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1152

Discussion Papers

Deutsches Institut für Wirtschaftsforschung 2011

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IMPRESSUM

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http://www.diw.de

ISSN print edition 1433-0210
ISSN electronic edition 1619-4535

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Stationarity Changes in Long-Run Fossil Resource Prices: Evidence from Persistence Break Testing

Aleksandar Zaklan[1], Jan Abrell[2] and Anne Neumann[3]

Abstract

This paper considers the question of whether changes in persistence have occurred during the long-run evolution of U.S. prices of the non-renewable energy resources crude oil, natural gas and bituminous coal. Our main contribution is to allow for a structural break when testing for a break in persistence, thus disentangling the effect of a deterministic break from that of a stochastic break and advancing the existing literature on the persistence properties of non-renewable resource prices. The results clearly demonstrate the importance of specifying a structural break when testing for breaks in persistence, whereas our findings are robust to the exact date of the structural break. Our analysis yields that coal and natural gas prices are trend stationary throughout their evolution, while oil prices exhibit a break in persistence during the 1970s. The findings suggest that especially the coal market has remained fundamentals-driven, whereas for the oil market exogenous shocks have become dominant. Thus, our results are consequential for the treatment of energy resource prices in both causal analysis and forecasting.

JEL Codes: C12, C22, Q31, Q41

Keywords: non-renewable resource prices, primary energy, persistence, structural breaks

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We thank Christian von Hirschhausen and Robert Taylor, as well as conference and seminar participants in Dresden, Stockholm and Dallas, for insights and fruitful discussion. We especially thank Eyal Dvir for kindly sharing his code. All remaining errors are ours, the usual disclaimer applies.

1 Introduction

The long-run behavior of non-renewable resource prices has long been a topic of considerable interest; both in the theoretical and in the applied literature (cf. Krautkraemer, 1998, for a literature review). One key property of each price series is its persistence. A series may exhibit persistence of the same type over the entire sample period, e.g. (trend) stationarity, or it may experience a change in persistence, from stationarity to non-stationarity or vice versa. Understanding the character of resource price paths with respect to persistence as well as determining whether a series has experienced changes in persistence is relevant, both from an analytical and a policy perspective.

Determining the persistence properties of a resource price series is one approach to testing the validity of theoretical approaches to modeling resource markets. As Aherns and Sharma (1997) note, (trend) stationarity indicates that resource markets may be mostly driven by market fundamentals (Hotelling, 1931), whereas non-stationarity is consistent with the view that exogenous shocks may dominate (Slade, 1988). Evaluating whether a time series is piecewise stationary may allow us to distinguish periods during which market fundamentals dominated from those in which exogenous shocks may have played a more prominent role.

Furthermore, knowing whether a change in persistence has occurred in energy resource prices will be beneficial for the purposes of inference and forecasting. As Lee et al. (2006) point out a variable's persistence characteristics determine the admissibility of certain estimation frameworks. This consideration extends to changes in persistence as well. When estimating a relationship using a price series that exhibits a break in persistence at some point of its evolution, one may need to consider the pre and post break periods separately. Also, correctly handling the persistence properties of price series can substantially improve the performance of forecasting, as shown by Berck and Roberts (1996) and Lee at al. (2006).

We are particularly interested in analyzing the long-run persistence properties of primary energy commodity prices since bituminous coal, crude oil and natural gas may be considered partial substitutes in electricity generation. Thus, understanding whether the persistence properties of all three fuel prices are similar may shed some light on the degree of their connectedness, i.e. whether these three commodity markets could react differently to shocks caused by the implementation of certain direct or indirect policy options.

There is a substantial empirical literature devoted to the analysis of the persistence properties of resource price time paths. Implicitly assuming trend stationarity, Slade (1982) analyzes the evolution of a number of resource prices and concludes that prices follow a U-shaped time path. Pindyck (1999) analyzes the price paths of bituminous coal, crude oil and natural gas, respectively, and finds a quadratic trend in the data, which is unstable over time. Slade (1988) finds empirical support for prices being non-stationary and concludes that uncertainty appears to be a strong determinant of price formation, as opposed to Hotelling (1931) type deterministic models. Berck and Roberts (1996) also find that resource prices are non-stationary. Ahrens and Sharma (1997) conclude that the evidence is more mixed after analyzing the long-term development of eleven non-renewable resource prices, finding non-stationarity for five price series and trend stationarity for six. Their analysis is partly based on Perron (1989), allowing for a single exogenous structural break per series in 1929, 1939 and 1945, respectively. Using the same data and applying the unit root test by Lee and Strazicich (2003), which allows for up to two endogenously determined structural breaks, Lee et al. (2006) find overwhelming evidence against non-stationarity and in favor of trend stationarity. These studies have in common that they assume that each price series is stable with respect to persistence over its entire sample period.

However, recent developments in persistence testing theory allow us to test the null hypothesis of (trend) stationarity over the entire sample period against the alternative hypothesis that a change in persistence has occurred either from stationarity to non-stationarity or vice versa (Kim, 2000; Kim et al., 2002; Busetti and Taylor, 2004). In addition, the period in which the break has occurred can be estimated. Dvir and Rogoff (2010) apply this methodology to an analysis of long-term crude oil prices. They find that oil prices switch from non-stationarity to stationarity in 1877 and back to non-stationarity in 1973 without allowing for structural breaks. Moreover, we are not aware of a study in the field of resource economics applying the persistence break testing methodology while allowing for a structural break. However, as Perron (1989) shows and as Busetti and Taylor (2004) acknowledge in the context of persistence break testing, an unaccounted-for structural break of a significant magnitude typically biases unit root tests in favor of finding a unit root. In other words, a break in the deterministic trend causes the unit root test to erroneously conclude that a series contains a stochastic trend.

Our contribution to the literature is to allow for a structural break when testing for a break in persistence, thus aiming to disentangle the effect of a deterministic break from that of a stochastic break and adjusting for potential biases from disregarding a possible structural break. A range of potential structural break years is chosen from the existing literature (Perron, 1989; Ahrens and Sharma, 1997; Lee et al., 2006).

Using annual U.S. price data from the 19th century to the early 21st century we first analyze the three price series without allowing for a structural break. In this case we find that all three series exhibit changes from trend stationarity to non-stationarity. However, once we allow for structural breaks, our results diverge for the three series. We find that bituminous coal and natural gas prices are trend stationary throughout their sample periods. However, crude oil prices still exhibit a break from trend stationarity to non-stationarity, although the result is considerably weaker than in Dvir and Rogoff (2010). In our analysis, the persistence breaks for the case of crude oil are all estimated to have occurred during the 1970s, either during the first or the second oil price crisis, depending on the structural break year chosen. Overall, our results are fairly robust to the choice of the structural break period. Thus, we demonstrate that specifying a structural break at all appears to be more important than doing so at a specific point in time.

The remainder of this article is structured as follows: Section 2 provides a descriptive analysis of the data, while section 3 introduces the persistence break testing methodology. We present our results in Section 4 and discuss them in Section 5. The final section summarizes and concludes.

2 Data

Descriptive statistics on the three annual price series from the U.S. are presented in Table I. While the coverage for bituminous coal and crude oil prices is comparable[4], the sample period of natural gas prices is considerably shorter. Furthermore, we notice that oil prices exhibit the greatest variation around their mean, natural gas prices being an intermediate case and coal prices being the most stable. We deflated all three price series using the U.S. CPI, as in Hamilton (2011).

Table I: Real oil, coal and natural gas prices: descriptive statistics

Variable	Obs.	From	To	Mean	SD	Min	Max
Real bituminous coal price per ton	140	1882	2009	5.76	2.10	3.40	11.60
Real crude oil price per barrel	128	1870	2009	25.73	18.35	9.15	96.91
Real natural gas price per barrel	88	1922	2009	12.75	11.04	3.19	47.57

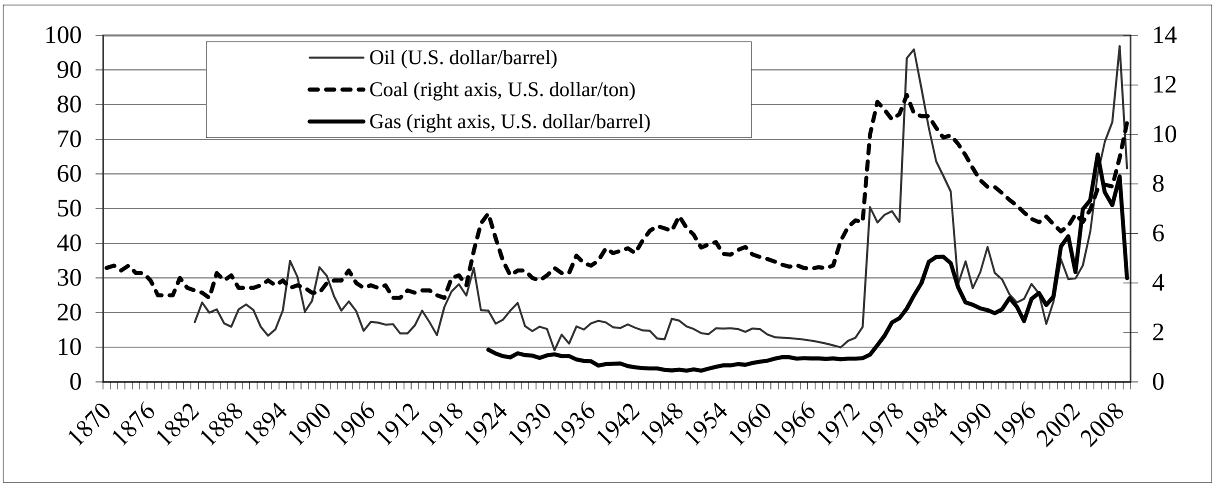
Sources: Bituminous coal prices are from Manthy (1978) and from the U.S. Energy Information Administration (EIA), crude oil prices from BP, and natural gas prices from the EIA.
Note: All prices are annual and were deflated using the U.S. CPI index with the basis year 2009.

The evolution of the three prices series over time is depicted in Figure 1. The coal price series remained stable until after World War I and did not stray far from that level over the following half-century. However, two deviations are notable in the period from the early 1920s until the late 1960s. First, there was a sharp although relatively short-lived spike after World War I. Second, there was a more gradual increase starting in the mid-1930s, reaching a peak in 1948, and from there gradually decreasing again, until by the early 1960s it returned

[4] Oil prices are available from 1861. However, in our analysis we use oil price data starting in 1882, as prior to this the oil market was in its infancy and thus in considerable turmoil (Dvir and Rogoff, 2010; Yergin, 1991).

to the level of the late 1930s. A great increase in real terms followed, before the price gradually settled down toward the end of the 20th century, until rising again during the past decade.

Figure 1: Evolution of real crude oil, bituminous coal and natural gas prices



The oil price is also relatively stable from the 1880s until the early 1970s, with peaks at the end of the 19th century, as well as during and after World War I. Strikingly, World War II only had a small impact on the oil price, most likely due to regulation by U.S. state regulatory bodies, mainly the Texas Railroad Commission (Hamilton, 2011) and price controls during World War II (Yergin, 1991). The oil price remained stable in real terms until the two oil crises in the 1970s, when it rose sharply. During the 1980s and 1990s it returned to a level slightly higher than before the crises. As in the case of the coal price, another significant price increase is observed during the past decade. Furthermore, the oil price is characterized by intermediate and low volatility from the late 19th century until the 1970s, respectively, and then by a marked increase in volatility since then (Dvir and Rogoff, 2010). Overall, the variation in oil prices is significantly greater than that of coal prices.

The natural gas price was stable from the 1920s until the early 1970s. After that it mainly followed the price of oil, although it exhibited smaller price swings. We observe similar increases in the natural gas price as in the case of oil both during the 1970s and in the first decade of the end of the 21st century. In terms of variation around its mean, the natural gas price is an intermediate case between the prices of oil and coal.

3 Methodology

We consider the Gaussian unobserved components model, as presented in Busetti and Taylor (2004):

$$y_t = d_t + \mu_t + \varepsilon_t, \quad t = 1, ..., T \tag{1}$$

$$\mu_t = \mu_{t-1} + 1_{(t > \lfloor \tau T \rfloor)} \eta_t \tag{2}$$

where $\varepsilon_t \sim N(0, \sigma^2)$ and $\eta_t \sim N(0, \sigma_\eta^2 \sigma^2)$ are mutually independent IID processes and $\tau \in]0, 1[$. $1_{(.)}$ is an indicator function taking on the value of 1 for $t > \lfloor \tau T \rfloor$. Thus, starting at point $\lfloor \tau T \rfloor$, η_t from (2) affects (1). Suppose that the point $\lfloor \tau T \rfloor$ is known. Following Busetti and Taylor (2004) we set $d_t = d$ and the starting value $\mu_0 = 0$, without loss of generality.[5] Plugging in for μ_t in (1) yields

$$y_t = d + \varepsilon_t \quad \forall\, t \le \lfloor \tau T \rfloor \tag{3}$$

We thus obtain a stationary process. We then substitute for μ_t:

$$\begin{aligned} \mu_t &= \mu_{t-1} + \eta_t \\ &= \mu_{t-2} + \eta_{t-1} + \eta_t \\ &\vdots \\ &= \sum_{i=1}^{t} \eta_i \end{aligned}$$

Using this result in (1) yields

$$y_t = d + \sum_{i=\lfloor \tau T \rfloor}^{t} \eta_i + \varepsilon_t \quad \forall t > \lfloor \tau T \rfloor \tag{4}$$

This represents a non-stationary process if η_t has non-zero variance. In this manner the summation of a (trend) stationary process yields a non-stationary one (Cavaliere and Taylor, 2008). Therefore, in order to determine whether y_t is stationary throughout we must test

[5] In our actual analysis we allow for both an intercept and a trend.

whether the variance of the η_t process is different from zero, leading to the following null and alternative hypotheses (Busetti and Taylor, 2004):

$$H_0 : \sigma_\eta^2 = 0 \quad \forall t$$

$$\begin{aligned} H_1^a : \sigma_\eta^2 &= 0 \quad for\ t \le \lfloor \tau T \rfloor \\ \sigma_\eta^2 &> 0 \quad for\ t > \lfloor \tau T \rfloor \end{aligned}$$

Thus, H_1^a posits that the series is I(0) until $t = \lfloor \tau T \rfloor$ and I(1) thereafter. We next consider an alternative case for μ_t, in contrast to the representation in (2)

$$\mu_t = \mu_{t-1} + 1_{(t \le \lfloor \tau T \rfloor)} \eta_t \qquad (5)$$

Our null hypothesis remains the same, while the alternative hypothesis changes as follows:

$$\begin{aligned} H_1^b : \sigma_\eta^2 &> 0 \quad for\ t \le \lfloor \tau T \rfloor \\ \sigma_\eta^2 &= 0 \quad for\ t > \lfloor \tau T \rfloor \end{aligned}$$

Thus, we are now testing the null hypothesis of constant stationarity against an alternative hypothesis of I(1) behavior up until the break point and I(0) behavior afterwards. Finally, we can also combine both approaches and test the null hypothesis of constant I(0) behavior against an alternative hypothesis of a change in persistence in either direction, as follows:

$$\begin{aligned} H_1^c : \sigma_\eta^2 &> 0 \quad for \quad t \le \lfloor \tau T \rfloor \\ \sigma_\eta^2 &= 0 \quad for \quad t > \lfloor \tau T \rfloor \\ &OR \\ \sigma_\eta^2 &= 0 \quad for \quad t \le \lfloor \tau T \rfloor \\ \sigma_\eta^2 &> 0 \quad for \quad t > \lfloor \tau T \rfloor \end{aligned}$$

In the following we lay out a procedure for testing H_0 against H_1^a, H_1^b and H_1^c, respectively. Kim (2000) develops a ratio-based statistic to test H_0 against H_1^a. We first define the following partial sum process S_t, as in Kim (2000) and Kim et al. (2002):

$$S_t = \sum_i^t \tilde{\varepsilon}_i, \quad t = 1, \ldots, T \qquad (6)$$

where $\tilde{\varepsilon}_t,\ t = 1, \ldots, T$ are ordinary least square (OLS) residuals from a regression of a time series on an intercept and a trend. Suppose that we wish to test whether a change in persistence has occurred at a specific point $t = \lfloor \tau T \rfloor,\ \tau \in]0, 1[$.[6] We next consider the process before and after $\lfloor \tau T \rfloor$:

$$S_{0,\,t}(\tau_0) = \sum_{i=1}^{\lfloor \tau_0 T \rfloor} \hat{\varepsilon}_{0,\,i}, \quad t = 1, ..., \lfloor \tau T \rfloor \tag{7}$$

$$S_{1,\,t}(\tau_0) = \sum_{i=\lfloor \tau_0 T \rfloor + 1}^{t} \hat{\varepsilon}_{1,\,i}, \quad t = \lfloor \tau T \rfloor + 1, ..., T, \tag{8}$$

where $\hat{\varepsilon}_{0,\,t},\ t = 1, ..., \lfloor \tau T \rfloor$ and $\hat{\varepsilon}_{1,\,t},\ t = \lfloor \tau T \rfloor + 1, ..., T$ are OLS residuals from a regression of y_t on intercept and trend for the periods before and after the proposed break, respectively. We thus obtain the components necessary for computing the test statistic:

$$K(\tau) = \frac{[(1-\tau)T]^{-2} \sum_{i=\lfloor \tau_0 T \rfloor + 1}^{T} S_{1,\,i}(\tau)^2}{[\tau T]^{-2} \sum_{i=1}^{\lfloor \tau_0 T \rfloor} S_{0,\,i}(\tau)^2} \tag{9}$$

For large values of the test statistic we reject H_0 in favor of H_1^a.

Busetti and Taylor (2004) show that we can use the inverse of Kim's (2000) test statistic $K(\tau)^{-1}$ to test H_0 against H_1^b. Again, we reject H_0 for large values of $K(\tau)^{-1}$. Furthermore, Busetti and Taylor (2004) also show that we can test H_0 against H_1^c by means of the following maximum statistic:

$$K(\tau)^{\max} = \max\left\{K(\tau),\ K(\tau)^{-1}\right\} \tag{10}$$

Once again, we reject H_0 in favor of H_1^c for large values of $K(\tau)^{\max}$.

It is not necessary to make any pre-judgement about a potential break point. Rather, we can apply three approaches to computing the test statistics for unknown break points, as suggested by Kim (2000):

[6] For computational reasons τ is chosen such that $\tau \in [\tau_l,\ \tau_u]$ to obtain finite value test statistics (Kim, 2000), with the requirement that the interval be symmetric around τ to ensure consistency of the test statistic (Busetti and Taylor, 2004). We choose $\tau \in [0.1,\ 0.9]$.

$$MX(K(.)) = \max_{\tau \in \mathrm{T}} K(\tau) \tag{11}$$

$$MS(K(.)) = \int_{\tau \in \mathrm{T}} K(\tau) d\tau \tag{12}$$

$$ME(K(.)) = \log \left\{ \int_{\tau \in \mathrm{T}} \exp \left(\frac{1}{2} K(\tau) \right) d\tau \right\}, \tag{13}$$

where $MX(K(.))$ is the maximum over the sequence of statistics for all possible break points (Andrews, 1993), $MS(K(.))$ the mean score statistic (Hansen, 1991) and $ME(K(.))$ the mean-exponential statistic (Andrews and Ploberger, 1994).

Additionally, Kim's (2000) ratio statistic is robust to autocorrelation in the time series under consideration. In order to allow for non-stationarity volatility of a very general form, we employ a wild bootstrap procedure following Cavaliere and Taylor (2008), using 10,000 re-samplings.

In case we reject H_0, i.e. if we find evidence in favor of the series containing a break in persistence, we estimate the period in which the break occurs by determining τ^* in the following criterion, as suggested by Kim (2000):

$$\Lambda(\tau) = \frac{[(1-\tau)T]^{-2} \sum_{i=\tau T+1}^{T} \widehat{\varepsilon}_{1,\, i}(\tau)^2}{[\tau T]^{-2} \sum_{i=1}^{\tau T} \widehat{\varepsilon}_{0,\, i}(\tau)^2}, \tag{14}$$

where τ^* is chosen such that $\tau^* = \arg\max_{\tau \in [\tau l, \tau u]} \Lambda(\tau)$.

Finally, as shown in Busetti and Taylor (2004) the presence of a structural break can seriously distort the test statistic. Perron (1989) and Lee and Strazicich (2003) also demonstrate the existence of this problem in the context of other unit root tests. For this reason we also conduct versions of the test allowing for structural breaks in both level and trend to occur.[7] Since the persistence break literature has yet to develop a way of endogenously determining structural breaks while performing the persistence break test, we

[7] Based on our descriptive analysis of the three commodity price series we hypothesize that at most one structural break has occurred during the sample period for each series.

will treat structural breaks as exogenous and aim to disentangle the effects of structural breaks from those of persistence breaks in this manner. Thus, for the case of a structural break in both the constant and trend our regressor matrix becomes $x_{t\,break} = (1,\ t,\ w_t,\ w_t t)$, where

$$w_t = \begin{cases} 0 & \text{for } t \leq \lfloor \delta T \rfloor \\ 1 & \text{for } t > \lfloor \delta T \rfloor \end{cases}, \tag{15}$$

and $\delta \in]0,1[$.[8] We then perform the persistence break test using structural break points suggested by the existing literature (Perron, 1989; Ahrens and Sharma, 1997; Lee et al., 2006).

4 Results

For conciseness we only report the results based on $K(\tau)$ and $K(\tau)^{-1}$, since the $K(\tau)^{\max}$ statistic unequivocally yields that $K(\tau)$ is always the larger of the two test statistics, for all three commodity price series and for all possible structural break points considered. Thus, our first general finding is that, if anything, all three price series change from I(0) to I(1).

We then first consider the results for bituminous coal prices (Table II). Computing the test statistics without allowing for a structural break leads us to rejecting H_0 in favor of H_1^a, i.e. to finding that the coal price series has become non-stationary. The change in persistence is estimated to have occurred in 1964.[9] However, once we allow for an exogenous structural break in both level and trend our conclusion about the stationarity properties of bituminous coal prices changes. Allowing for a break in 1945 (Ahrens and Sharma, 1997) the result in favor of H_1^a is already significantly weakened.[10] When considering structural breaks in 1902,

[8] We have also performed calculations allowing for breaks in trend and constant in different periods. These are available upon request.

[9] We only estimate a persistence break date if at least two versions of the $K(\tau)$ statistic are significant.

[10] The results for the other cases that Ahrens and Sharma (1997) consider are similar and available upon request.

1915, 1972 and 1973, respectively (Perron, 1989; Lee et al., 2006), we can no longer reject the hypothesis of trend stationarity of coal prices throughout the sample period.[11]

Table II. Testing for change in persistence of bituminous coal prices

Structural Break Type and Year	Mean Score Statistic		Mean-Exponential Statistic		Maximum Statistic	
No Structural Break	MS	18.2224 **	ME	42.1462 ***	MX	93.7435 ***
		(0.030)		(0.007)		(0.007)
	MS^R	0.8269	ME^R	2.14 **	MX^R	12.3016 **
		(0.172)		(0.026)		(0.020)
	Estimated change point I(0) to (1)		1964			
Level and trend break in 1902 Lee et al. (2006)	MS	43.5966 *	ME	51.7712	MX	111.902
		(0.098)		(0.132)		(0.133)
	MS^R	0.2235	ME^R	0.1587	MX^R	2.936 *
		(0.209)		(0.124)		(0.069)
Level and trend break in 1915 Lee et al. (2006)	MS	49.9862	ME	50.9666	MX	109.8864
		(0.104)		(0.175)		(0.178)
	MS^R	0.1904	ME^R	0.1279	MX^R	2.5984 *
		(0.225)		(0.140)		(0.059)
Level and trend break in 1945 Ahrens and Sharma (1997)	MS	18.3296	ME	31.9403 *	MX	73.1709 *
		(0.101)		(0.062)		(0.061)
	MS^R	0.1795	ME^R	0.1081	MX^R	2.4281 *
		(0.668)		(0.577)		(0.072)
	Estimated change point I(0) to (1)		1965			
Level and trend break in 1972 Lee et al. (2006)	MS	3.4708	ME	3.5868	MX	14.9986
		(0.378)		(0.341)		(0.301)
	MS^R	0.502	ME^R	0.2656	MX^R	1.3896
		(0.615)		(0.624)		(0.722)
Level and trend break in 1973 Perron (1989)	MS	3.5195	ME	3.6529	MX	15.1371
		(0.319)		(0.341)		(0.314)
	MS^R	0.4814	ME^R	0.2533	MX^R	1.354
		(0.730)		(0.737)		(0.807)

*, ** and *** indicate significance at the 10%, 5% and 1% levels, respectively. Bootstrap p-values are in parentheses. The bootstrap has been performed using 10,000 samplings.

Table III summarizes the results for crude oil prices. Again, when considering the persistence break test without allowing for a structural break we find strong evidence in favor of a switch from I(0) to I(1). The changepoint is estimated for 1973, as already shown by Dvir and Rogoff (2010). However, contrary to the results for coal prices, allowing for the structural break points identified by the existing literature no longer reverses this result unequivocally. In fact, for most of the suggested structural break points we still find evidence that the series has become non-stationary, although the result is significantly weaker than when ignoring a possible structural break. For all cases in which the persistence break test yields a significant

[11] For the structural breaks from Lee et al. (2006) we only consider the break periods based on their analysis using linear trends.

result the change in persistence is estimated to have occurred during the 1970s, either around the time of the first or the second oil price shocks.

Table III. Testing for change in persistence of crude oil prices

Structural Break Type and Year	Mean Score Statistic		Mean-Exponential Statistic		Maximum Statistic	
No Structural Break	MS	39.3676 ***	ME	201.0695 ***	MX	411.447 ***
		(0.009)		(0.000)		(0.000)
	MS^R	0.9087 *	ME^R	4.7968 ***	MX^R	18.6498 ***
		(0.053)		(0.006)		(0.006)
	Estimated change point I(0) to (1)		1973			
Level and trend break in 1896 Lee et al. (2006)	MS	65.0405 ***	ME	176.1284 ***	MX	361.5646 ***
		(0.007)		(0.007)		(0.007)
	MS^R	0.6898 **	ME^R	2.6517 ***	MX^R	13.9871 ***
		(0.039)		(0.006)		(0.005)
	Estimated change point I(0) to (1)		1972			
Level and trend break in 1945 Ahrens and Sharma (1997)	MS	30.6408	ME	49.5806	MX	108.4637
		(0.188)		(0.179)		(0.177)
	MS^R	0.1745	ME^R	0.1342	MX^R	3.773 **
		(0.404)		(0.209)		(0.015)
Level and trend break in 1971 Lee et al. (2006)	MS	29.7929	ME	74.8542 *	MX	159.0152 *
		(0.123)		(0.100)		(0.100)
	MS^R	0.1132	ME^R	0.0593	MX^R	1.0073
		(0.813)		(0.807)		(0.512)
	Estimated change point I(0) to (1)		1977			
Level and trend break in 1973 Perron (1989)	MS	31.8084	ME	80.6055 *	MX	170.5188 *
		(0.108)		(0.096)		(0.095)
	MS^R	0.091	ME^R	0.0464	MX^R	0.4438
		(0.903)		(0.903)		(0.887)
	Estimated change point I(0) to (1)		1978			

*, ** and *** indicate significance at the 10%, 5% and 1% levels, respectively. Bootstrap p-values are in parentheses. The bootstrap has been performed using 10,000 samplings.

Finally, the results for natural gas are presented in Table IV. Unfortunately, the small sample size does not allow us to perform a full analysis for this case.[12] However, the results again clearly display a discrepancy between allowing for a structural break or not. When not allowing for a structural break we clearly reject trend-stationarity in favor of a change from I(0) to I(1). When allowing for a structural break to occur in 1973 (Perron, 1989; Lee et al., 2006) we again fail to reject trend-stationarity for the entire sample period.

[12] We cannot compute the test statistics using the truncation limits applied to the cases of coal and oil prices. We are forced to truncate the gas price series more severely to $\tau_{gas} \in [0.4, 0.6]$ in order to obtain finite value test statistics, limiting our evaluation of the structural break points proposed by the existing literature. However, fortunately we are still able to evaluate the period of the early 1970s.

Table IV. Testing for change in persistence of natural gas prices

Structural Break Type and Year	Mean Score Statistic		Mean-Exponential Statistic		Maximum Statistic	
No Structural Break	MS	19.2233 **	ME	18.9935 ***	MX	43.8405 ***
		(0.018)		(0.008)		(0.008)
	MS^R	0.0585	ME^R	0.0293	MX^R	0.0741
		(0.987)		(0.987)		(0.997)
	Estimated change point I(0) to (1)		1973			
Level and trend break in 1973	MS	17.8299	ME	17.338	MX	40.5187
Perron (1989), Lee et al. (2006)		(0.724)		(0.609)		(0.602)
	MS^R	0.0634	ME^R	0.0317	MX^R	0.0842
		(0.273)		(0.274)		(0.362)

*, ** and *** indicate significance at the 10%, 5% and 1% levels, respectively. Bootstrap p-values are in parentheses. The bootstrap has been performed using 10,000 samplings.

5 Discussion

Our results clearly demonstrate the importance of allowing for a structural break when testing for a change in persistence, providing empirical support for results in the existing literature (Perron, 1989; Lee and Strazicich, 2003; Busetti and Taylor, 2004).

Concretely, we find that the behavior of coal prices differs from that of oil prices both in terms of finding a persistence break and in terms of the timing of the break when the null hypothesis of trend stationarity is rejected.[13] Once we allow for a structural break in the coal price series we mostly fail to reject the hypothesis of trend stationarity, confirming results from the existing literature (Ahrens and Sharma, 1997; Lee et al., 2006). For the cases in which we do find a persistence break it is estimated to occur in the middle of the 1960s, a period of consolidation in the U.S. coal industry (EIA, 1993). In contrast, for the oil price the evidence is in favor of a change from trend stationarity to non-stationarity, corroborating the finding by Dvir and Rogoff (2010), although this result is considerably weakened once we allow for a structural break. In all significant cases the break is found to occur during the 1970s, indicating that the coal and oil markets may have diverged with respect to persistence during that time. Our descriptive analysis has shown that the natural gas market appears to be similar to the oil market in terms of the direction of price movement, but closer to the coal

[13] Since our analysis of natural gas prices is constrained by data limitations, in our further discussion we will mostly focus on the results for the other two fuels under consideration, bituminous coal and crude oil prices.

market in terms of price variation. Thus, it is plausible that a break in the persistence of natural gas prices may have occurred during the 1970s as well. Our result on natural gas prices is consistent with this assertion when we ignore a possible structural break. However, as for coal prices, this result no longer holds once we account for a possible structural break, suggesting that the natural gas market may be an intermediate case between the coal and oil markets in terms of persistence.

Overall, our results are consistent with the view that the coal market may be predominantly determined by fundamentals, whereas the oil market appears to be more strongly affected by exogenous shocks, with the natural gas market again an intermediate case. Thus, when performing inference using the oil price and when forecasting it may be advantageous to consider the periods before and after the estimated persistence break points separately.

6 Summary and Conclusions

This paper applies recent developments in persistence testing to the question of whether the long-run U.S. prices of the key non-renewable energy resources bituminous coal, crude oil and natural gas exhibit a change in stationarity. We test the hypothesis of trend stationarity over the entire sample period against an alternative of a change in persistence, from trend stationarity to non-stationarity and vice versa. We also estimate the persistence breakpoints. We advance the literature by allowing for a structural break when testing for a change in persistence, thus aiming to avoid a biased test statistic on account of ignoring a potential structural break. To our knowledge this is the first study in the field of resource economics attempting to disentangle a deterministic break from a stochastic break in a price series.

Our findings clearly show the importance of specifying a structural break when evaluating the persistence of a resource price series. When ignoring a structural break the prices of all three resources appear to switch from stationarity to non-stationarity. However, this result is reversed for the cases of bituminous coal and natural gas when allowing for a structural break. Furthermore, while for crude oil prices we still find that they have switched from trend stationarity to non-stationarity in the 1970s, the result is considerably weaker when compared to the case in which we ignore a possible structural break.

Our results indicate a divergence between the markets for crude oil and of those for bituminous coal and natural gas with respect to persistence, at least in a U.S. context, suggesting that oil market analysts may want to take the switch in stationarity into account when estimating relationships in this market and when forecasting oil prices. Our analysis also indicates that a policy intervention may be more promising in the coal market, since it exhibits greater price stability than the oil market, thus facilitating policy targeting.

A fruitful avenue for further research could be to consider whether an analysis of long-run international energy resource prices confirms our results for the U.S. case.

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