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EVIDENCE FROM STATIONARITY TESTS WITH STRUCTURAL CHANGES

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Abstract

This study aims to investigate the efficient market hypothesis for a number of non-renewable resources over the period 1980Q1-2019Q4. We use two different stationarity tests, one is designed to capture smooth breaks, and the other one is designed to detect abrupt changes in the prices. With the use of the stationarity tests, we aim to overcome the low power issue of the commonly utilized unit root tests with stationary but persistent data. Moreover, given the inference of the existing studies for the importance of structural breaks in the analysis of stochastic properties of non-renewable natural resource prices, we utilize both smooth and instant breaks in our analysis to account for the fact that misspecification of the functional form of the breaks could be as problematic as ignoring the breaks. Our empirical results reveal significant evidence of trend stationarity in almost all prices with structural changes related to market-specific and global economic events, though concerns on economic uncertainties appeared to be effective, especially on precious metals. The only exception is silver, with stationarity being rejected for all specifications considered in the paper, suggesting that shocks to silver are mostly permanent in nature and it is characterized by the efficient market hypothesis.

Keywords: Long-run variance estimation; Market efficiency; Non-renewable resource price; Stationarity test; Structural change

JEL classification numbers: C12, C22, Q31

1. Introduction

Price determination has been a great concern in the study of economics since price signals are thought to reflect significant information about markets. The price behavior of exhaustible or non-renewable resources has been discussed for a long time due to a number of reasons, including their scarce nature. Comprehending stochastic structures of commodity prices is essential for commodity-dependent economies to manage and forecast their export earnings. However, it is also crucial for all other economies due to the capacity of commodity prices to influence the current and upcoming production and investment decisions. Moreover, understanding the stochastic properties of commodity prices, especially those of precious metals, is essential for forming and forecasting firms' short-term and long-term investment decisions and diversification strategies. As underlined by Georgiev (2001) and Chan and Young (2006), gold and silver, the most acclaimed precious metals, play a crucial role as a store of value, especially during financial turmoil times, since their inclusion yields risk reduction in portfolios through hedging strategies.

Furthermore, analyzing the price behavior of non-renewable resources conveys significant information about the efficiency of the relevant resource market, which is important for arbitrageurs and speculators. A market is said to be efficient if all relevant and available information is simultaneously reflected in prices so that no participant can earn excess returns consistently by utilizing past, current or new information (Fama, 1970). Depending upon the level of available information, three types of market efficiency are defined. While the weak form efficiency is based on an information set that involves only historical prices, the semi-strong and strong forms, respectively, account for publicly available information and any publicly or privately provided information. In the field of energy, the most commonly examined form of efficiency is the weak form of market efficiency, which states that future energy prices cannot be forecasted by using historical prices. In turn, this suggests that prices follow random walk processes with arbitrary successive price changes. There are many studies available on the empirical validity of the efficient market hypothesis in the field of energy. While a major strand of this literature has focused on the efficiency of oil markets (e.g., Maslyuk and Smyth, 2008; Charles and Darné, 2009; Ozdemir et al., 2013; Stevens and de Lamirande, 2014), the literature on the market efficiency of non-renewable resources is relatively thin. When examining the literature for the market efficiency of non-renewable resources, it appears that earlier studies, including Macdonald and Taylor (1988), Chowdhury (1991), Moore and Cullen (1995), and Berck and Roberts (1996), have employed standard augmented Dickey-Fuller (ADF) and Phillips and Perron (PP) tests. Despite some differences, almost all of these studies concluded that most of the metal prices are non-stationary, and therefore, they can be characterized by the efficient market hypothesis.

Later on, the studies including Ahrens and Sharma (1997), Lee et al. (2006), Kellard and Wohar (2006), Narayan and Liu (2011), Ghoshray (2011), and Presno et al. (2014) have ascribed the nonstationarity finding of the earlier studies to the low power of conventional unit root tests in the presence of structural breaks. Since the seminal paper of Perron (1989), it is well understood that when the time series is stationary around a break, then the conventional unit root tests might produce inferences that are biased towards accepting the false null hypothesis of a unit root. Given that natural resource markets are sensitive to macroeconomic conditions and episodes of world geopolitical tensions, as underlined by Lee and Lee (2009) and Gil-Alana et al. (2015), the use of the conventional unit root tests that ignore potential structural breaks, might produce some serious limitations. In that sense, by imposing exogenously-determined structural breaks associated with the Great Depression in 1929, the outbreak of World War II in 1939 or the end of World War II in 1945, Ahrens and Sharma (1997) investigated stochastic properties of 11 non-renewable natural resource real price series over the period 1870-1990. Their findings indicate that most natural resource prices (6 of 11) appear to be characterized by a trend stationary process with a structural break. Although the assigned break date(s) of Ahrens and Sharma (1997) are quite reasonable economically, it is evident that the applied procedure may subject to a pre-test bias.

Accordingly, Kellard and Wohar (2006) employed the endogenously-determined one-break unit root test of Zivot and Andrews (1992) and the two-break unit root test of Lumsdaine and Papell (1997) to investigate the efficiency hypothesis for several primary commodities, including the non-renewable resources, aluminum, lead, copper, silver, tin, and zinc. Kellard and Wohar (2006) observed strong evidence for stationarity of lead, silver, and zinc around a deterministic trend displaying sharp changes over the period 1900-1998. Although the structural break unit root tests of Zivot and Andrews (1992) and Lumsdaine and Papell (1997) are more reliable if the data is subject to some sudden breaks, it is central to note that they allow for structural breaks only under the alternative of stationarity. Due to this setup, these tests might suffer from serious size distortion problems when the actual data contains a unit root with a structural break, which results in spurious rejections of the null hypothesis of a unit root (Lee and Strazicich, 2003). In that sense, Lee et al. (2006) employed the endogenously-determined two-break unit root test of Lee and Strazicich (2003), which does not suffer from the spurious rejection of the null hypothesis due to allowing for structural breaks under the null of a unit root. By re-analyzing the data of Ahrens and Sharma (1997), Lee et al. (2006) observed that all real price series follow a stationary process around deterministic trends with sharp structural breaks over the period 1870-1990, contrary to the findings of Ahrens and Sharma (1997). Similar to Lee et al. (2006), Ghoshray (2011) utilized the structural break unit root test of Lee and Strazicich (2003) to investigate stochastic properties of 24 diversified primary commodity prices, including those of non-renewable natural resources, aluminum, copper, tin, silver, lead, and zinc. By using annual data over the period 1900 and 2003, Ghoshray (2011) observed empirical evidence for stationarity of copper, tin, and lead, while silver and aluminum prices are found to be non-stationary.

Another study that has utilized the dataset of Ahrens and Sharma (1997) is the one by Presno et al. (2014). Unlike the approach of Lee et al. (2006), Presno et al. (2014) followed a two-step testing procedure with the focus being on the circularity problem between tests for structural breaks and stationarity or unit root behavior of the process. In the first step, they specified the structural change(s) by utilizing a procedure that assumes no prior knowledge regarding the integration order of the process. Once the dates of the structural changes are observed, they continued with testing the null hypothesis of trend stationarity against the alternative of a unit root. In that respect, they applied two types of tests, one allowing for sharp structural breaks and the other one allowing for gradual breaks through the use of a nonlinear logistic function where the transition variable is time. They observed that all real prices except silver and natural gas follow a stationary path with sharp or smooth changes in trend depending on the price analyzed.

Unlike Ahrens and Sharma (1997), Kellard and Wohar (2006), Lee et al. (2006), Ghoshray (2011), and Presno et al. (2014), Narayan and Liu (2011), and Gil-Alana et al. (2015) investigated the efficient market hypothesis over a more recent dataset. In that respect, Gil-Alana et al. (2015) analyzed the stochastic properties of five major precious metal prices, gold, silver, rhodium, palladium, and platinum, by using monthly data from 1972:1 to 2003:12. They used a fractional integration framework enhanced with the identification of sudden structural breaks and observed that even if the prices show a mean-reverting structure at the beginning of the sample, the degree of persistence of almost all prices increases across time. Narayan and Liu (2011) analyzed the stochastic properties of daily prices of 10 non-renewable natural resources for the period ending in March 2010. They employed two different unit root tests, one allowing for two sharp structural breaks in intercept and trend terms and the other one accounting for both the structural breaks and the potential autoregressive conditional heteroscedasticity (ARCH) structure, which is necessary to address the heteroscedasticity issue that emerges due to the use of daily data. Their findings reveal that while the unit root null hypothesis can be rejected for most of the prices, a number of series, including gold, silver, platinum, aluminum, and copper, follow a random walk structure. Some of the other studies utilizing daily frequency to investigate the efficient market hypothesis are the ones by Cagli et al. (2019), Pathak et al. (2020), and Rehman and Vo (2021).

For the period January 1985 to February 2019, Cagli et al. (2019) observed that both the spot and future prices of all analyzed industrial and precious metals show a non-stationary structure through a unit root test designed to capture nonlinear properties of the prices in the presence of heteroscedastic variances. Using daily data ranging from July 31, 2000 to July 31, 2020, Rehman and Vo (2021) reported that all metal prices follow a random walk structure based on a structural break unit root test. Similar to the nonstationarity findings of Cagli et al. (2019) and Rehman and Vo (2021), employing the generalized spectral test for daily prices of precious metals, gold, palladium, and platinum, Pathak et al. (2020) reported that markets for these metals are efficient for most times of the period 2009-2018 despite disrupting effect of the crises on the level of efficiency. As underlined by Goshray and Johnson (2010), the common nonstationarity finding of the studies employing higher frequency data might be because of their analysis resting on relatively shorter periods, dominated by the persistence of shocks to metal prices.

In this paper, we intend to investigate whether quarterly real prices of copper, lead, tin, nickel, zinc, aluminum, gold, platinum, and silver can be characterized by the efficient market hypothesis or not over the period 1980Q1 and 2019Q4. Following the existing literature, real metal prices are selected to eliminate the potential cyclicality of the exchange rate. Methodologically, unlike the majority of the current studies, we utilize two different stationarity tests, which are improved versions of the conventional KPSS (Kwiatkowski et al., 1992) test for structural breaks. Since tests with the null of a unit root produce low power when applied to stationary but persistent data, these tests cannot reject the null hypothesis of nonstationarity unless there is very strong evidence against it. Therefore, the market efficiency hypothesis can be more naturally tested under the null of stationarity, as in Presno et al. (2014). Moreover, following the proposal of Lee et al. (2006) that structural breaks and trends are important considerations for the analysis of stochastic properties of non-renewable natural resource prices, we adopt two different tests, one is designed to capture smooth breaks, and the other one is designed to detect abrupt changes in trend. Given that the consequences of misspecification of the functional form of the breaks are as severe as those of ignoring the breaks, we consider both smooth and instant breaks in our analysis.

More specifically, we first use the endogenously-determined two-break stationarity test of Carrion-i Silvestre and Sanso (2007) to detect potential sharp breaks in the price series. Then, accounting for the possibility that structural changes might take a period of time to affect an economy and might not be captured well by dummy variables, we utilize the stationarity test of Becker et al. (2006), designed to detect multiple smooth breaks of unknown form through the use of a Fourier function. The advantage of using a Fourier function over the nonlinear logistic function used by Presno et al. (2014) is that the Fourier approximation can also detect u-shaped breaks and smooth breaks located near the end of the series. Both approaches of Carrion-i Silvestre and Sanso (2007) and Becker et al. (2006) are modified versions of the standard KPSS test. Given the sensitivity of the KPSS test to the estimation of the long-run variance, we provide careful consideration of this drawback throughout our analysis.

The rest of the study is organized as follows. The next section describes the econometric methodologies we adopt. While Section 3 discusses the data and the empirical results, Section 4 concludes the study.

2. Methodology

This section describes first the sharp break stationarity test of Carrion-i Silvestre and Sanso (2007) and then continues with the stationarity test of Becker et al. (2006) that allows for smooth breaks of unknown form.

2.1 Sharp Break Stationarity Test of Carrion-i-Silvestre and Sanso (2007)

To test the null hypothesis of stationarity in the presence of sharp structural breaks, Carrion-i Silvestre and Sanso (2007) considers the following setting for the time series process y_t :

$$\begin{aligned} y_t &= f\left(t, T_{b1}, T_{b2}\right) + r_t + \varepsilon_t \\ f\left(t, T_{b1}, T_{b2}\right) &= \theta_0 + \gamma_0 t + \sum_{i=1}^{2} \theta_i DU_{i,t} + \sum_{i=1}^{2} \gamma_i DT_{i,t} \\ r_t &= r_{t-1} + u_t \end{aligned} \tag{1}$$

where ε_t is the stationary error term, u_t is $iid\left(0, \sigma_u^2\right)$, $DU_{1,t}$ and $DU_{2,t}$ are the dummy variables for mean shifts and $DT_{1,t}$ and $DT_{2,t}$ are the dummies for trend shifts which occur at times $T_{b1} = \lambda_1 T$ and $T_{b2} = \lambda_2 T$, where $\lambda_1, \lambda_2 \in (0,1)$ and $T_{b1} \neq T_{b2} \mp 1$. That is:

$$DU_{i,t} = \begin{cases} 1 & if\ \ t > T_{bi} \\ 0 & otherwise \end{cases} \quad \text{and} \quad DT_{i,t} = \begin{cases} t - T_{bi} & if\ \ t > T_{bi} \\ 0 & otherwise \end{cases}$$

Given the regression, the KPSS test statistic for the null of trend stationarity, i.e. $\sigma_u^2 = 0$, has the form:

$$\hat{\eta} = \hat{\sigma}^2 T^{-2} \sum_{t=1}^{T} S_t^{\,2} \tag{2}$$

where $S_t = \sum_{i=1}^{t} \hat{e}_i^{\,2}$ with $\hat{e}_i$ being the ordinary least square (OLS) residual obtained from the regression of y_t on the intercept and trend terms and $\hat{\sigma}^2$ being the estimated long-run error variance[1].

In this setup, one needs to clarify two unknown points, break dates, and estimation of the long-run variance $\hat{\sigma}^2$, which is expected to capture the unknown autocorrelated structure. Carrion-i Silvestre and Sanso (2007) suggest estimating the unknown break dates through the minimization of the sequence of the sum of squared residuals (SSR). That is, the regression (1) is estimated by OLS for each potential structural break T_{b1} and T_{b2}, and then the dates which produce the minimum SSR are selected. More specifically, the break dates are estimated as:

$$\left(\hat{T}_{b1}, \hat{T}_{b2}\right) = \underset{\lambda_1, \lambda_2 \in \Lambda}{\arg\min}\, SSR\left(T_{b1}, T_{b2}\right)$$

where the interval Λ is set as $\Lambda = \left[\frac{2}{T}, \frac{T-1}{T}\right]$.

The estimation of the long-run variance, which is the main drawback of stationarity tests is the next problem that needs to be accommodated to calculate the KPSS test statistic. In their original paper, KPSS has proposed a nonparametric estimator of $\hat{\sigma}^2$ in the form of

$$\hat{\sigma}^2 = T^{-1}\sum_{t=1}^{T} \hat{e}_t^{\,2} + 2T^{-1}\sum_{s=1}^{l} w(s,l) \sum_{t=s+1}^{T} \hat{e}_t \hat{e}_{t-s}$$

where $\hat{e}_t$ is the residual obtained from the OLS estimation of the regression of y_t on the deterministic terms and $w(s,l)$ is the Bartlett kernel set with a truncation lag

[1] In their papers Carrion-i Silvestre and Sanso (2007) have proposed seven different specificants for the deterministic terms that are represented by $f\left(t, T_{b1}, T_{b2}\right)$. Among them, we select the most general form that allows for two structural changes in the intercept and trend terms. For other deterministic specifications, see Carrion-i Silvestre and Sanso (2007).

$l = \text{integer}\left[q\left(T/100\right)^{1/4} \right]$ and with $q = 0,\ 4$ or 12. Although Bartlett kernel is used by KPSS to weigh the estimated autocovariances, different kernels, like Parzen and Quadratic Spectral kernel, can be also applied. Among all available kernels, the Quadratic Spectral kernel appears to be the most preferred one in the literature due to Andrews (1991) and Newey and West (1994), who have shown that the Quadratic Spectral kernel has the optimal asymptotic mean squared error properties and yields more accurate long-run variance estimates than other kernels in finite samples. The other determinant of the estimation of the long-run variance is the truncation lag l and its calculation has received more attention in the literature. Regarding the optimal truncation lag, there are different suggestions. In that respect, while Lee (1996) suggested using Andrews' (1991) method, Hobijn et al. (2004) preferred to employ the method proposed by Newey and West (1994). However, Choi (1994), Choi and Ahn (1995, 1999), Kurozumi (2002) and Sul et al. (2005) have criticized the data-based selection methods of Andrews (1991) and Newey and West (1994) for leading to inconsistency of the test under random walk alternatives and advised to eliminate the inconsistency by imposing some bounds to control the estimated truncated lag. Recently, Carrion-i-Silvestre and Sanso (2006) provided a comparative analysis to investigate the finite sample properties of the KPSS test under different estimation methods for the long-run variance. According to their Monte Carlo analysis, the procedure suggested by Sul et al. (2005) appears to be the best one with less size distortion and reasonable power. Therefore, Carrion-i Silvestre and Sanso (2007) have advised using the prewhitened heteroscedasticity and autocorrelation consistent (HAC) estimator of Sul et al. (2005) for the long-run variance in their two-break KPSS test.

The procedure suggested by Sul et al. (2005) is as follows: First, an autoregressive (AR) model is estimated for the residuals $\hat{e}_t$ as:

$$\hat{e}_t = \rho_1 \hat{e}_{t-1} + \ldots.. + \rho_p \hat{e}_{t-p} + \psi_t \tag{3}$$

where the appropriate lag order can be specified for example using the Bayesian information criterion (BIC). Once the AR model is estimated, the long-run variance of the estimated residuals $\tilde{\sigma}_{\psi}{}^{2}$ is obtained through the use of a HAC estimator by using the Quadratic Spectral window to take the presence of heteroscedasticity into account. In the second step, the estimated long-run variance $\tilde{\sigma}_{\psi}{}^{2}$ is recolored as:

$$\hat{\sigma}^{2} = \frac{\tilde{\sigma}_{\psi}{}^{2}}{\tilde{\rho}(1)^{2}}$$

where $\tilde{\rho}(1)$ represents the autoregressive polynomial $\tilde{\rho}(L) = 1 - \tilde{\rho}_{1}L - ..... - \tilde{\rho}_{p}L^{P}$ evaluated at $L = 1$. Finally, in order to eliminate the inconsistency of the KPSS test statistic that is emerged due to the use of prewhitened long-run variance estimates, Sul et al. (2005) set the following boundary condition:

$$\hat{\sigma}^{2} = \min\left\{ T\tilde{\sigma}_{\psi}{}^{2}, \frac{\tilde{\sigma}_{\psi}{}^{2}}{\tilde{\rho}(1)^{2}} \right\} \tag{4}$$

Once the estimation of break dates and long-run variance are handled, Carrion-i Silvestre and Sanso (2007) derived the nonstandard asymptotic distribution of the KPSS-type statistic $\hat{\eta}$, which depends on the relative positions of the breaks in the sample. Hence, they estimated finite sample critical values at the 5 percent significance level using response surface regressions, where the critical values converge to their asymptotic values as the sample size increases.

2.2 Smooth Break Stationarity Test of Becker et al. (2006)

Different from the sharp break stationarity test of Carrion-i Silvestre and Sanso (2007), which is based on the implicit assumption that breaks occur at a particular point in time and their effects are felt instantaneously, the test of Becker et al. (2006) considers the possibility that structural changes can occur gradually. It is designed to capture multiple smooth breaks in the series through the use of a Fourier function, which can approximate any integrable functions to any desired degree of accuracy. In this approach, the time-series process y_t is modeled as:

$$\begin{aligned} y_t &= \alpha + \beta t + \alpha(t) + r_t + \varepsilon_t \\ \alpha(t) &= \gamma_1 \sin\left(\frac{2\pi kt}{T}\right) + \gamma_2 \cos\left(\frac{2\pi kt}{T}\right) \\ r_t &= r_{t-1} + u_t \end{aligned} \tag{5}$$

where $\alpha(t)$ represents the time-varying deterministic component that requires no prior knowledge regarding the number and forms of breaks, k is the frequency selected for the approximation, γ_1 and γ_2 measure the amplitude and displacement of the frequency component, t is a trend term, T is the sample size, $\pi = 3.1416$, ε_t is a stationary disturbance term and u_t is the iid disturbance term with the variance $\sigma_u{}^2$. In this setting, it is obvious that under the null hypothesis $\sigma_u^2 = 0$, y_t is trend stationary and the KPSS statistic has the form:

$$\tau_\tau(k) = \hat{\sigma}^2 T^{-2} \sum_{t=1}^{T} \tilde{S}_t(k)^2 \tag{6}$$

where $\hat{\sigma}^2$ being the estimated long-run error variance and $\tilde{S}_t(k) = \sum_{i=1}^{t} \tilde{e}_i{}^2$ with $\tilde{e}_i$ being the OLS residual obtained from the regression of

$$y_t = \alpha + \beta t + \gamma_1 \sin\left(\frac{2\pi kt}{T}\right) + \gamma_2 \cos\left(\frac{2\pi kt}{T}\right) + e_t$$

The calculation of the KPSS statistic requires estimated long-run variance and an appropriate frequency k. Although Becker et al. (2006) followed the approach of KPSS to estimate the long-run variance, in our analysis we prefer to use the approach of Sul et al. (2005) due to its previously discussed merits in long-run variance estimation. Given that the true value of k is unknown, the next issue we need to clarify is finding the appropriate frequency k. To do that, Becker et al. (2006) advise employing a grid search procedure. That is, equation (5) is estimated through OLS for each integer value of $k \in [1,5]$ and the optimum frequency that produces the smallest residual sum of squares is selected. Becker et al. (2006) showed that the asymptotic distribution of the proposed KPSS statistic $\tau_{\tau}(\tilde{k})$, where $\tilde{k}$ is the optimum frequency observed from the grid search procedure, is non-standard and depends on the frequency of the Fourier series. Hence, for different integer values of k, they tabulated the critical values through Monte Carlo simulations.

If the application of the test provides empirical support for the stationarity of y_t, then Becker et al. (2006) suggest continuing further with testing for the statistical significance of smooth breaks since the null hypothesis of $\sigma_u^2 = 0$ includes no specific assumption regarding the significance of the breaks. In that sense, the null hypothesis $H_0 : \gamma_1 = \gamma_2 = 0$ is tested in model (5) by using an *F*-statistic, $F(\tilde{k})$. Obviously, the distribution of the *F*-test is non-standard due to the presence of the nuisance parameter under the null; therefore, the critical values are simulated through Monte Carlo simulations and tabulated in Becker et al. (2006).

3. Data and Empirical Results

We use quarterly real metal prices of copper, lead, tin, nickel, zinc, aluminum, gold, platinum, and silver covering the period of 1980Q1 and 2019Q4. The real metal prices are calculated by deflating the end-period nominal metal prices by the seasonally adjusted US producer price index (PPI) (1982 =100), as in Ahrens and Sharma (1997), Lee et al. (2006) and Presno et al. (2014). The data is extracted from the World Bank Commodity Price dataset (the Pink Sheet) and the Federal Reserve database (FRED). Before the analysis all real prices are transformed into natural logarithms.

Our empirical analysis commences with the application of the standard KPSS test where the approach of Sul et al. (2005) is utilized for estimation of the long-run variance. According to the results reported in the second column of Table 1, the standard KPSS test provides evidence for stationarity for zinc and aluminum only. As stated earlier, the KPSS test might be seriously misleading when the real price series are subject to structural breaks. Hence, to provide more reliable results, we continue with the stationarity tests of Becker et al. (2006) and Carrion-i Silvestre and Sanso (2007), which account for the presence of potential structural breaks in the price series.

We apply the smooth break stationarity test of Becker et al. (2006) and start with modeling smooth breaks in real metal prices. As such, we estimate the Fourier regression in (5) for each integer frequency, and the optimal frequency $\tilde{k}$ is selected through the minimization of the sum of squared residuals. Since Becker et al. (2006) have proposed that one or two frequencies should be sufficient to capture the important breaks in the series and higher frequencies are likely to be related with stochastic parameter variability rather than structural breaks, we employ the grid search procedure for values $k=1$ and 2[2]. Estimated frequencies $\tilde{k}$ and the KPSS-type test statistics $\tau_{\tau}\left(\tilde{k}\right)$ obtained from (6) are reported in the second and third columns of Table 1. According to the results, the stationarity test of Becker et al. (2006) not only supports the stationary inferences of the standard KPSS test for zinc and aluminum prices but also reveals further evidence for stationarity of lead, tin, and nickel prices. Once stationarity of the prices of lead, tin, nickel, zinc, and aluminum is ensured, we proceed with testing for the statistical significance of the smooth breaks through the $F\left(\tilde{k}\right)$ test. The results reported in the second column of Table 1 reveal the statistical significance of smooth breaks in all cases found

[2] When we extend the grid search procedure for each integer value of $k \in [1,5]$, we observe no substantial change in the reported results.

to be stationary. Hence, we observe more supportive results for stationarity of real metal prices with the consideration of smooth breaks.

We further plot the real metal prices together with the estimated Fourier functions in Figure 1 to provide an illustration of the observed smooth structural breaks. It seems from the Figure that the Fourier approximations are plausible to capture the overall pattern of the real price series, though they do not seem to perform very well in capturing the sharp breaks in terms of break magnitudes and timings. This finding is consistent with the proposal of Jones and Enders (2014) that Fourier approximations can perform plausibly well mimicking sharp breaks though they might struggle with identification of the time and magnitude of the breakpoint. This is an issue that might affect the performance of the stationarity test of Becker et al. (2006) since, as suggested by Harvey and Mills (2004), the size distortions of the smooth break stationarity tests in the presence of instant breaks might worsen with the magnitude of the break.

Table 1: Results of the Stationarity Tests

	KPSS Test	*Stationarity Test of Becker et al. (2006)*			*Stationarity Test of Carrion-i-Silvestre and Sanso (2007)*		
		$\tilde{k}$	$\tau_\tau(\tilde{k})$	$F_\tau(\tilde{k})$	$\hat{\eta}$	$\hat{T}_{b1}$	$\hat{T}_{b2}$
Copper	0.367	2	0.281	na	0.027**	1987Q2	2004Q4
Lead	0.643	1	0.043**	59.298**	0.037**	1986Q3	2003Q4
Tin	2.642	1	0.045**	378.689**	0.017**	1990Q2	2006Q3
Nickel	0.176	2	0.047**	45.505**	0.022**	1987Q3	2003Q3
Zinc	0.100**	2	0.100**	12.028**	0.035**	1988Q1	2004Q4
Aluminum	0.057**	2	0.028**	12.542**	0.067**	1987Q2	1993Q4
Gold	4.093	1	0.212	na	0.038**	2000Q4	2013Q1
Platinum	0.496	2	0.112	na	0.014**	1999Q3	2010Q3
Silver	2.040	2	0.067	na	0.047	1992Q3	2010Q2

Notes: $\tau_\tau(\tilde{k})$ indicates the KPSS-type test statistic of Becker et al. (2006) with $\tilde{k}$ representing the selected frequency of the Fourier function and $F_\tau(\tilde{k})$ being the *F*-statistic to test for the presence of the smooth breaks. The critical values for $\tau_\tau(\tilde{k})$ and $F_\tau(\tilde{k})$ are taken from Table 1 of Becker et al. (2006). na stands for 'Not Applicable' since F-statistic is not reliable when stationarity is rejected. $\hat{\eta}$ is the KPSS-type test statistic of Carrion-i-Silvestre and Sanso (2007) with the observed structural breaks $\hat{T}_{b1}$ and $\hat{T}_{b2}$. The critical values for $\hat{\eta}$ are obtained from the response surfaces in Table 2 of Carrion-i-Silvestre and Sanso (2007). In all tests, the long-run variance is estimated by using the approach of Sul et al. (2005), where the appropriate lag order of the AR regression (3) is chosen using the Bayesian Information Criterion (BIC) with the maximum autoregressive order set to be 8. ** denotes stationarity of the price series at 5 percent significance level.

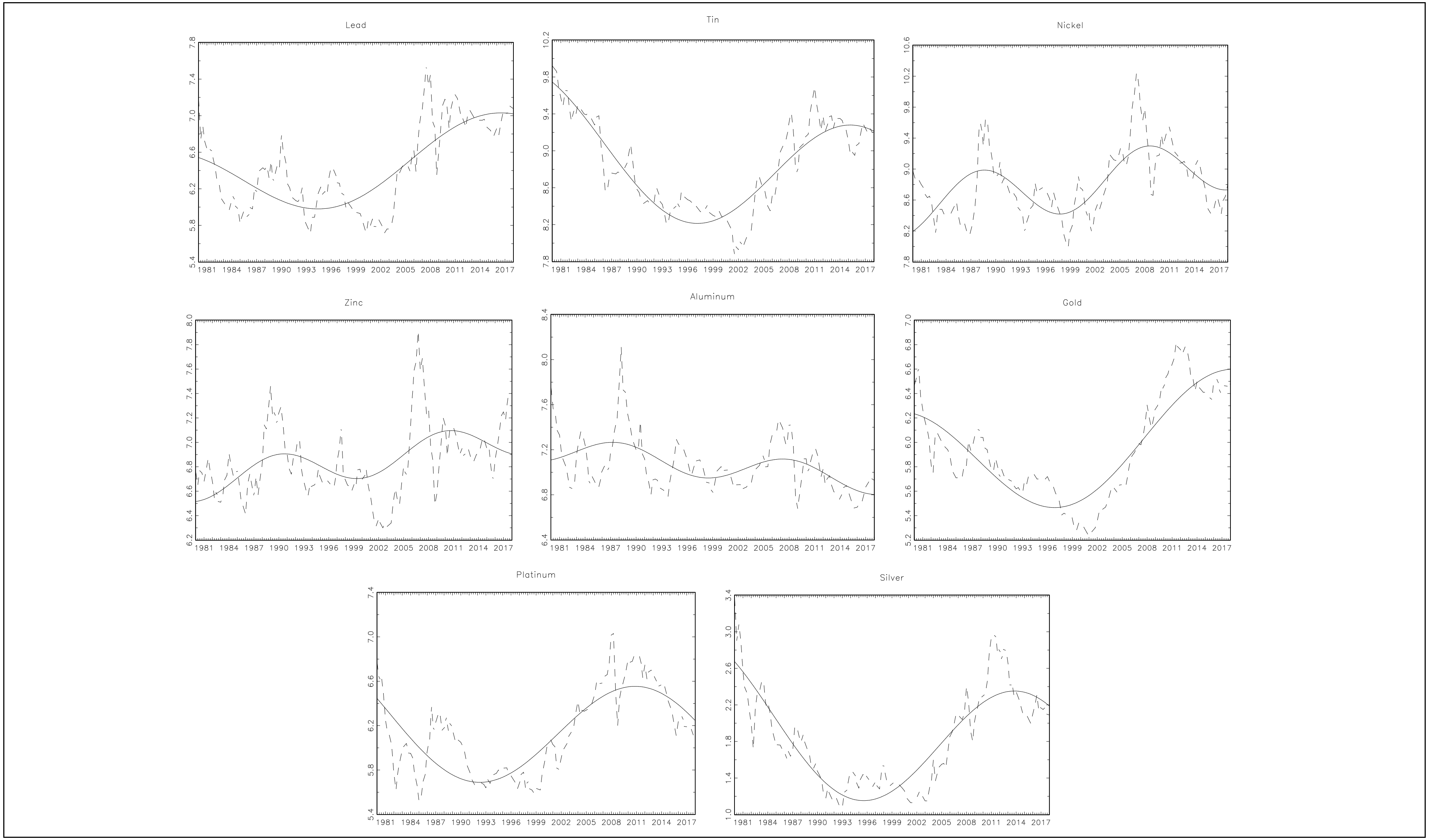

Figure 1: Real Metal Prices with Fitted Fourier Functions

Taking this potential problem into account, we continue with the sharp break stationarity test of Carrion-i Silvestre and Sanso (2007). The test statistic $\hat{\eta}$ calculated from (2) and the corresponding estimated structural break dates are presented in the last three columns of Table 1. It appears from the results that the improvements observed from the test of Becker et al. (2006) become more apparent with the application of the test of Carrion-i Silvestre and Sanso (2007) so that all price series except silver are found to be stationary at the 5 percent significance level. Compared to the most recent studies utilizing low-frequency data, our results, especially those in favor of stationarity of copper, aluminum, platinum, and gold, differ from the outcomes of Ghoshray (2011) and Narayan and Liu (2011), who performed their analysis from the perspective of unit root testing. However, despite the difference in the time span of the data, our empirical findings are entirely in line with the inferences proposed by Presno et al. (2014), who utilized a stationarity testing approach.

The observed structural breaks are then plotted together with the original price series in Figure 2. Overall, it seems that the observed dates and magnitudes of the breaks are more coherent with the original price series compared to those derived from the approach of Becker et al. (2006). This comparative coherence is more apparent for copper, gold, and platinum, which are found to be non-stationary through the use of the test of Becker et al. (2006) and display sharp upward movements after 2004Q4, 2000Q4 and 1999Q3, respectively.

Regarding the observed structural changes for copper, the change points are detected as 1987Q2 and 2004Q4. The first change can be associated with the transformation of the 1980-1986 oversupply and depressed copper price period into the tight supply and high-price period in 1987 due to the highly boosted production resulted from the previous oversupply and low-price period. The second change is quite sharp compared to the first change, and it might be related to the explosive growth in Chinese demand for copper that is coupled with mine production cutbacks. For lead, the break dates appeared as 1986Q3 and 2003Q4. It seems that following the hard times experienced by the lead industry between 1982 and 1986, the change observed in 1986 corresponds to the beginning of the recovery period of the industry due to expansion in demand for lead and to substantial cost-cutting in the primary and secondary producing sectors (USGS, 2012:81). The rapid growth in the lead-acid battery industry for automotive and industrial uses together with the development of emerging economies in Asia, especially China, boosted the demand for lead and constituted the second change in 2003Q4 with substantial price increases. The breaks for tin prices are detected in 1990Q2 and 2006Q3. The year 1990 can be named as the start of the recovery period of the tin industry following the major tin crisis in 1985 and its delisting from trading on the London Metal Exchange for about 3 years. After following a stable path from 1990 through 2006, tin prices began to increase due to several factors including mine disasters and floodings in China, the world's leading tin producers, shortfalls in tin production in Indonesia, the world's second-leading tin producer and substantial growth in world tin consumption (USGS, 2012:181).

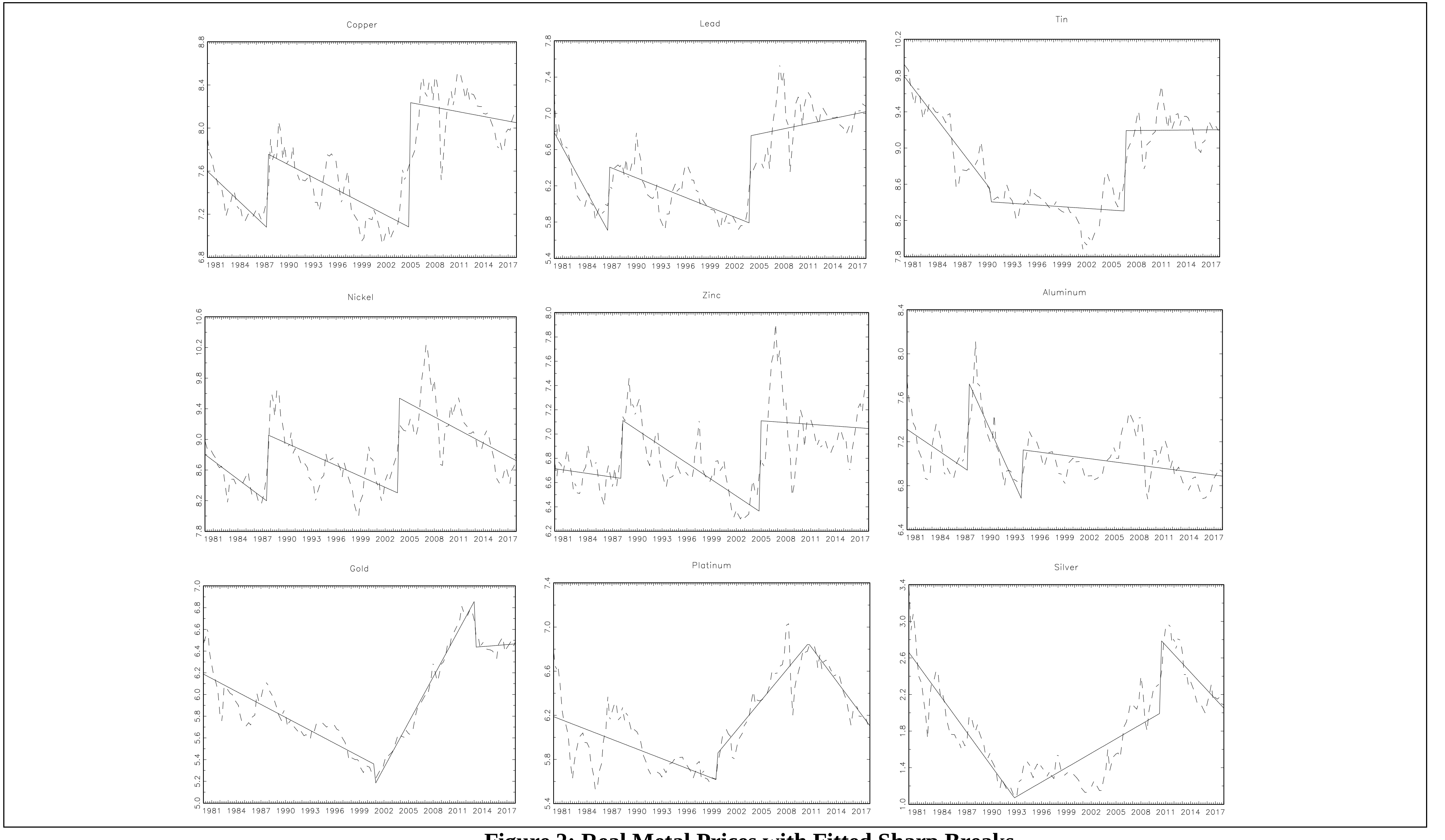

Figure 2: Real Metal Prices with Fitted Sharp Breaks

For the nickel price, the first structural change is observed in 1987Q3 with a sharp increase in prices. The factors that fuelled the observed increase are summarised in Kuck (1999) and USGS (2012:107) as reduced world production of nickel due to its low prices during the early and mid-1980s and a substantial increase in demand for stainless steel, the largest end-use for nickel. The second change in 200Q3 could be related to the increased nickel consumption due to the rapidly expanding economy in China and the increased use of steel worldwide. Moving on to the zinc price, the first break is detected in 1988Q1 when the prices increased sharply due to the factors including tight supply resulted from strong world demand, strikes and hurricane-related delays of zinc shipments from Mexico. After following a decreasing trend triggered by an oversupply in the metal market, which is mainly caused by low global demand and a substantial increase in Chinese exports, in 2004 (the second break) zinc prices shifted upward with increased global demand for zinc. The main impetus for this increase could be the rapidly growing Chinese economy and infrastructure coinciding with constrained zinc production by curtailments and mine closures that were implemented as a result of low prices during the late 1990s and early 2000s. Regarding aluminum prices, after suffering from a period of oversupply during the early 1980s, leading aluminum prices to tumble, the increasing use of aluminum in beverage cans and automobile parts led the worldwide demand for aluminum to raise in 1987 (USGS, 2012:2). The break we observed in 1987Q1 coincides with this tight supply situation and the resultant sharp increase in aluminum prices. In the early 1990s, with the dissolution of the Soviet Union, large quantities of Russian[3] aluminum entered into the world market, which created a period of oversupply, high inventories, and depressing world aluminum prices. Our break date 1993Q4 corresponds to the rebound of the aluminum prices with production cutbacks and increased demand.

[3] Russia was the second largest primary aluminum producer after the U.S.A. in years 1992 and 1993 (Bureau of Mines,1993:107)

For the gold price, after following a decreasing trend from the early 1980s to 2000, the gold price started to increase in 2001 with high investor demand spurred by concerns after September 11 terrorist attacks, lack of sales of central banks, reduced mine production, safe-haven demand and creation of exchange-traded funds[4] which ease small investors to invest in gold (USGS, 2012:60). While the gold price continued to climb and reached its peak level in 2011 with the European sovereign debt crisis, the price downturn in 2013 may be attributed to decreased gold investment coinciding with the improvement of United States' economic fundamentals in the early part of 2013 signaling the reversal of expansionary monetary policy or expectation of U.S. interest rates (Bureau of Mines, 2013:7). Turning to the platinum price, the change points are detected as 1999Q3 and 2010Q3. Despite two moderate peaks reached in 1983 and 1986 due to the increased demand for platinum for use in catalytic converters to control air pollution, the aluminum price fluctuated around a decreasing trend until the end of 1999. From late 1999 through late 2010, platinum prices rose owing to its increased consumption by the automobile industry, the electrochemical and electronics sectors. From late 2010, the prices have started to decrease, probably due to the global uncertainties regarding European sovereign debt problems (Bureau of Mines, 2011:4). Finally, for the silver price, although the turning points, namely peaks and troughs, seem to be captured relatively well by the observed structural change dates, the estimates of break dates are inconsistent due to the nonstationarity of the silver prices and should not be interpreted, as underlined by Bai (1994). Our conclusion about nonstationarity of silver prices is in line with the inferences proposed by the most recent studies of Goshray (2011), Narayan and Liu (2011), and Presno et al. (2014), though the embodied methodologies and the time span of the data differ.

Overall, our results indicate that while industrial metals are affected mainly by market-specific conditions, precious metals are driven specifically by global macroeconomic conditions. This finding is quite revealing given that precious metals play an important role as

[4] The first exchange traded fund backed by gold is established in 2003.

a store of value especially in times of financial turmoil since their inclusion yields risk reduction in portfolios through hedging strategies.

4. Conclusion

This study has scrutinized the efficient market hypothesis for a number of non-renewable resources, including copper, lead, tin, nickel, zinc, aluminum, gold, platinum, and silver over the period 1980Q1-2019Q4. Following the inference of the existing literature that structural breaks and trends are important considerations for the analysis of stochastic properties of non-renewable natural resource prices, we have utilized two stationarity tests, one is designed to detect smooth breaks, and the other one is designed to capture abrupt changes in trend. The motivation behind the consideration of smooth and sharp breaks is to avoid any misspecification of the functional form of the breaks, which could be as problematic as ignoring the breaks.

Our results have revealed that the empirical evidence in favor of stationarity of metal prices increases when structural breaks are properly accommodated. In that sense, the sharp break stationarity test of Carrion-i Silvestre and Sanso (2007) has appeared to be decisive in uncovering evidence for stationarity of metal prices, while the smooth break stationarity test of Becker et al. (2006) seemed to perform relatively poor in capturing the sharp breaks in terms of break magnitudes and timings. More specifically, we have observed that almost all metal prices follow a trend stationary path with the observed structural changes being related to market-specific and economic events, though global economic conditions have appeared to be effective especially on precious metal prices. This finding implies inefficiency in all these markets and necessitates one to perform technical analysis to predict prices and making profits. Moreover, for these markets, stabilization policies will be effective in dealing with exogenous shocks, which will be temporary and short-lived. The only exception to this finding is silver, which has appeared to be characterized by the efficient market hypothesis with stationarity being rejected for all specifications considered in the paper. This result suggests that the effects of exogenous shocks on silver prices would be permanent, and strong policy measures should be implemented to return silver prices back to their original trends. Since the utilization of non-renewable resources is important for future sustainability of economies, predicting their prices is of utmost importance. The results are important for both hedgers, arbitrageurs, speculators but also for policymakers.

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