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A reexamination of trends in primary commodity prices

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ABSTRACT

This paper conducts a reexamination of the Prebisch–Singer hypothesis by employing the unit root test proposed by Lee and Strazicich (2003) and Lee and Strazicich (2004) that allow for up to two structural breaks. Given the higher power of these tests compared to the Zivot and Andrews (1992) and Lumsdaine and Papell (1997) tests, rejection of the null can be considered as genuine evidence of stationarity. The main findings of this paper are that eleven out of twenty-four commodity prices are found to be difference stationary implying that shocks to these commodities tend to be permanent in nature. The remaining thirteen prices are found to exhibit trend stationary behavior with either one or two structural breaks. Most of the commodities that do not exhibit difference stationary behavior seem to contain no significant trends. There are fewer cases, in relation to past studies, of commodities that display negative trends thereby weakening the case for the Prebisch–Singer hypothesis.

1. Introduction

An issue which has dominated much of the literature when analyzing commodities is the possible existence of a long run trend in primary commodity prices relative to the price of manufactures. Classical economists such as David Ricardo and John Stuart Mill held the view that this trend should be positive as the supply of primary commodities would be constrained by the fixed amount of land while the supply of manufactures would be augmented by technical progress. However, Prebisch (1950) and Singer (1950) reversed this view claiming that commodity prices should decline in relation to manufactured goods in the long run, referred to as the Prebisch–Singer hypothesis. Prebisch argued that strong labor unions in countries that export manufactures cause wages to ratchet upwards during times of boom but prevent wages from falling during times of recession. On the other hand, countries that export primary commodities have weaker labor unions which are unable to increase wages during times of boom and cannot prevent a fall in wages during times of recession. Thus primary commodity prices increase by less than manufactures during upswings but fall more than manufactures during times of downswings leading to a secular decline in real commodity prices. Singer argued that the manufacturing sector has monopoly power which prevents the technical progress from lowering prices. Besides, the low income elasticity of demand for primary commodities would cause the decline in primary commodity prices in relation to manufactured goods.

The subject of primary commodities has led to much debate as it has been used to explain the widening gap between developed and less developed countries leading to a large volume of studies that empirically test the Prebisch–Singer hypothesis. The evidence has been mixed which leads to serious policy implications as to whether developing countries should specialize in primary commodity exports. The World Bank for instance, has encouraged long term primary commodity projects in developing countries (Ardeni and Wright, 1992). Besides, free market solutions were provided to developing countries to deal with primary commodities instead of positive intervention (Maizels, 1994). The upshot is that countries which rely on the exports of primary commodities must understand the nature of commodity prices in order to devise their development macroeconomic policy (Deaton, 1999).

The continuing interest in the Prebisch–Singer hypothesis stems from the fact that the central question is an empirical one. When considering the possibility of the existence of a trend in real commodity prices one is naturally led to the question of whether the prices are trend stationary or difference stationary.[1] The standard econometric tool used to discriminate between the two alternatives is the unit root test. The unit root test is useful as it aids in distinguishing

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[1] In recent developments in the time series literature, test for the significance of the trend can be carried out irrespective of whether the data contains a unit root or not. See Harvey et al., 2007. We would like to thank an anonymous referee for this comment. In the context of this study, unit root tests are useful as they lend empirical support as to whether real commodity prices contain stochastic trends or not.

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whether the real commodity prices are characterized by stochastic trends or not. If the price series is found to contain a unit root then the series is said to contain stochastic trends such that the effect of shocks to the underlying price series will be permanent. If however, the underlying price series is found to reject a unit root then the series is considered to be trend stationary and the effect of shocks on the price series would be transitory in nature. The original work on the Prebisch–Singer hypothesis has assumed that the underlying data series is trend stationary (Prebisch (1950), Singer (1950)). While Sapsford (1985), Grilli and Yang (1988), Helg (1991) and Ardeni and Wright (1992) among others have advocated for commodity prices to follow a trend stationary model, Cuddington and Urzua (1989), Cuddington (1992), Bleaney and Greenaway (1993) and Newbold et al. (2005) recognized that commodity prices may be difference stationary. The evidence on the Prebisch–Singer hypothesis has been mixed. While Sapsford (1985) and Helg (1991) tend to support the hypothesis, in contrast, Newbold and Vougas (1996) cannot provide compelling evidence as to whether the data is trend stationary or difference stationary, while Kim et al. (2003) suggest that commodity prices generally display unit root behavior and that there is limited evidence for the Prebisch–Singer hypothesis.

Perron (1989) showed that if a structural break is ignored the power of the unit root test is lowered. His paper however, was criticized for the fact that he assumed that the date of the structural break is known. Zivot and Andrews (1992), developed a unit root test that allowed for the break to be unknown and determined endogenously from the data. Since the unit root test suffers from low power by ignoring a single structural break, it was argued by Lumsdaine and Papell (1997), that the loss of power is greater if there are two structural breaks. They formulated a method which is an extension of the Zivot–Andrews, which tests for a unit root under the null hypothesis against the alternative of two structural breaks determined endogenously from the data.

The issue of structural breaks applied to primary commodity prices has been of great interest to researchers. Leon and Soto (1997) applied the single break Zivot–Andrews test on primary commodity prices and found evidence of structural change. In contrast to Kim et al. (2003) their results show that most commodity prices can be described as trend stationary models and that the trend coefficients generally support the Prebisch–Singer hypothesis. Zanias (2005) and Kellard and Wohar (2006) have employed the Lumsdaine–Papell test to allow for two structural breaks. Zanias employs the test to the extended aggregate Grilli Yang index and concludes the data to be a trend stationary process with two intercept shifts. The breaks are identified in 1920 and 1984 which cumulatively account for a 62% drop. However, Zanias (2005) considered the aggregate index, which has been questioned by recent studies, arguing that individual commodities behave in a different manner. The study by Kellard and Wohar (2006), KW hereafter, is a case in point. KW conduct a study of the disaggregated commodity prices over the period of 1900–1998. Out of 24 commodity prices, their results indicate that 14 are trend stationary allowing for 1 or 2 structural breaks. Overall, they show that the deterioration of commodity prices has been discontinuous. Following Leon and Soto (1997) KW argue that the existence of a single linear trend would be 'strong evidence' in favor of the Prebisch–Singer hypothesis. KW state that a single trend is a 'summary measure' of several trends which may be positive or negative. Arguing that reliance on a single trend may be misleading to policy makers, KW create a measure to define the prevalence of a trend. The study by KW can be seen as an attempt to extend the results obtained by Cuddington and Urzua (1989) and Kim et al. (2003) by noting that where evidence is found for the real commodity prices to be difference stationary, those prices are found to lend less support to the Prebisch–Singer hypothesis.

More recently, Balagtas and Holt (2009) test and estimate for a linear unit root model against smooth transition and time varying alternatives. They note that while previous research focused on breaks in the intercept and trend, a more complete analysis would allow for breaks in the autoregressive and moving average terms. However, though Balagtas and Holt (2009) argue that the nonlinearity in commodity price adjustment arises due to the impossibility of negative storage (Deaton and Laroque, 1995), they do not attempt to provide any economic intuition for such breaks. Further, no explanation is provided as to why such breaks should be smooth. For example, Ocampo and Parra (2003) have argued that the decline in terms of trade have been discontinuous and notable, especially for periods in the 1920s and 1980s. A drawback of the past studies is the choice of a null hypothesis that allows for a linear unit root. This paper attempts to address this gap by allowing breaks in the null hypothesis of a unit root.

However, when considering unit root tests that allow for structural breaks, there may be a case of size distortion which leads to spurious rejection of the null hypothesis of a unit root when the actual time series process contains a unit root with a structural break (Nunes et al., 1997). The test developed by Lumsdaine and Papell (1997) suffers from the same size distortion problems and consequent spurious rejection of the null hypothesis (Lee and Strazicich, 2003). The minimum two-break LM test developed by Lee and Strazicich (2003) allows for structural break under the null hypothesis and does not suffer from the spurious rejection of the null hypothesis. Besides, the minimum LM test possesses greater power than the Lumsdaine–Papell test. The upshot is that under the LM test setting, rejection of the null hypothesis can be considered as genuine evidence of stationarity. A further disadvantage of the Zivot–Andrews test and the Lumsdaine–Papell test is that the tests tend to estimate the breakpoint incorrectly where bias in estimating the unit root test is the greatest (Lee et al., 2006). This leads to size distortion which increases with the magnitude of the break. This size distortion does not occur when using the LM test as it employs a different detrending method (Lee et al., 2006).

This paper tests for unit roots in commodity prices by employing the method proposed by Lee and Strazicich (2003, 2004) to allow for up to two structural breaks. Following the study by KW, the prevalence of trends is examined by calculating the ratio of the number of years a statistically negative trend is found to exist to the whole sample. Finally, the paper describes the policy implications of the results. The organization of the paper is as follows: Section 2 describes the econometric methods, Section 3 presents the data and the empirical results, Section 4 outlines a discussion of the empirics and finally Section 5 concludes.

2. Econometric methodology

The Prebisch–Singer hypothesis argues that (log) relative commodity prices steadily decrease over time. When conducting an empirical test on the hypothesis one needs to take into account the underlying nature of the data as the conclusions reached by researchers will be dependent on the econometric model. The price series under consideration may be trend stationary or difference stationary. If the underlying commodity price series were to be trend stationary, then to test the Prebisch–Singer hypothesis one typically estimates the following log-linear time trend model:

$$P_t = \alpha + \beta t + \varepsilon_t \tag{1}$$

where $P_t = \ln(P_t^C / P_t^M)$, that is, the log ratio of commodity price series to the manufacturing unit value index and the errors denoted by ε_t is a white noise process. The coefficient β on the time index t measures the growth rate. If $\beta > 0$ then it indicates an improvement in the terms of trade otherwise for $\beta < 0$ we conclude a deterioration in the terms of trade. The error process in Eq. (1), that is ε_t, is assumed to follow an ARMA process to allow for cyclical fluctuations of relative commodity prices around their long run trend. If the relative commodity price series P_t, were to contain a unit root then estimating the trend stationary model given by Eq. (1) will generate spurious estimates of the trend, by concluding that the trend is significant when it is actually not. An appropriate strategy for estimating the trend is to adopt the following difference stationary model:

$$\Delta P_t = \beta + \eta_t \tag{2}$$

where η_t is a stationary and invertible error process. In the difference stationary model, if we find a negative estimate of β which is statistically significant, we may conclude that the underlying price series supports the Prebisch–Singer hypothesis. An important point to note is that if the real commodity price series is a trend stationary process but is treated as a difference stationary process, then tests based on Eq. (2) are inefficient, lacking power relative to those based on Eq. (1).

However, as discussed in the earlier section, if structural break(s) are ignored the power of the unit root test is lowered. In this paper we consider the unit root test subject to two endogenously determined structural breaks (Lee and Strazicich, 2003) and a single structural break (Lee and Strazicich, 2004). To briefly describe the Lee and Strazicich (2003) method, consider the following data generating process (DGP):

$$P_t = \psi' X_t + u_t \text{ and } u_t = \phi u_{t-1} + \varepsilon_t; \text{ where } \varepsilon_t \sim \text{iid} N\left(0, \sigma^2\right) \tag{3}$$

where P_t is the price series and the two changes in level and trend are given by $X_t = [1, t, D_{1t}, D_{2t}, DT_{1t}, DT_{2t}]'$, where

$$DT_{jt} = \begin{cases} t - TB_j \text{ for } t \geq TB_j + 1 \\ 0 \text{ otherwise} \end{cases} \quad \text{for} j = 1, 2.$$

TB_j denotes the points at which the breaks occur. Note that the DGP contains breaks in the null hypothesis when $H_0{:}(\phi = 1)$ and the alternative hypothesis when $H_A{:}(\phi < 1)$. The break fractions are denoted as $\lambda_j = TB_j/T$ where T denotes the total number of observations.

When employing the Lee and Strazicich (2004) method which considers a single structural break, the single change in level and trend in Eq. (3) is now given by $X_t = [1, t, D_t, DT_t]'$, where

$$DT_t = \begin{cases} t - TB \text{ for } t \geq TB + 1 \\ 0 \text{ otherwise} \end{cases}$$

TB denotes the points at which the breaks occur. And the break fraction is denoted as $\lambda = TB/T$.

The LM unit root test statistic can be estimated by the following regression:

$$\Delta P_t = \phi' \Delta X_t + \gamma \overline{\Upsilon}_{t-1} + \sum_{i=1}^{p} \psi_i \Delta \overline{\Upsilon}_{t-i} + u_t \tag{4}$$

where $\overline{\Upsilon}_t = P_t - \mu - X_t \bar{\phi}$, $t = 2, 3, \ldots, T$; $\bar{\phi}$ are coefficients on the regression of ΔP_t on ΔX_t; μ is given by $P_1 - X_1 \bar{\phi}$.[2] The lagged terms $\Delta \overline{\Upsilon}_{t-i}$ are added to correct for serial correlation. The augmentation is determined using the general to specific method. The LM test statistics are given by the τ statistic testing the null hypothesis $H_0{:}(\gamma = 0)$. The LM unit root test determines the break points endogenously by utilizing a grid search. To eliminate endpoints trimming of the *infimum* (inf) is made at 10%. The breakpoints are determined where the test statistic is minimized. The LM test is given as $LM_\tau = \inf \hat{\tau}(\lambda)$.

In small samples it has been argued that asymptotic critical values can be misleading. Besides, to facilitate a comparison of our results with those of KW we compute finite sample critical values for the unit root tests statistic using a parametric bootstrap procedure. Following KW, we initially difference each price series and attempt to fit an autoregressive model. The optimal AR order is chosen using the Schwartz Bayesian Criterion (SBC). For each commodity price series we then construct 1000 pseudo series of first differences with the sample size equal to the actual series using the AR coefficients from the optimal model with randomly drawn residuals which are independently and identically distributed with zero mean and constant variance. We then compute the cumulative sum of the pseudo simulated first differences of the series to generate the pseudo unit root series.[3] The Lee–Strazicich test is then carried out on the pseudo unit root series and the minimum t-statistic is calculated. From this distribution the 1%, 5% and 10% critical values are constructed.

If we find that we can identify up to 2 structural breaks in any price series, in the next step we attempt to determine whether the sign of the trend is negative or positive and whether it is significant or not. Besides, it would be of interest to observe whether the trend changes signs in the different regimes which are outlined by the structural breaks. In order to determine how the trend has shifted over time we proceed to model the data generating process in the manner conducted by KW. For the trend stationary process the logarithm of the commodity price series is regressed against a constant and a time trend and one or two intercept and slope dummy variables corresponding to the results of the structural break tests. The error structure is modeled as an ARMA(p,q) process. Thus the estimation process is carried out using the following equations:

$$P_t = \gamma + \delta_1 t + \delta_2 D_{L1,t} + \delta_3 D_{T1,t} + \delta_4 D_{L2,t} + \delta_5 D_{T2,t} + u_t \tag{5}$$

$$u_t - \phi_1 u_{t-1} - \ldots\ldots - \phi_p u_{t-p} = \varepsilon_t - \psi_1 \varepsilon_{t-1} \ldots\ldots - \psi_q \varepsilon_{t-q} \tag{6}$$

where ε_t is a white noise process. D_{Li} and D_{Ti}($i = 1,2,3$) denote the level and slope dummy respectively; i refers to the regime defined by the prior identification of the break dates. The regimes are defined as:

regime 1 = start date to $TB1$
regime 2 = ($TB1$ + 1) to $TB2$
regime 3 = ($TB2$ + 1) to end date.

Following KW, Eq. (5) is reparameterized to facilitate the estimation of the trend coefficient in the three different regimes. Eq. (5) was reparameterized in the following way:

$$\begin{aligned} P_t^* = {} & \gamma R_{L1,t} + \delta_1 R_{T1,t} + (\gamma + \delta_2) R_{L2,t} + (\delta_1 + \delta_3) R_{T2,t} + (\gamma + \delta_2 + \delta_4) R_{L3,t} \\ & + (\delta_1 + \delta_2 + \delta_3) R_{T3,t} + u_t \end{aligned} \tag{7}$$

where $R_{\mathrm{Li},t}$ denotes the intercept dummy for a level shift in regime i, ($i = 1,2,3$) and $R_{\mathrm{Ti},t}$ denotes the slope dummy for a trend shift in regime i, ($i = 1,2,3$). For the three regime models, P_t^* is defined as:

$$P_t^* = \begin{cases} P_t \text{ if } 1 \leq t \leq TB1 \\ P_t - \delta_1 (TB1 - 1899) \text{ if } TB1 + 1 \leq t \leq TB2 \\ P_t - \delta_1 (TB1 - 1899) - \delta_3 (TB2 - TB1) \text{ if } t > TB2 \end{cases}.$$

To facilitate comparison with KW, the estimation was conducted by exact maximum likelihood and the ARMA order (p,q) was selected through the SBC allowing all possible models with $p + q \leq 6$.

3. Data and empirical results

Prior to Grilli and Yang's (1988) seminal paper, there were many inadequacies in the availability of consistent commodity price data in the economic field. The *Economist* Index and the W. A. Lewis Index were then two of the main indices which provided a large backlog of data on commodity prices, even though their level of data accuracy

[2] Where P_1 and X_1 denote the first observations of the P_t and X_t sequences respectively.

[3] See Kellard and Wohar (2006) for details.

had often been called into question. The former had been subject to frequent revisions and was weighted by "the relative values of commodities in the import trade of industrial countries" (Grilli and Yang, 1988, p.3), resulting in a one-sided focus on the import patterns of developed countries, and it also excluded fuel-based commodities. On the other hand, the W. A. Lewis Index only ran until 1938 and used the export unit values of certain countries rather than international market quotations.

In terms of the available data on manufactured goods series, again the W. A. Lewis Index started in 1870 but had gaps in the data owing to both World Wars. Maizels (1970, cited by Grilli and Yang, 1988, p.3) also constructed a series but this only comprised average prices of certain periods. Finally, although the United Nations' data on manufacturing unit values (MUVUN) spanned the entire twentieth century, again it contained gaps in the data during the war years.

Given the then situation regarding data inadequacy, Grilli and Yang (henceforth referred to as GY) opted to construct "a U.S. dollar index of prices of twenty-four internationally-traded nonfuel commodities, beginning in 1900" (GY, 1988, p.3). The Grilli Yang Commodity Price Index (GYCPI) is "base-weighted, with 1977–1979 values of world exports of each commodity used as weights" (GY, 1988, p.3) and is a means of capturing the evolution of international prices of a basket of primaries.

As for the updated version of the MUVUN (the MUV), the gaps in the United Nations data are corrected for, via interpolations using U.S. and U.K. data on export and import unit values. This updated series "reflects the unit values of exports of manufactures of a number of industrial countries" (GY, 1988, p.5), with weights which vary to show the changing importance of different manufactured goods over time — such changes are shown in several updates, which occur every few years until 1938 and then again in: 1959, 1963, 1970, 1975, and 1980 (United Nations 1969, 1972, 1976, 1982, and 1987, cited in GY, 1988, p.5). GY then go on to use the series GYCPI/MUV which "measures the evolution of the purchasing power of a basket on nonfuel primary commodities in terms of traded manufactures, valued at 'international prices'" (GY, 1988, p.7).

One caveat to this construction process is that neither the manufactured goods price series nor the primary commodity price series can be complete proxies for the components of the net barter terms of trade, as developing countries' level of total imports comprise more than manufactured goods, whilst developed countries' total imports comprise more than just primary commodities. Furthermore, GY place emphasis on the fact that a declining trend in relative prices should not be taken solely as a declining real income effect, the income effect relies not only on movements in relative prices but also on the evolution of the purchasing power of exports; what is more, the authors highlight that "one has to account simultaneously for the movements in the relative prices of exports and for the quantity of exports" (GY, 1988, p.7) — an expression which reflects this is the income terms of trade which "reflects the purchasing power of total exports in terms of imports" (p.7).

An extended data set of the original GYCPI is employed in this study.[4] The data set consists of 24 primary commodity prices measured annually over the period 1900–2003 and deflated by the MUV index. Fig. 1 plots the 24 commodity prices deflated by the MUV index.

As a prelude to the test of unit roots allowing for structural breaks, we employ two unit root tests developed by Ng and Perron (2001) which are modified versions of the Phillips Perron unit root test and have superior size and power properties. We report the Ng and Perron tests labeled MZ_ρ and MZ_t along with the lag length k selected according to the modified AIC as in Table 1 below.

The preliminary results at this stage show that 17 commodity prices were found to contain a unit root. For the remaining 7 prices (sugar, beef, lamb, rubber, timber, aluminum and zinc) the null hypothesis of a unit root was rejected in favor of a trend stationary alternative. Comparing the results with KW we find that the test results are affected by the sample size. At the 10% level 5 of the 7 price series are an exact match. The remaining two prices (aluminum and beef) which are found to be trend stationary are different to the findings by KW.[5]

The unit root tests due to Ng and Perron (2001) ignore the possibility of structural breaks in either the level or trend of the price series. Since the non-rejection of the null hypothesis of a unit root may be as a result of not accounting for breaks in the intercept or trend, we proceed to investigate the issue of structural breaks in commodity prices. Before embarking on the LM test we employ the Perron and Rodrìguez (2003) test which is an extension of the Ng and Perron (2001) test where a change in the trend function is allowed to happen at an unknown time adopting the GLS detrending approach by Elliot et al. (1996).[6] We choose the model where the break point is selected by minimizing the t-statistic on the coefficient of the trend for the GLS detrended ADF and M-type test regressions due to Perron and Ng (1996) as it is shown to have a higher power than a model which maximizes the absolute value of the t-statistic on the coefficient of the trend. The modified AIC is chosen to determine the autoregressive truncations lag as it leads to tests with better overall properties (Perron and Rodrìguez, 2003). Table 2 below shows the results of the Perron–Rodriguez test below.

The results in Table 2 show that there is no evidence against the unit root for the following commodity series: coffee, cocoa, tea, beef, lamb, hides, aluminum and silver. For the remaining 16 commodities[7] we can reject the null hypothesis of a unit root accepting the alternative that the process is trend stationary with a single structural break.[8] Comparing our results with those of Leon and Soto (1997) we find as one would expect, fewer rejections of the null.[9] However, the Perron–Rodriguez test only allows for a single break. For the reasons described in the earlier section we will now proceed to test for up to two structural breaks using the method developed by Lee and Strazicich (2003) and Lee and Strazicich (2004).

Table 3 reports the unit root tests allowing for structural breaks. The top half of the table tests for unit roots allowing for two endogenously determined structural breaks (Lee and Strazicich, 2003) and the lower half of the table tests for unit roots allowing for a single endogenously determined structural break (Lee and Strazicich, 2004).

The results in Table 3 (under the column heading Lee–Strazicich tests) confirm that 11 price series are found to be trend stationary with two structural breaks, 2 price series are trend stationary with a single structural break and the remaining 11 price series (not reported here) are non-stationary containing a unit root. When comparing the Lee–Strazicich test results with those of the Perron–Rodriguez test we

[4] The extension of the original Grilli Yang Index was obtained from the supplementary data set in Pfaffenzeller et al. (2007). A detailed explanation of how the data is updated is documented in the paper by Pfaffenzeller et al. (2007).

[5] Note that the Ng–Perron tests conducted by KW use the 5% significance level to reach their conclusions. When considering the 10% level one can obtain a closer match. The different results that appear are due to the different lag length chosen for the unit root test which was obtained using an extended sample.

[6] We thank an anonymous referee for raising this point.

[7] Note that for Timber and Wool the estimated statistics are just about significant at the 10% level.

[8] One can note that when comparing across the M-type and ADF tests, for 5 commodities, (that is, rice, cotton, wool, timber and jute) the evidence is mixed where the ADF test results do not match with the M-tests. Further, there is a slight discrepancy in the selected break dates for 3 commodity price series and though the discrepancy is marginal for wheat and maize, it is more pronounced in the case of sugar. However, the ADF type tests are reported for completeness, we only rely on the M-type tests for comparison.

[9] The number of rejections falls further when the comparison is made with the sample size used by that of Leon and Soto (1997). This result lends support to the case made by Lee et al. (2006). The results are available from the authors on request.

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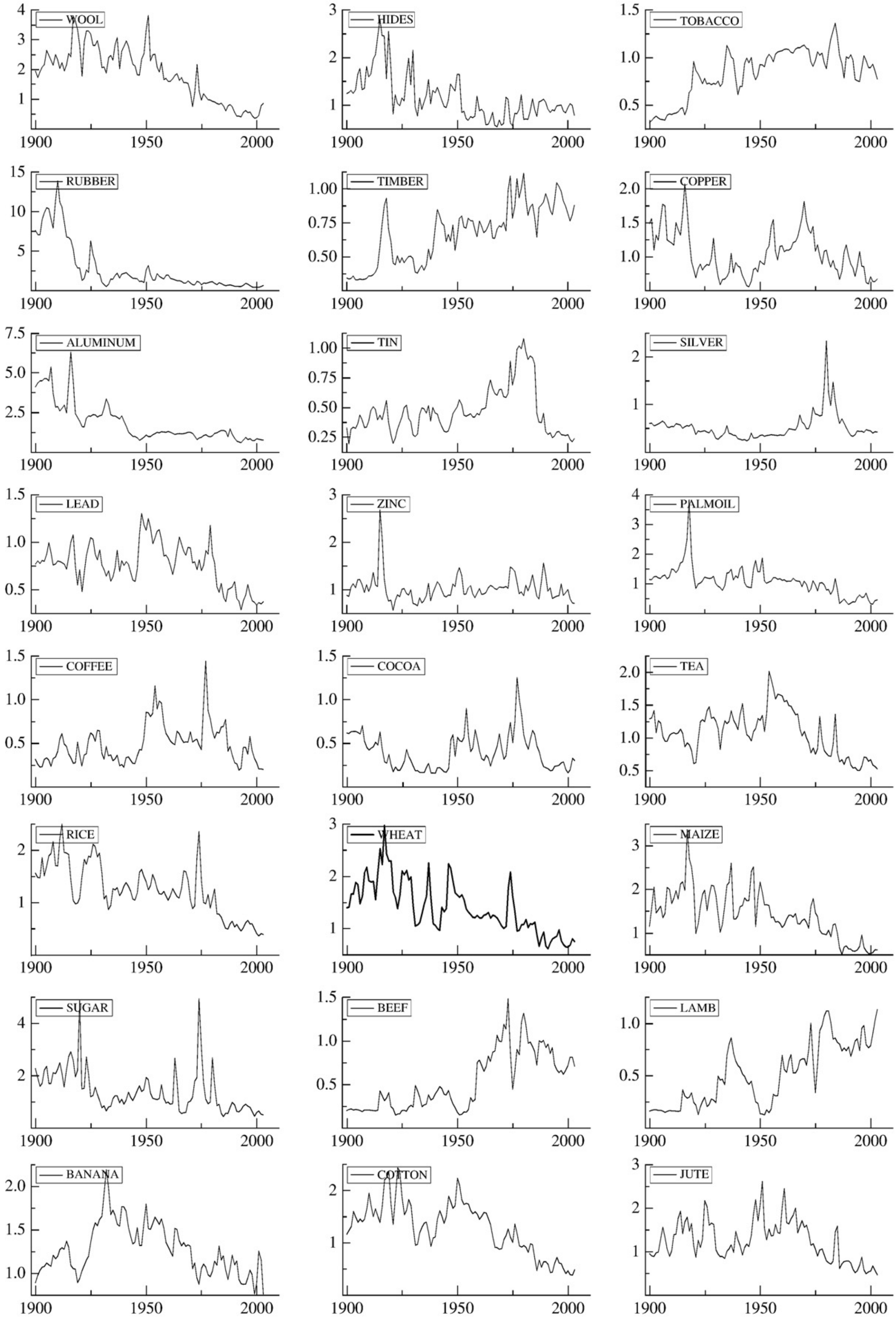

Fig. 1. A reexamination of trends in primary commodity prices.

Table 1
Unit root tests (Ng and Perron, 2001).

	MZ_ρ	MZ_t	k
Coffee	−14.15	−2.50	0
Cocoa	−8.25	−2.03	2
Tea	−9.20	−2.07	2
Rice	−10.26	−2.14	4
Wheat	−2.68	−1.11	8
Maize	−4.70	−1.45	4
Sugar	−16.63*	−2.86*	5
Beef	−14.57*	−2.67*	0
Lamb	−15.40*	−2.77*	0
Banana	−3.53	−1.12	2
Palm oil	−8.42	−2.01	5
Cotton	−4.24	−1.35	3
Jute	−7.31	−1.76	2
Wool	−3.59	−1.34	4
Hides	−3.62	−1.34	10
Tobacco	−1.45	−0.58	4
Rubber	−17.45**	−2.92**	0
Timber	−18.91**	−3.04**	0
Copper	−14.03	−2.64*	0
Aluminum	−14.59*	−2.68*	0
Tin	−9.15	−2.00	0
Silver	−6.53	−1.80	2
Lead	−13.93	−2.57	0
Zinc	−29.97***	−3.81***	0

***, ** and * denote significance at the 1%, 5% and 10% levels respectively. *k* denotes the number of lags chosen according to the modified AIC.

find that the null is rejected for 13 commodities using the former method and 16 when using the latter. The results are broadly similar in that there are 13 exact matches. While the LM test concludes that maize, sugar and jute contain a unit root, Perron–Rodriguez concludes trend stationary with a structural break.

To determine how our results compare with the recent study by KW which employs the Lumsdaine–Papell unit root test, we observe that the results are different. KW find 10 price series to be trend stationary with two structural breaks and 4 are trend stationary with a single structural break. The difference lies in the fact that only 4 price series out of the 13 match those of KW. These include wool, rubber, timber and tin. While KW find maize, hides and silver to be trend stationary with two structural breaks, our findings suggest that these prices are non-stationary with a unit root. Further, KW find 7 price series (these include cocoa, wheat, sugar, banana, cotton, tobacco and copper) to contain a unit root, whereas our findings note that 4 of these prices (wheat, cotton, tobacco and copper) are trend stationary with two structural breaks. When considering a single structural break, KW find rice, palm oil, jute and aluminum to exhibit a trend stationary process. Our findings reveal that none of these prices display a single structural break; while jute is found to be a unit root process, the remaining three price series are trend stationary with two breaks. Note, however, that these results are sensitive to the sample size.

The sample size chosen in this study is slightly longer than that of KW and more so in the case of Leon and Soto (1997). To facilitate comparison, the Lumsdaine–Papell test was employed to all the commodities over the sample period chosen in this study, which is 1900 to 2003.[10] When considering 2 structural breaks, the results (see Table 3, column heading Lumsdaine–Papell Tests) show that the null hypothesis of a unit root is rejected for only 7 commodities using the Lumsdaine–Papell test compared to 11 using the LM test. Comparing with the LM test, out of the 7 commodities, 4 are common, that is rice, rubber, tin and lead. Interestingly, these results are quite different to those obtained by KW highlighting the difference the results can make when choosing a sample with different end points. When considering a single structural break, the Lumsdaine–Papell test reject the null of a unit root for 3 commodities[11] whereas only 2 are rejected using the LM test. None of the commodities for each test show any match. In total, when employing the Lumsdaine–Papell test, we find 10 commodities that exhibit trend stationary behavior with structural break(s). For 5 commodities, (that is, hides, tea, maize, aluminum and silver) we find support for the view by Lee and Strazicich (2003) that in the presence of breaks under the null, the Lumsdaine–Papell test is prone to over-rejection. Overall, we observe that the results change when employing different unit root tests and considering different sample sizes. Clearly both sample size and the choice of the unit root test can affect the results. The subsequent analysis will build on the findings from the Lee and Strazicich (2003, 2004) tests.

Given the different results, the remaining part of this paper now proceeds to investigate the prevalence of trends in the primary commodity prices. Table 4 shows the result of the estimated trend stationary model with 2 and 1 structural breaks respectively.

Out of the 13 trend stationary price series, 11 prices were found to exhibit a significant negative trend, though not necessarily for the entire sample. For timber we find a significant positive trend, whereas for zinc we find no evidence of a significant positive or negative trend. Comparing the results with KW we find that there are 9 commodities in common that exhibit a trend stationary process. While the break dates are somewhat closer for lead and rubber, for the rest of the commodities the break dates are far apart. It is not surprising therefore that the estimated trends within the regimes are different. What is interesting is that for a small difference in the break dates (for example, lead) the impact on the estimated slope coefficients can be great.[12] This study provides a more accurate picture of the trends as the break dates are estimated with more precision than the Lumsdaine–Papell test (Lee et al., 2006).

Table 5 shows the result of the estimated difference stationary models for those price series which were found to exhibit unit root behavior. A difference stationary model is estimated with the error

Table 2
Unit root tests with a single structural break (Perron and Rodrìguez, 2003).

	MZa	TB	MZt	TB	ADF	TB
Coffee	−18.85(2)	1953	−3.06(2)	1953	−3.18(1)	1911
Cocoa	−17.30 (2)	1930	−2.93 (2)	1930	−3.02 (2)	1930
Tea	−20.51(2)	1956	−3.20(2)	1956	−3.28 (2)	1956
Rice	−32.04 (4)***	1921	−3.98 (4)***	1921	−3.25 (4)	1921
Wheat	−63.65 (2)***	1960	−5.61 (2)***	1960	−5.76 (1)***	1957
Maize	−53.59 (2)***	1926	−5.16 (2)***	1926	−5.70 (1)***	1928
Sugar	−38.75(2)***	1924	−4.31 (2)***	1924	−4.13 (1)*	1940
Beef	−21.35 (1)	1944	−3.22 (1)	1944	−3.35 (1)	1944
Lamb	−11.13 (2)	1959	−2.34 (2)	1959	−2.36 (2)	1959
Banana	−42.16 (1)***	1967	−4.58 (1)***	1967	−5.89 (1)***	1967
Palm oil	−25.26 (2)*	1921	−3.53 (2)*	1921	−3.87 (2)*	1921
Cotton	−24.04 (3)*	1953	−3.36 (3)*	1953	−3.09 (3)	1953
Jute	−23.99 (2)*	1943	−3.46 (2)*	1943	−3.71 (2)	1943
Wool	−22.69 (2)	1929	−3.36 (2)*	1929	−3.55 (2)	1929
Hides	−4.90 (10)	1951	−1.55 (10)	1951	−1.95 (10)	1951
Tobacco	−62.26 (1)***	1915	−5.57 (1)***	1915	−6.07 (1)***	1915
Rubber	−41.87 (1)***	1964	−4.57 (1)***	1964	−5.00 (1)***	1964
Timber	−22.75 (2)	1966	−3.37 (2)*	1966	−3.79 (2)	1966
Copper	−29.84 (1)**	1936	−3.83 (1)**	1936	−4.11 (1)*	1936
Aluminum	−20.33 (1)	1963	−3.17 (1)	1963	−3.28 (1)	1963
Tin	−28.32 (1)**	1929	−3.76 (1)**	1929	−3.96 (1)*	1929
Silver	−12.57 (2)	1930	−2.46 (2)	1930	−2.46 (2)	1929
Lead	−32.49 (1)***	1923	−4.03 (1)***	1923	−4.33 (1)**	1923
Zinc	−47.07 (1)***	1916	−4.85 (1)***	1916	−5.52 (1)***	1916

***, ** and * denote significance at the 1%, 5% and 10% levels respectively. The numbers within parentheses denote the lag length. TB denotes the break date.

[10] We thank an anonymous referee for raising this point.

[11] In this case, the test is essentially the Zivot and Andrews (1992) test.

[12] In the case of lead, we chose the break dates of KW and obtained the same results for the signs and statistical significance of the trend coefficients. However, when choosing our break dates, which are close to the break points of KW the conclusions about the trend coefficients change. Our results indicate that the second and third regimes exhibit significant negative trends while for the break dates employed by KW we find significant negative trends for the first two regimes.

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Table 3
Unit root test with structural breaks.

	Lee–Strazicich unit root tests						Lumsdaine–Papell tests					
Unit root test with 2 breaks in intercept and trend												
	TB1	TB2	*t*-min (lags)	Bootstrapped crit. values			TB1	TB2	*t*-min (lags)	Bootstrapped crit. values		
				1%	5%	10%				1%	5%	10%
Rice	1970	1982	−6.59 (7)***	−6.03	−5.74	−5.53	1946	1972	−6.47 (12)*	−7.10	−6.75	−6.31
Wheat	1921	1985	−6.37 (3)**	−6.61	−6.16	−5.88						
Palm oil	1922	1981	−6.04 (1)*	−6.88	−6.22	−5.92						
Cotton	1928	1950	−6.59 (1)**	−6.74	−6.21	−5.98						
Wool	1956	1988	−6.72 (3)*	−7.41	−6.96	−6.69						
Tobacco	1922	1956	−6.17 (2)*	−6.85	−6.34	−6.02						
Rubber	1928	1940	−5.52 (1)*	−6.31	−5.68	−5.40	1929	1942	−6.35 (1)*	−6.91	−6.61	−6.31
Timber	1925	1938	−5.85 (4)***	−6.12	−5.66	−5.43						
Copper	1949	1974	−5.43 (1)*	−6.20	−5.68	−5.40						
Tin	1957	1984	−5.76(11)***	−6.75	−6.01	−5.72	1941	1985	−6.94 (1)**	−7.33	−6.75	−6.41
Lead	1945	1983	−7.12 (3)***	−6.29	−5.76	−5.51	1946	1981	−6.56 (0)*	−7.43	−6.73	−6.47
Hides							1919	1953	−7.20 (3)**	−7.05	−6.86	−6.48
Silver							1946	1978	−6.82 (6)**	−7.11	−6.64	−6.35
Zinc							1917	1973	−7.37 (1)**	−7.17	−6.71	−6.35
Unit root test with 1 break in intercept and trend												
Banana	1928		−5.75 (1)**	−5.22	−5.44	−6.09						
Zinc	1972		−5.28 (1)*	−5.07	−5.35	−6.02						
Tea							1952		−5.26(2)*	−5.74	−5.49	−5.25
Maize							1985		−6.12(0)*	−5.70	−5.47	−5.21
Aluminum							1939		−5.14(1)*	−5.73	−5.49	−5.17

***, **, and * denote significance at the 1%, 5% and 10% levels respectively. The numbers in parentheses denote the lag length selected according to the General to Specific methodology as suggested by Lee and Strazicich (2003). TB1 and TB2 denote the first and second break dates respectively.

term allowed to follow an ARMA (p,q) process. Out of the 11 commodities considered, only 4 commodities (that is, tea, maize, hides and jute) showed significant signs of a trend, while the remaining 7 commodities are found to be trendless. Comparing the results with KW we find that there are only 3 out of 11 commodities that are in common that display a difference stationary process, being cocoa, coffee and lamb. Similar to KW's findings however, these three commodities exhibit a difference stationary process with no drift.

Given that we find evidence of significant structural breaks under the null hypothesis for three commodities (that is, hides, tea and maize) we find support to the claim by Lee and Strazicich (2003) that the Lumsdaine–Papell test is prone to over-rejection of the null hypothesis for these commodities.

Following KW, we synthesize the results from the analysis done so far by constructing $\Psi(-)$ that measures the prevalence of a negative trend. For each commodity we calculate $\Psi(-)=\lambda(-)/T$ where $\lambda(-)$ equals the number of years that a statistically significant negative trend exists. In the same way we calculate the measure of the prevalence of a positive trend $\Psi(+)$ (defined as $\Psi(+)=\lambda(+)/T$) and trendless behavior $\Psi(.)$ (defined as $\Psi(.)=1-\Psi(-)-\Psi(+)$). Table 6 displays the relative measure results for all 24 commodities.

The prevalence of a negative trend is found in 16 commodities. That is, each of these commodities displays at least one significant negative trend segment. Out of these 16 commodities, 6 commodities display a negative trend for more than 70% of the sample period and the number rises to 10 commodities if we were to consider at least 50% of the sample period. When considering positive trends, only 2 commodities display a positive trend for more than 70% of the sample and 3 commodities for more than 50%. For trendless behavior we find evidence of 8 commodities that cover at least 70% of the sample and 11 when considering at least 50% of the sample. When comparing our results with those of KW only 3 commodities show similar trends, being coffee, cocoa and lamb which exhibit a trendless behavior. If one were to adopt the approach by KW, that evidence for the Prebisch–Singer hypothesis can be attributed to finding a negative trend which exists for at least 70% of the sample, then our results further weaken the case for the Prebisch–Singer hypothesis. While KW find 8 commodities (that is, a third of the commodities) to support the

Table 4
Estimated trend stationary model with breaks.

	TB1	TB2	Regime 1	Regime 2	Regime 3	ARMA
Rice	1970	1982	−0.54*** (−2.92)	−5.01** (−2.49)	−3.41*** (−3.14)	2,0
Wheat	1921	1985	1.81** (2.43)	−0.60*** (6.91)	−0.70 (0.67)	2,2
Palm oil	1922	1981	1.35 (1.11)	−0.50* (1.73)	−2.40** (2.05)	0,2
Cotton	1928	1950	1.20*** (2.78)	2.07*** (3.28)	−2.84*** (−16.14)	2,0
Wool	1956	1988	0.13 (0.68)	−3.11*** (−6.74)	−4.15*** (−2.91)	0,1
Tobacco	1922	1956	3.95*** (7.92)	0.76*** (2.87)	−0.54*** (3.23)	3,0
Rubber	1928	1940	−7.26*** (−8.98)	−2.25 (−0.69)	−2.62*** (−15.19)	2,1
Timber	1925	1938	2.14*** (2.60)	0.50 (0.30)	0.61** (2.54)	1,0
Copper	1949	1974	−1.92*** (−11.01)	1.72*** (3.16)	−1.42*** (−3.30)	2,2
Tin	1957	1984	0.39 (1.26)	2.99*** (2.89)	−10.46*** (−2.60)	2,1
Lead	1945	1983	−0.09 (−0.31)	−1.19* (−1.88)	−3.51* (−1.87)	2,1
Banana	1928	NA	1.25*** (3.15)	−0.93*** (−9.08)	NA	0,2
Zinc	1972	NA	0.02 (0.12)	−1.07 (−1.50)	NA	1,0

TB1 and TB2 denote the first and second break dates respectively. The slope coefficients are recorded for regimes 1, 2 and 3. ***, **, and * denote significance at the 1%, 5% and 10% levels respectively. The final column presents the ARMA (p,q) specification. The numbers in parentheses denote the t-ratios.

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Table 5
Estimated difference stationary models with breaks.

	TB1	TB2	Regime 1	Regime 2	Regime 3	ARMA
Coffee	1931	1948	0.0002	0.016	−0.01	0,0
			(0.006)	(0.27)	(−0.40)	
Maize	1919	1974	0.01**	−0.008***	−0.03***	0,2
			(2.21)	(−4.43)	(−7.36)	
Tea	1922	1953	−0.01	0.015	−0.02***	0,2
			(−0.98)	(1.32)	(−2.31)	
Cocoa	1945	1988	−0.02	0.007	−0.01	2,0
			(−0.78)	(0.25)	(−0.19)	
Sugar	1924	1971	−0.009	−0.005	−0.028	0,2
			(−0.32)	(−0.26)	(−1.20)	
Lamb	1945	1960	0.02	0.03	0.01	0,0
			(0.59)	(0.55)	(0.32)	
Jute	1930	1966	0.002	0.007**	−0.03***	0,2
			(0.45)	(2.21)	(−9.27)	
Silver	1943	1984	−0.02	0.03	−0.05	0,0
			(−0.70)	(1.21)	(−1.19)	
Hides	1918	1973	0.01	−0.01***	0.015**	0,2
			(1.18)	(−7.22)	(1.98)	
Aluminum	1941	1987	−0.02	−0.01	−0.01	0,0
			(−0.89)	(−0.47)	(−0.37)	
Beef	1947	1963	0.01	0.04	0.0001	0,0
			(0.35)	(0.79)	(0.003)	

TB1 and TB2 denote the first and second break dates respectively. The final row presents the ARMA (p,q) specification. ***, **, and * denote significance at the 1%, 5% and 10% levels respectively. The numbers in parentheses denote the t-ratios.

Table 6
Relative measures of a prevalence of a trend.

	$\Psi(-)$	$\Psi(+)$	$\Psi(.)$
Coffee	0.000	0.000	1.000
Cocoa	0.000	0.000	1.000
Tea	0.481	0.000	0.519
Rice	1.000	0.000	0.000
Wheat	0.615	0.202	0.183
Maize	0.818	0.182	0.000
Sugar	0.000	0.000	1.000
Beef	0.000	0.000	1.000
Lamb	0.000	0.000	1.000
Banana	0.731	0.269	0.000
Palm oil	0.778	0.000	0.222
Cotton	0.509	0.491	0.000
Jute	0.352	0.346	0.288
Wool	0.452	0.000	0.548
Hides	0.528	0.299	0.173
Tobacco	0.462	0.538	0.000
Rubber	0.885	0.000	0.115
Timber	0.000	0.875	0.125
Copper	0.760	0.240	0.000
Aluminum	0.403	0.000	0.597
Tin	0.182	0.259	0.559
Silver	0.000	0.000	1.000
Lead	0.557	0.000	0.443
Zinc	0.000	0.000	1.000

Prebisch–Singer hypothesis, the current study finds only 6 commodities (that is, a fourth of the commodities). Comparing our results with Leon and Soto (1997) the results are drastically different. While Leon and Soto (1997) find evidence that 17 commodity prices exhibit a negative trend for all or most of the (1990–1992) sample period, our results are much lower at 7 commodities. Overall, based on a longer data set and more powerful unit root tests, we find that there is evidence to further weaken, albeit marginally, the Prebisch–Singer hypothesis.[13]

The results suggest that the trend is variable, (except for rice) there is no evidence of a single negative trend. We can concur with Kellard and Wohar (2006) that forecasting of commodity prices should not typically occur about a single trend. The first break date is concentrated in the 1920s and 1940s (that is, for 10 out of the 13 commodities). The exceptions are rice (1970s), tin and wool (1950s). The second structural break is relatively varied. About half (that is, 7 out of 13) experience a break date in the 1980s. The exceptions are copper (1970s), beef (1960s), cotton and tobacco (1950s), rubber (1940s) and timber (1930s).

4. Discussion and policy implications

A wide range of unit root tests were carried out on an updated Grilli Yang Index. Starting with the unit root tests due to Ng and Perron (2001) that do not allow for structural breaks, we find 17 commodities that display unit root process and the remaining 7 commodity prices are trend stationary. However, when allowing for a single structural break using a method developed by Perron and Rodrìguez (2003) that extends the Ng–Perron test, we find 16 commodity prices to display trend stationary behavior. Applying the Lumsdaine and Papell (1997) test to allow for up to two structural breaks, we find the results to change with only 10 commodity prices to exhibit a trend stationary process. Finally, the Lee and Strazicich (2003, 2004) test concludes 13 prices to display trend stationary behavior. Given the nature of the unit root test of Lee and Strazicich (2003, 2004) we can conclude that the finding of stationarity can be considered as genuine evidence. In light of the recent study by KW which applied the Lumsdaine–Papell test on the commodity prices but for a slightly shorter period, we can observe that the results are sensitive to both the choice of test and the sample size. We further observe, that the results vary for the individual commodities concerned.

The Lee and Strazicich (2003, 2004) unit root tests determine the break points endogenously where the t-test statistic is minimized. One has to be careful as to not interpret these tests as a test for structural breaks.[14] The test is to determine whether the commodity prices contain a unit root. This does however require one to locate the breaks, which is carried out using the Lee and Strazicich (2003) method of utilizing a grid search. The location of the breaks does show some signs of coinciding with major events that took place in commodity prices. For instance, in the case of rice, the pattern of trends coincides with an in-depth study of rice markets made by Kajisa and Akiyama (2005). When considering wheat, the sharp change in trend in 1921 coincides with the sharp downturn in manufacturing (Powell, 1991). In the case of tin, the trend is found to be positive between first and the last International Tin Agreement (that is, 1956 and 1985); the dates seem to match closely the observed location of the breaks. The trend patterns observed for copper match closely to a study by Crowson (2007).

What do these results mean in terms of policy? KW argue that the underlying price movements, whether they be trend stationary or difference stationary can seriously affect the income and consumption levels of developing countries. Stabilization policies were introduced to smooth income flows. While it has been argued that stabilization policies are effective when the price series is trend stationary, they may be difficult to implement if the transitory shocks are large and take a lot of time to dissipate, or if the price series have a varying trend (Reinhart and Wickham, 1994). Our evidence indicates that such policies would be either ineffective or difficult to implement. Besides, most commodities do not maintain buffer stocks and have abandoned price stabilization policies. These include cocoa where buffer stock operations ended in 1988; coffee, where regulated exports were

[13] We find evidence that even after choosing the same sample used in this study (1900–2003) and employing the Lumsdaine–Papell test, we find limited evidence of a negative prevalent trend. The results, from the Lee–Strazicich test in comparison with the Lumsdaine–Papell test, still show relatively less support for the Prebisch–Singer hypothesis. The results are not reported here for brevity but are available from the authors on request.

[14] We thank an anonymous referee for raising this point.

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abandoned in 1989; sugar, where price stabilization measures were removed in 1992; and jute where price stabilization ended with the 1989 agreement. This study shows that for these commodities we find evidence of difference stationary behavior. More recently, in the case of tin and rubber, the buffer stock has been gradually sold off. When countries experience a setback due to a period of negative trend in primary commodity prices, the impact of the shock can be partly offset by obtaining finance from the IMF under CFF (Compensatory Financing Facility) scheme or from the EU under STABEX (Stabilization System for Export Earnings) scheme[15]. These policies should be based on the underlying prevalence of the trend, and evidence from this study suggests that the design and form of assistance would be difficult to implement given the mixed and varying trend results.

It has been argued that countries that face a declining trend in their terms of trade (largely due to primary commodities comprising a large portion of their exports) can lead to a current account deficit. This deficit can be compensated by allowing for a capital account surplus driven by increased public and private borrowing. In the event where a country experiences both current and capital account deficits, one may observe a decline in international monetary reserves and currency instability. Given that for several commodities (coffee, cocoa, tea, maize, sugar, lamb, jute, hides, silver and zinc) we find no evidence of a trend, one may infer from these results that foreign exchange constraints facing developing countries can be relaxed (Newbold et al., 2005). However, Newbold et al. (2005) point out that the risk of revenue shortfall during the period of time when the repayments are anticipated can be quite high. Given that we find very little evidence of a persistent long run trend, one cannot draw conclusions regarding future terms of trade. We find evidence that some commodities experience segments of a downward trend interspersed by periods of approximate stability. However, forecasting of commodity prices may be difficult since the break points can be unpredictable.

The prevalence of a negative trend, albeit less in comparison to recent studies, is mainly found for commodities exported by developing countries. Singer (1999) argues that developing countries need to diversify their exports into manufacturing irrespective whether there is a declining trend or not. The experience of some of the countries in Asia, such as South Korea, Singapore, Hong Kong and Taiwan, shows that they moved towards successful export diversification. However, many other developing countries, especially in Africa have been left behind. Institutions such as the World Bank encouraged countries to diversify into new exports of primary commodities. It has been argued that exports should be diversified away from primary commodities that are in oversupply. Diversification of exports in to other primary commodities would depend on the existing and potential resource availability and possible export destinations. An argument that has been raised against this policy is that many commodity prices tend to move together since they are commonly substitutes. While empirical evidence to support this notion was provided by Pindyck and Rotemberg (1990), weak evidence for this relationship was found by Leybourne et al. (1994) and Deb et al. (1996). However, Page and Hewitt (2001) argue that this policy is an unlikely substitute for industrialization, particularly for small countries.

Would promoting industrialization be the major policy implication? Lutz and Singer (1994) indicate that policies by Bretton–Woods institutions have intentionally or not, promoted the production and expansion of primary commodities by developing countries, contributing to the declining trend of commodities. In more recent years, studies have hinted that the trends in prices of manufactures from developing countries may be in decline (Sarkar and Singer, 1991; Maizels et al., 1998). The exponents of inward looking development strategies tend to make a case for import substitution. The evidence suggests that such policy measures can be detrimental given the mixed results obtained on price behavior for different commodities. For example, when considering major exports of developing countries, we find that rice displays a trend stationary process with a negative trend for the entire period. This is in sharp contrast to cocoa and coffee which exhibit a driftless random walk. The heterogeneity of the results for the estimated trends obtained for individual prices confirms the evidence obtained by Leon and Soto (1997) that the use of aggregate measures (for example, Zanias (2005)) may be misleading. Besides, import substitution has become unpopular with countries such as Brazil and India. Both countries had initially embraced this policy but subsequently rejected them in favor of liberalized market policies with particular reference to exports (Sapsford and Balasubramanyam, 1994).

On a final note, one should exercise caution when drawing policy implications from studies that estimate and analyze trends in primary commodity prices. These results provide information about the trends of relative primary commodity prices and not the real value of exports or the distribution of export revenue (Diakosavvas and Scandizzo, 1991).

5. Conclusion

This paper employs unit root tests to determine whether real primary commodity prices contain stochastic trends. As described earlier, the presence of a structural break lowers the power of unit root tests. Lee and Strazicich (2003, 2004) develop a minimum LM test which incorporates structural change in the null hypothesis. These LM tests are more powerful than other tests for unit roots and are meaningful in the application of testing for trends in commodity prices as rejection of the null can be considered as genuine evidence of stationarity. Employing the minimum LM test we obtain quite different results from comparable unit root tests due to Lumsdaine and Papell (1997) and those obtained in recent studies by Kellard and Wohar (2006) and Leon and Soto (1997). However, as noted earlier, the different end points in all these studies can have an effect on the overall conclusions. The main findings of this paper are that eleven out of twenty-four commodity prices are found to be difference stationary implying that shocks to these commodities tend to be permanent in nature. The remaining thirteen prices are found to exhibit trend stationary behavior allowing for one or two structural breaks. The changes in economic conditions and environment over the length of time chosen for this study justify the case to allow for structural breaks. There are fewer cases, in relation to past studies, of commodities that display negative trends thereby weakening the case for the Prebisch–Singer hypothesis. With the different commodities analyzed in this study, we observe different patterns of trends. Given that we find evidence that some commodities experience segments of a downward trend interspersed by periods of approximate stability, forecasting of commodity prices proves to be difficult. The evidence from this study suggests that policy recommendations would be difficult to implement given the mixed and varying trend results.

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[15] See Engel and Meller (1993) for details.

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