 0\}$. We test whether cumulative short-run adjustment is symmetric for positive and negative input price changes. The null hypothesis is given by

$$H_{02} : \sum_{i=1}^{p} \gamma_i^{+} = \sum_{i=1}^{p} \gamma_i^{-}. \tag{5}$$

The corresponding F-statistics is denoted by F_{SR}.

The statistical properties of asymmetry tests in residual-based threshold cointegration models are discussed in Schild and Schweikert (2019). They find that asymmetry tests can be strongly undersized if conventional critical values are used. While we could in principle test for short-run asymmetry using critical values from an appropriate F-distribution[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 long-run 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 compute F_{LR}.

(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 $\Delta u_t^b = \hat{\varrho} u_{t-1}^b + \sum_{j=1}^{p} \hat{\gamma}_j \Delta u_{t-j}^b + \varepsilon_t^b$ 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-statistic F_{LR}^b.

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

(7) Repeat (2) to (6) sufficiently often to obtain the empirical distributions of the bootstrap F-statistic. Compute the p-value for F_{LR} based on the bootstrap distribution.

The algorithm can easily be adapted to account for regime-specific coefficients.

4 Empirical Analysis

We start our analysis by testing all price series for their order of integration. For this matter, we apply ADF 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 error correction models.

4.1 Data, unit root and cointegration tests

The transmission of input price changes to pre-tax retail gasoline and diesel prices is analyzed using daily data from August 2004 to March 2016. Austrian retail gasoline and diesel prices, expressed in Euro per 1000 liters, 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 adjustment of retail fuel prices is measured using regional spot prices for refined fuel as a proxy for the cost structure. An alternative would be to use crude oil prices but according to Borenstein et al. (1997) the demand of other products derived from crude oil might distort the perceived cost pass through relationship. European spot prices for gasoline and diesel are obtained from Thomson Reuters Datastream. All prices are reported in Euro

[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 scepticism.

per 1000 liters to ensure an ordinary comparison of prices. Retail gasoline and diesel prices without tax and duty are used to exclude the possibility that the taxation structure affects our results.

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 setting the corresponding dummy variables 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) reaches 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 Austrian retail gasoline and diesel prices as well as their respective spot prices for the full sample period. We observe that both retail-spot trajectories move closely together. Extreme fluctuations can be observed during the Great 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 liter in 2016).

For the empirical analysis, we first test all price series for their order of integration. The null hypothesis of the ADF test is nonstationarity in form of a unit root while the alternative hypothesis is (trend-)stationarity. 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

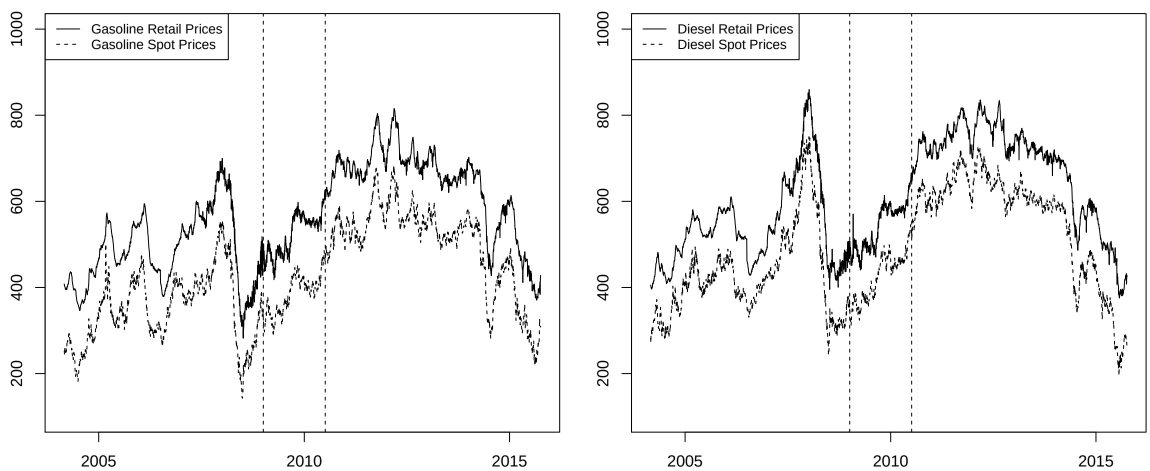

Figure 1: Fuel prices in Euro/1000L. The vertical dashed lines mark the timing of the first and second fuel price regulation in Austria.

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

	ADF				KPSS			
	drift	lags	*trend*	lags	*drift*	lags	*trend*	lags
p^g	−2.313	10	−2.086	10	10.140***	9	2.379***	9
$p^{spot,g}$	−2.243	1	−1.992	1	10.130***	9	2.360***	9
p^d	−1.971	10	−1.517	10	8.015***	9	2.455***	9
$p^{spot,d}$	−1.632	4	−1.157	4	6.726***	9	2.493***	9
Δp^g	−13.280***	9	-	-	0.220	9	-	-
$\Delta p^{spot,g}$	−38.620***	1	-	-	0.140	9	-	-
Δp^d	−13.638***	9	-	-	0.374*	9	-	-
$\Delta p^{spot,d}$	−27.210***	3	-	-	0.304	9	-	-

Note: p^g and p^d denote the daily retail gasoline and diesel price, respectively. The superscript *spot* indicates the corresponding spot fuel prices. Including an intercept in the ADF/KPSS test equation is indicated by *drift*, including an additional linear trend term by *trend*. The lag selection for the ADF test was achieved via Akaike Information Criterion (AIC). Since it is highly unrealistic that first differences of price series follow a linear trend, we do not conduct the unit root test using the trend specification for the first differences.
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

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.

Given that our variables are all integrated of order one, we can use the Engle-Granger

Table 2: Long-run equilibria and cointegration tests

Sample:	08/04 - 06/09		07/09 - 12/10		01/11 - 03/16	
	p^g	p^d	p^g	p^d	p^g	p^d
β_0	151.29	135.93	151.75	158.75	157.01	182.32
β_1	0.961	0.953	0.969	0.904	0.961	0.895
EG	-3.905^{***}	-4.520^{***}	-3.621^{**}	-4.366^{***}	-6.957^{***}	-6.121^{***}
Φ_{TAR}	16.36^{***}	21.02^{***}	14.41^{***}	23.83^{***}	52.88^{***}	44.56^{***}

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.

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 consider either gasoline spot prices ($p^{spot,g}$) or diesel spot prices ($p^{spot,d}$) as an exogenous variable and use the corresponding retail prices as the dependent variable. The European spot fuel market is considerably larger than the Austrian retail fuel market so that price shocks coming from an unelastic local fuel demand should not influence spot prices significantly. This implies that spot prices are at least weakly exogenous which is supported by results from additional error correction models for spot price adjustment (not reported). The second step of the Engle-Granger approach involves conducting a unit root test on the 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.

4.2 Results

According to the Federal Minister of Science, Research and Economy of Austria, the primary goal 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 positive APT in the first subsample. The price regulation should then lead to faster responses in case of falling crude oil prices. Reduced search costs for consumers should improve the overall speed of adjustment and also lead to more symmetric responses. If the policy has met its intended goals, we would expect generally faster adjustment rates and less evidence for positive 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.

In order to evaluate the Austrian Fuel Price Fixing Act we first estimate an asymmetric error correction model according to Equation (3) for the pre-regulation subsample.[6] The main results are reported in Table 3. [7]

4.2.1 Pre-regulation subsample

Before the introduction of the first regulation, we surprisingly find evidence of negative long-run asymmetry for gasoline and diesel. The speed of adjustment after disequilibrium states in which retail prices are relatively too high is much faster than in opposing situations in which

[6]Since we use national aggregates to estimate our models, we implicitly impose a representative agent assumption for retailers. Lippi (1988) shows that different adjustment characteristics on the micro-level might lead to cross-sectional aggregation bias in error correction models, which could overstate the estimated dynamic process toward the long-run equilibrium. Balaguer and Ripollés (2016) provide empirical support that error correction models overestimate the persistence of shocks for aggregated data. Ideally, we would estimate our models for individual station data. However, these data are not available over the full sample period.

[7]We have reestimated our models under the possibility of asymmetric autoregressive dynamics, i.e. distinguishing whether $\Delta y_{t-i} \geq 0$ or $\Delta y_{t-i} < 0$ for all lags $i = 1, \ldots, q$. The qualitative interpretation of our results does not change.

Table 3: Results of the error correction models

Sample:	08/04 - 06/09		07/09 - 12/10		01/11 - 03/16	
	p^g	p^d	p^g	p^d	p^g	p^d
	Panel (a): AECM coefficient estimates					
α^+	-0.117	-0.185	-0.128	-0.401	-0.133	-0.081
α^-	-0.054	-0.047	-0.173	-0.166	-0.175	-0.194
γ_0^+	-0.044	0.043	-0.026	-0.080	0.080**	0.098**
γ_1^+	-0.004	0.003	-0.134	-0.125	0.055	0.008
γ_2^+	0.056	0.060	0.083	-0.114	0.002	0.033
γ_3^+	0.055	0.071*	0.177*	0.051	0.083**	0.105**
γ_4^+	0.036	0.018	0.151	0.128	0.084**	0.135***
γ_5^+	0.132***	0.045	0.060	0.102	0.100**	0.060
γ_0^-	0.039	0.040	-0.074	-0.039	0.041*	0.033
γ_1^-	-0.127***	-0.146***	-0.054	-0.379***	-0.016	0.041
γ_2^-	-0.059*	-0.105**	-0.006	-0.249**	0.179***	0.248***
γ_3^-	0.067	-0.027	0.054	0.070	0.116***	0.112**
γ_4^-	0.026	-0.007	0.200**	-0.033	0.103**	0.079
γ_5^-	0.029	0.065	0.156	0.057	0.086**	0.137***
δ_1	-0.406***	-0.406***	-0.287***	-0.334***	-0.290***	-0.315***
δ_2	-0.216***	-0.155***	-0.157***	-0.245***	-0.202***	-0.201***
δ_3	-0.040	-0.074	-0.202***	-0.170***	-0.096***	-0.058*
δ_4	-0.128***	-0.039	-0.181***	-0.178**	-0.054*	-0.035
δ_5	0.153***	0.178***	0.326***	0.249***	0.166***	0.129***
	Panel (b): Hypothesis tests					
F_{LR}	10.09***	32.43***	0.430	10.35***	1.184	8.361***
F_{SR}	8.450***	25.69***	0.023	14.83***	1.235	5.443**

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-statistics computed for the null hypothesis, $H_{02} : \gamma_0^+ + \cdots + \gamma_p^+ = \gamma_0^- + \cdots + \gamma_p^-$. The number of lags is based on the AIC and on tests for residual autocorrelation.
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

spot fuel prices are relatively too low. Similarly, we find statistically significant asymmetries in the cumulative short-run reaction to cost shocks. In general, we report weak short-run adjustment behavior. It takes retail fuel prices one day to react to positive cost shocks, and yet more surprisingly, short-run coefficients at the first lag have the wrong sign which amplifies the spread between input and output prices. Cost shocks seem to be mostly passed through by the long-run adjustment component of the system. Since the combined short-run and long-run dynamics are responsible for the overall adjustment path, we cannot interpret single coefficients and must investigate the combined effect of cost shocks.

To better understand the complete price adjustment dynamics and to evaluate whether short-run or long-run asymmetries dominate the adjustment process, we analyze 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 curves are depicted in Figure 2 and Figure 3. We employ a stationary bootstrap according to Politis and Romano (1994) to draw bootstrap samples and compute confidence bands for the impulse-response curves.

The pass through of input price shocks is relatively slow for both retail markets in the first subsample. Both retail markets take more than four weeks to adjust 90% of a cost shock. In case of gasoline, we find that the pass through of positive cost shocks is slower than the pass through of negative cost shocks. 50% of input price increases are passed through after 15 days whereas input price decreases are passed through after only 10 days. Similarly, diesel passes through spot diesel price decreases a lot faster (by ten days) than increases.

Our results on the pre-regulation period are largely in line with the first study of the BWB (BWB 2008). They do not find significant evidence for long-run asymmetry in both retail markets but report substantial short-run asymmetries in the gasoline market. Using a different impulse specification, they also report faster adjustment of diesel prices for the same subsample. Most existing studies are limited to weekly frequencies so that the minimal speed of adjustment is one week per construction. The results of Meyler (2009) are based on weekly gasoline prices and he finds that 50% (90%) pass through is reached after three (ten) weeks. Arpa et al. (2006) also use weekly data and report no significant difference between the speed of adjustment after rising and falling crude oil prices in Austria.

In summary, we find it difficult to underpin the public debate about large oil companies delaying input price changes with empirical results. While the adjustment seems to be asymmetric for both retail markets, we do not find delays of negative cost shocks which might be detrimental to consumer welfare. Instead, positive cost shocks are delayed in the Austrian fuel markets which means that consumers benefit from retail fuel prices being relatively too low over an extended period of time. Since the discussion usually involves the relationship between retail prices and crude oil prices, we have reestimated our models using Brent prices instead of spot fuel prices.[8] The speed of adjustment after crude price changes is generally slower than after spot fuel price changes. Here, we can find slight empirical evidence of positive long-run asymmetric adjustment in the Austrian gasoline market. Still, we do not find empirical evidence for positive APT in the Austrian diesel market.

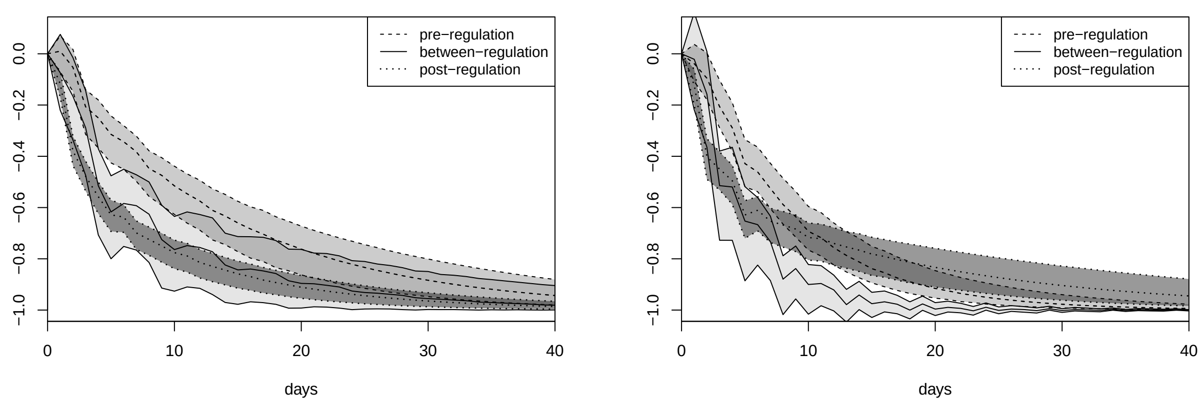

Figure 2: Response of gasoline (left) and diesel (right) after negative cost shocks

[8]The results are not reported but can be obtained from the authors upon request.

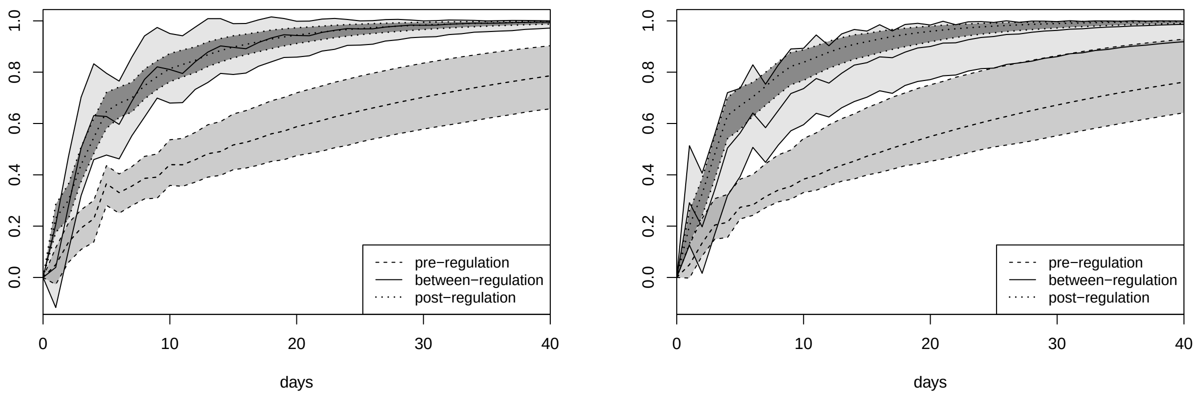

Figure 3: Response of gasoline (left) and diesel (right) after positive cost shocks

4.2.2 Between-regulations 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 positive 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 for gasoline is no longer statistically significant. This can in parts be attributed to larger standard errors in the smallest regime and, hence, provides only weak evidence for positive APT during this period. The change of α^- from the pre-regulation to the between-regulations period is significant at the 5% level (α^+ does not change significantly). Our results imply that it has become easier for retailers to adjust gasoline prices to increasing input price levels after the implementation of the first Fuel Price Fixing Act. In contrast, we do not find evidence for asymmetric short-run effects. A closer look at the short-run adjustment estimates reveals that retail prices respond slightly faster (third lag) to input price decreases than input price increases (fourth lag).

Again, we find significant evidence for negative APT in the diesel market. Diesel responds considerably faster to positive deviations from the long-run equilibrium in the between-regulations subsample compared to the pre-regulation subsample. Both long-run adjustment coefficients change significantly from the pre-regulation to the between-regulations period. Diesel reacts to short-run input price changes with a different pattern to the one observed for gasoline. Although we estimate a faster response to spot diesel price decreases than to price increases, the sign of the adjustment coefficients is negative which amplifies the spread between retail diesel and spot diesel prices. Combining short-run and long-run dynamics, we find that the speed of pass through after the first price regulation is now considerably faster for both retail markets. Gasoline passes through 50% of an input price decrease (increase) in three (four) days while diesel needs four (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. 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.

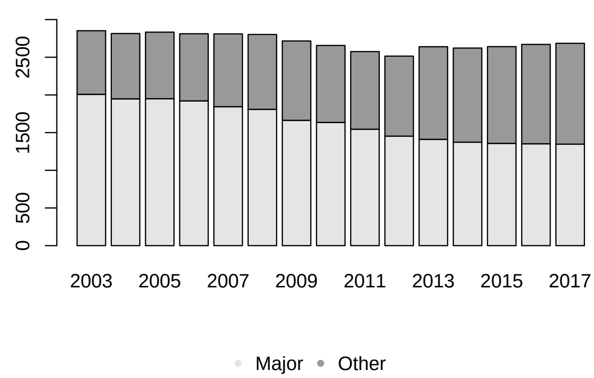

Figure 4: Major-branded and other retail filling stations. Data was obtained from WKO (2017).

4.2.3 Post-regulation subsample

The longer post-regulation subsample is still characterized by positive long-run asymmetry for retail gasoline. However, the speed of adjustment after negative disequilibria has not changed significantly from the between-regulation subsample and the difference between adjustment coefficients remains insignificant. Although we cannot reject the null hypothesis of symmetric cumulative short-run response, we find that the short-run response after input price decreases is faster (two days) than after input price increases (three days). All significant coefficients have the anticipated positive sign. The impulse-response analysis for this period shows that asymmetric short-run and long-run effects cancel out in the first few days after a cost shock. The first 50% of input price changes are passed through in four days independent of the sign of the change. Slightly asymmetric long-run coefficients shape the remaining adjustment process. It takes gasoline 15 days to transmit 90% of an input price decrease while it takes 19 days to transmit 90% of an input price increase. Compared to the results reported in Meyler (2009), it appears that the speed of adjustment has substantially increased. Still, our findings reveal that consumers do not fully benefit from both price regulations in terms of adjustment behavior. The speed of adjustment increases much more after situations in which retail prices are relatively too low. Hence, it is now possible for oil companies to pass through higher gasoline prices faster.

Following the implementation of the second Fuel Price Fixing Act, we reveal statistical evidence for positive long-run asymmetry in the diesel market. While the numerical values of the adjustment coefficients in the between-regulations subsample indicate negative APT, we report positive APT in the longer post-regulation period. Combining short-run and long-run adjustment, we find that it takes diesel 29 days to transmit 90% of a positive cost shock while it only takes 14 days to transmit 90% of a negative cost shock. Price increases are now transmitted faster and price decreases are transmitted slower compared to the pre-regulation period. Consequently, consumers are now faced with less efficient price transmissions.

Figure 2 and Figure 3 show the differences in pass through after the implementation of the Austrian Fuel Price Fixing Act. While the impulse-response curve for retail gasoline becomes slightly steeper in case of negative cost shocks, the curve becomes much steeper in case of positive cost shocks. Similarly, the impulse-response curves for retail diesel have mostly overlapping confidence bands in case of negative cost shocks and are more distinct in case of positive cost shocks. To investigate whether differences in the impulse-response functions between the three regimes are statistically significant, we employ the stationary bootstrap algorithm to bootstrap the differences between impulse-response curves at each step. First, we consider the changes in the impulse-response functions from the pre-regulation to the between-regulation period. While the differences after positive cost shocks are statistically significant at the 5% level for the complete adjustment process of both retail markets, we find significant differences only between steps 4 and 14 (steps 3 to 16) in case of negative cost shocks in the gasoline (diesel) market. The same analysis for post-regulation and pre-regulation curves reveals significant differences after positive cost shocks but not after negative cost shocks. Further, we can only report significant differences for the post-regulation and between-regulation curves in case of negative cost shocks in the retail diesel market. It seems that the first Fuel Price Fixing Act in 2009 mostly affected pass through after positive cost shocks and much more than its amendment in 2011.

Unfortunately, our subsampling methodology is susceptible to 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. Overall, the market structure has remained relatively stable during our sampling period indicated by the amount of retail filling stations depicted in Figure 4. It could be argued that the market power by large oil companies has been slightly reduced as the share of major-branded filling stations has converged to 50% over time and a less concentrated market structure might have an additional positive effect on the efficiency of price transmissions. Since we were not able to obtain high-frequency data on the Austrian fuel market structure, we use a linear trend to account for the decreasing share of major-branded filling stations. As a robustness check, we reestimate our model with a linear trend as a control and interaction term with the error correction terms analogously to the regression models proposed in Oladunjoye (2008). The results obtained for the baseline specification (3) seem to be robust against the possibility of a smoothly transforming market structure.

5 Conclusion and Policy Implications

Our findings in this study shed new light on the implications of price regulations in the Austrian retail fuel market. Our results indicate that the decision to implement fuel price regulations cannot clearly be linked to previously existing APTs in the Austrian retail fuel market. Public concerns about possibly collusive behavior 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 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 input prices are relatively too low. In contrast, we mostly find negative short-run asymmetry which indicates competitive pressure to delay input price increases. Only if we consider the combined effect of cost shocks, we observe that the speed of pass through has become faster for both markets. However, it appears that the speed of pass through improves mainly after input price increases. This implies that consumers benefit less from fuel price regulations than retailers.

Since this study is among the first to evaluate the effects of fuel price regulations empirically and we are limited to data from one country, it remains difficult to find a clear answer whether similar fuel price regulations should be recommended for other countries or whether the current policy should be amended. Subsequent studies might obtain deeper insights using a wider data base which is currently not available. Particularly, we were not able to obtain sufficient data on the changing market structure over our sample period and, hence, we cannot precisely control for this kind of variation. We assume that increasing market shares of independent retail filling stations should foster competition and thus have a positive effect on the efficiency of price transmission. This means that our results might be positively biased in this regard. Future empirical work should be directed at other instances of fuel price regulations to provide a better understanding of how certain aspects of fuel price regulations affect the pricing behavior. For example, the more restrictive Western Australian price regulation should have a stronger effect reducing search costs in the retail fuel market (prices stay fixed for 24h vs. one price increase per day and unlimited price decreases) but would also make it easier for retailers to maintain collusive agreements.

As it stands, we find that the Austrian Fuel Price Fixing Act affected the pricing behavior in various ways. While price volatility is naturally reduced, previous studies find that it also has a significantly negative effect on price levels. However, the effects in terms of efficiency in price transmission are ambiguous. The efficiency measured as the speed of adjustment after input price changes has improved compared to the pre-regulation subsample, but we do not identify substantial positive effects for consumers. Since price regulation policies are commonly introduced to increase competition, it remains doubtful whether the Austrian Fuel Price Fixing Act has reached its intended goals.

References

Arpa, M., Cuaresma, J. C., Gnan, E., Silgoner, M. A., 2006. Oil Price Shock, Energy Prices and Inflation - A Comparison of Austria and the EU. Monetary Policy & the Economy (1), 53–77.

Asane-Otoo, E., Schneider, J., 2015. Retail fuel price adjustment in Germany: A threshold cointegration approach. Energy Policy 78, 1–10.

Bachmeier, L. J., Griffin, J. M., 2003. New evidence on asymmetric gasoline price responses. The Review of Economics and Statistics 85 (3), 772–776.

Bacon, R. W., 1991. Rockets and feathers: the asymmetric speed of adjustment of UK retail gasoline prices to cost changes. Energy Economics 13 (3), 211–218.

Balaguer, J., Ripollés, J., 2012. Testing for price response asymmetries in the Spanish fuel market. New evidence from daily data. Energy Economics 34 (6), 2066–2071.

Balaguer, J., Ripollés, J., 2016. Asymmetric fuel price responses under heterogeneity. Energy Economics 54, 281–290.

Balke, N. S., Brown, S. P. A., Yücel, M. K., 1998. Crude Oil and Gasoline Prices: An Asymmetric Relationship? Economic and Financial Policy Review I, 2–11.

Ball, L., Mankiw, N. G., 1994. Asymmetric Price Adjustment and Economic Fluctuations. The Economic Journal 104 (423), 247–261.

Berninghaus, S., Hesch, M., Hildenbrand, A., 2012. Zur Wirkung regulatorischer Preiseingriffe auf dem Tankstellenmarkt. Wirtschaftsdienst 92 (1), 46–50.

BMDW, 2010. Mitterlehner: Spritpreis-Verordnung schafft bessere Orientierung und mehr Wettbewerb. https://www.bmdw.gv.at/Presse/Archiv/Archiv%202010/Seiten/MitterlehnerSpritpreis-VerordnungschafftbessereOrientierungundmehrWettbewerb.aspx, [Online; accessed 28-December-2017].

Borenstein, S., 1991. Selling costs and switching costs: Explaining retail gasoline margins. RAND Journal of Economics 22 (3), 354–369.

Borenstein, S., Cameron, A. C., Gilbert, R., 1997. Do Gasoline Prices respond asymmetrically to Crude Oil Price Changes. The Quarterly Journal of Economics 112 (1), 305–339.

Borenstein, S., Shepard, A., 2002. Sticky Prices, Inventories, and Market Power in Wholesale Gasoline Markets. The RAND Journal of Economics 33 (1), 116–139.

Bundeskartellamt, 2011. Sektoruntersuchung Kraftstoffe - Abschlussbericht Mai 2011. http://www.bundeskartellamt.de/SharedDocs/Publikation/DE/Sektoruntersuchungen/Sektoruntersuchung%20Kraftstoffe%20-%20Abschlussbericht.html, [Online; accessed 28-December-2017].

Bundeswettbewerbsbehörde, 2008. Untersuchung spezifischer Problemstellungen der Märkte für Mineralölprodukte. https://www.bwb.gv.at/Untersuchungen/Documents/Spritpreisuntersuchung_gesamt.pdf, [Online; accessed 28-December-2017].

Bundeswettbewerbsbehörde, 2009. Treibstoffpreise in Salzburg: Entwicklungen und Einflussfaktoren. https://www.bwb.gv.at/Untersuchungen/Documents/EndberichtTreibstoffmarktSalzburg_20091.pdf, [Online; accessed 11-October-2017].

Byrne, D. P., 2014. Fuelling Australia: Structural Changes and New Policy Challenges in the Petrol Industry. Australian Economic Review 47 (4), 523–539.

Deltas, G., 2008. Retail Gasoline Price Dynamics and Local Market Power. The Journal of Industrial Economics 56 (3), 613–628.

Dewenter, R., Heimeshoff, U., Lüth, H., 2017. Less Pain at the Pump? The Effects of Regulatory Interventions in Retail Gasoline Markets. Applied Economics Quarterly 63 (3), 256–274.

Enders, W., Siklos, P. L., 2001. Cointegration and Threshold Adjustment. Journal of Business & Economic Statistics 19 (2), 166–176.

Engle, R. F., Granger, C. W. J., 1987. Co-Integration and Error Correction: Representation, Estimation and Testing. Econometrica 55 (2), 251–276.

Galeotti, M., Lanza, A., Manera, M., 2003. Rockets and feathers revisited: an international comparison on European gasoline markets. Energy Economics 25 (2), 175–190.

Granger, C. W. J., Lee, T. H., 1989. Investigation of Production, Sales and Inventory Relationships Using Multicointegration and Non-Symmetric Error Correction Models. Journal of Applied Econometrics 4, S145–S159.

Grasso, M., Manera, M., 2007. Asymmetric error correction models for the oil-gasoline price relationship. Energy Policy 35 (1), 156–177.

Haucap, J., Müller, H. C., 2012. The Effects of Gasoline Price Regulations: Experimental Evidence. DICE Discussion Paper (47), 1–25.

IEA, 2014. Energy Policies of IEA Countries - Austria 2014 Review. https://www.iea.org/publications/freepublications/publication/energy-policies-of-iea-countries---austria-2014-review.html, [Online; accessed 11-October-2017].

Johnson, R. N., 2002. Search Costs , Lags and Prices at the Pump. Review of Industrial Organization 20, 33–50.

Kirchgässner, G., Kübler, K., 1992. Symmetric or asymmetric price adjustments in the oil market: An empirical analysis of the relations between international and domestic prices in the Federal Republic of Germany, 1972-1989. Energy Economics 14, 171–185.

Lewis, M., Noel, M., 2011. The Speed of Gasoline Price Response in Markets with and without Edgeworth Cycles. Review of Economics and Statistics 93 (2), 672–682.

Lewis, M. S., 2011. Asymmetric Price Adjustment and Consumer Search: An Examination of the Retail Gasoline Market. Journal of Economics and Management Strategy 20 (2), 409–449.

Lippi, M., 1988. On the dynamic shape of aggregated error correction models. Journal of Economic Dynamics & Control 12, 561–585.

Manning, D. N., 1991. Petrol prices, oil price rises and oil price falls: Some evidence for the UK since 1972. Applied Economics 23 (9), 1535–1541.

Meyler, A., 2009. The pass through of oil prices into euro area consumer liquid fuel prices in an environment of high and volatile oil prices. Energy Economics 31 (6), 867–881.

Obradovits, M., 2014. Austrian-style gasoline price regulation: How it may backfire. International Journal of Industrial Organization 32 (1), 33–45.

OECD, 2013. Competition in road fuel. https://www.oecd.org/competition/CompetitionInRoadFuel.pdf, [Online; accessed 11-October-2017].

Oladunjoye, O., 2008. Market structure and price adjustment in the U.S. wholesale gasoline markets. Energy Economics 30 (3), 937–961.

Perdiguero, J., 2013. Symmetric or asymmetric oil prices? A meta-analysis approach. Energy Policy 57, 389–397.

Politis, D. N., Romano, J. P., 1994. The Stationary Bootstrap. Journal of the American Statistical Association 89 (428), 1303–1313.

Remer, M., 2015. An empirical investigation of the determinants of asymmetric pricing. International Journal of Industrial Organization 42, 46–56.

Schild, K.-H., Schweikert, K., 2019. On the Validity of Tests for Asymmetry in Residual-Based Threshold Cointegration Models. Econometrics 7 (1), 1–13.

Schweikert, K., 2019. Asymmetric price transmission in the US and German fuel markets: a quantile autoregression approach. Empirical Economics 56 (3), 1071–1095.

Sen, A., Clemente, A., Jonker, L., 2011. Retail gasoline price ceilings and regulatory capture: Evidence from Canada. American Law and Economics Review 13 (2), 532–564.

Shepard, A., 1991. Price Discrimination and Retail Configuration. Journal of Political Economy 99 (1), 30–53.

Suvankulov, F., Lau, M. C. K., Ogucu, F., 2012. Price regulation and relative price convergence: Evidence from the retail gasoline market in Canada. Energy Policy 40 (1), 325–334.

Tappata, M., 2009. Rockets and feathers: Understanding asymmetric pricing. RAND Journal of Economics 40 (4), 673–687.

Verlinda, J. A., 2008. Do rockets rise faster and feathers fall slower in an atmosphere of local market power? Evidence from the retail gasoline market. Journal of Industrial Economics 56 (3), 581–612.

Wang, Z., 2009. (Mixed) Strategy in Oligopoly Pricing: Evidence from Gasoline Price Cycles Before and Under a Timing Regulation. Journal of Political Economy 117 (6), 987–1030.

WKO, 2017. Tankstellenstatistik 2010-2017. https://www.wko.at/branchen/handel/energiehandel, [Online; accessed 17-June-2018].

Wlazlowski, S., Giulietti, M., Binner, J., Milas, C., 2009. Price dynamics in European petroleum markets. Energy Economics 31 (1), 99–108.

Wlazlowski, S., Giulietti, M., Binner, J., Milas, C., 2012. Price transmission in the EU wholesale petroleum markets. Journal of Business and Economic Statistics 30 (2), 165–172.

Yang, H., Ye, L., 2008. Search with learning: Understanding asymmetric price adjustments. RAND Journal of Economics 39 (2), 547–564.