On the Prevalence of Trends in Primary Commodity Prices

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Abstract

This paper applies new time-series procedures to examine the Prebisch-Singer hypothesis of a secular deterioration in relative primary commodity prices. Previous literature has generally approached the issue by analyzing whether a single downward trend exists. We allow for two structural breaks giving scope to assess whether the trend is more usefully represented as a segmented phenomena. Employing a dataset of 24 relative commodity prices for the 1900-98 period, unit root tests find that 14 series are trend-stationary after allowance for (up to) two breaks. Furthermore, for the majority of commodities, the trend is not well represented by a single downward slope, but instead by a shifting trend that often changes sign over the sample period. Specifically, of the 15 commodity prices shown to present at least one negative trend, a measure of the prevalence of a negative trend shows that for 7 cases the deterioration exists for less than 70% of the sample period. Unlike some recent work that has also allowed for structural breaks, these results provide much less support for the Prebisch-Singer hypothesis.

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1. Introduction

The time series properties of the prices of primary commodities relative to an index of manufacturing prices has important implications for both producer and consumer countries. Examining long-run trends, Prebisch (1950) and Singer (1950) presented both theoretical justification and empirical evidence that there was a downward secular trend in relative primary commodities prices over the period 1870-1945. What has subsequently become popularized as the Prebisch-Singer (PS) hypothesis therefore argues that (log) relative commodity prices are steadily decreasing over time. To clarify, assume that y_t (the logarithm of the relative commodity price) is generated by a trend-stationary (TS) process:

$$y_t = \alpha + \beta t + u_t \tag{1}$$

where t is a linear trend and the random variable u_t is stationary with mean zero. The focal point of interest is the slope parameter β which the PS hypothesis predicts will be less than zero. Early empirical evidence focusing on estimations of (1) found strong support for the PS hypothesis[1]. Some of the explanations that have been offered include productivity differentials between countries, asymmetric market structure (where manufacturing industries capture oligopolistic rents relative to competitive firms earning zero economic profits and producing primary commodities) and high income elasticity of demand for manufacturing goods relative to that of primary commodities. One corollary of these findings is that developing countries, to the degree that they export primary commodities and import manufactures, will be subject to a secular deterioration in their net barter terms of trade. The clear policy implication is to diversify exports away from primary commodities or stimulate domestic production of manufactures.

The alternative generating process for the data is represented by the difference-stationary (DS) model:

$$\Delta y_t = \beta + v_t \tag{2}$$

where the generating process for v_t is stationary and invertible. Recent empirical studies employing techniques which permit commodity prices to contain a stochastic trend, have found evidence against the PS hypothesis. Notably, Kim *et al.* (2003) suggest that commodity prices exhibit unit root behavior and that in the small samples used this results in the slope coefficient in (1) being biased downwards. Modeling the 24 commodity prices that comprise the Grilli-Yang index as a DS processes, Kim *et al.* (2003) finds that just five commodity prices had the negative trend predicted by the PS hypothesis[2].

Of course Perron (1989), *inter alia*, has shown that the appearance of unit root behavior can be caused by the presence of infrequent structural breaks. A number of studies have investigated the possibility of shifting deterministic components in international commodity prices[3]. Specifically, Leon and Soto (1997) apply unit root tests allowing for structural change. Employing the Zivot and Andrews (1992) endogenous break point methodology to individual commodity prices in the Grilli-Yang index, they allow for one break in the deterministic trend. Of the 24 commodities, and after allowance for (up to) 1 break, 20 are classified as TS models for the 1900-92 period, implying that shocks to commodity prices, in several cases, do not possess the permanent component suggested by Kim *et al.* (2003). Moreover, 17 commodity prices report a negative trend and thus provide evidence in support of the PS hypothesis.

Many studies have emphasized the existence of *multiple* turning points in commodity prices including Popkin (1974), Cooper and Lawrence (1975), Enoch and Panic (1981), Bosworth and Lawrence (1982) and Chu and Morrison (1984). Therefore, in an extension to the work of Leon and Soto (1997), this study provides new evidence by applying the Lumsdaine and Papell (1997) unit root test that allows for two shifts in the mean and trend under the alternative hypothesis to individual commodity prices in an extended 1900-1998 Grilli-Yang index[4]. As a further extension of previous work in this area we note that there is substantial serial correlation observed in many commodity prices. We thus employ a parametric bootstrap to generate appropriate unit root test critical values which allow for temporal dependence under the null hypothesis of a unit root. Allowing for temporal dependence under the unit root null generally results in larger critical values. Thus, previous studies that employ either asymptotic critical values or use bootstrap critical values that do not account for temporal dependence under the unit root null will tend to reject a true unit root null too often.[5] Our results indicate that 14 of our 24 commodity price series are trend-stationary after allowance for (up to) two breaks and it is shown that 15 of the 24 commodity prices present a significant negative trend *over some segment of the sample period*.

Previous literature has generally approached the issue by analyzing whether a single downward trend exists. If a downward trend exists this is taken as evidence in favor of the PS hypothesis. Thus, as Leon and Soto (1997), under such criteria we would now claim strong evidence for the PS hypothesis. Of course the estimation of a single trend as such is a summary measure which may include a comprise of several trends, both positive and negative. A reliance on a single negative measure may mislead policy makers into believing that a relative commodity price is steadily decreasing over time when there is no reason *a priori* why this should be the case. However, allowing for the possibility of two endogenously determined break dates allows assessment of whether the trend is more usefully represented as a segmented phenomena.

In this paper, a simple measure of the prevalence of a negative trend is proposed as a means to assess the contribution of allowing for multiple breaks in commodity prices on our evidence for the PS hypothesis. Specifically, negative trend prevalence is computed as a ratio of the number of years that a statistically significant negative trend exists to the total number of sample years. Thus, we find that although 15 of the 24 commodity prices present at least one significant negative trend, 12 of these last for less than 90% of the sample period; 7 last for less than 70% of the sample period; and 4 for less than 50%. When one defines as evidence for the PS hypothesis that a negative trend lasts for at least 70% of the sample period, the evidence is clearly less than overwhelming.

The remainder of the paper is organized as follows: Section 2 describes the empirical estimation methodology. Section 3 presents the data and the empirical results. Section 4 proposes some measures of the prevalence of a trend and a discussion of the results. Finally, section 5 concludes.

2. Empirical Methodology

Perron (1989), Zivot and Andrews (1992), and Lumsdaine and Papell (1997), *inter alia*, have shown that the investigation of whether a series is TS or DS using standard Dickey-Fuller (1979, 1981), Phillips and Perron (1988) or Said and Dickey (1984) unit root tests can lead to wrong inferences if there are infrequent structural breaks in the intercept and/or trend terms. Perron (1989) allows for one exogenous predetermined shift in the deterministic trend. Christiano (1992) and Stock and Watson (1988a, 1988b) show that an exogenously chosen break date may lead to false inferences. In response, Bannerjee *et al.* (1982), Zivot and Andrews (1992), Perron and Vogelsang (1992a, 1992b), and Perron (1994) have developed recursive and sequential unit root tests in which the break point is estimated rather than selected a priori. Lumsdaine and Papell (1997) extends the Zivot and Andrews methodology from one endogenously chosen break date to two.

Consider the unit-root test developed by Lumsdaine and Papell (1997). The procedure allows for two distinct shifts in both the intercept and trend terms which are determined endogenously. The null and alternative hypotheses are given as follows:

$$\begin{aligned} &H_0 : y_t = \mu_0 + y_{t-1} + \varepsilon_t \\ &H_1 : y_t = \mu_0 + \beta t + \mu_1 D_{L1,t} + \mu_2 D_{T1,t} + \mu_3 D_{L2,t} + \mu_4 D_{T2,t} + \varepsilon_t \end{aligned} \tag{3}$$

where the μ's are coefficient parameters and ε_t is a well-defined error term. $D_{L1,t}$ and $D_{L2,t}$ are level dummy variables defined as follows:

$D_{L1,t} = 1$, if $t > TB1$, zero otherwise;

$D_{L2,t} = 1$, if $t > TB2$, zero otherwise.

$D_{T1,t}$ and $D_{T2,t}$ indicate shifts in the trend function defined as follows:

$D_{T1,t} = (t - TB1)$, if $t > TB1$, zero otherwise;

$D_{T2,t} = (t - TB2)$, if $t > TB2$, zero otherwise.

TB1 and *TB2* are defined as first and second hypothesized break dates. In other words, the null hypothesis of a unit root is tested against the alternative that the series is TS with two distinct shifts in the intercept and deterministic trend. The actual test is based on the following regression:

$$\begin{aligned} \Delta y_t = {} & \mu_0 + \beta t + \rho y_{t-1} + \mu_1 D_{L1,t} + \mu_2 D_{T1,t} + \mu_3 D_{L2,t} + \mu_4 D_{T2,t} \\ & + \sum_{j=1}^{k} d_j \Delta y_{t-j} + \varepsilon_t \end{aligned} \tag{4}$$

where ρ and d_j are coefficient parameters. If ρ is not significantly different from zero, then shocks to the logarithm of relative commodity prices are permanent and have a unit root. On the other hand, if ρ is significantly less than zero, the unit root null hypothesis is rejected and shocks have temporary effects. The k extra regressors Δy_{t-j} are intended to eliminate possible nuisance-parameter dependencies in the asymptotic distributions of the test statistics caused by serial correlation in the error terms. We examine the above model (which Lumsdaine and Papell, 1997, refer to as Model CC) as well as a version which allows for only a single break in intercept and trend term.

At each possible break date, we search for the optimal lag length *k*. A general-to-specific method is employed starting with k_{max} equal to 5[6]. If the coefficient of the last included lag difference term is significant at the 10% level, select $k = k_{max}$. Otherwise, reduce the order of lags by one until the coefficient on the last included lag differenced term is statistically significant[7]. The selection of break dates *TB1* and *TB2* correspond to the specification that yields the minimum t-statistic (largest in absolute value) associated with the coefficient ρ.

The null hypothesis of a unit root with drift is tested against the alternative hypothesis of trend stationarity allowing for two shifts in the trend and/or intercept terms. As asymptotic critical values are often misleading in small samples, small sample critical values are computed for the ADF test statistic using a parametric bootstrap procedure. We compute finite sample critical values for our test statistics for each individual commodity price series. Initially autoregressive (AR) models are fitted to the first differences of each commodity price series, using SBC to select the optimal AR order of the model. We then (for each commodity price series) construct 1000 pseudo series of first differences with sample size equal to the actual series, using the AR coefficients from the optimal model (for each commodity price series) with randomly drawn residuals from iid $N(0,\sigma^2)$ errors, where σ^2 is the estimated error variance of the optimal model. Pseudo unit root series without structural change, but allowing for temporal dependence in the data, are subsequently generated by computing the cumulative sum of the pseudo simulated first differences. For each of the 1000 pseudo unit root series, the Lumsdaine and Papell (1997) unit root testing methodology, allowing for breaks in trend described earlier, is then carried out. The specification with the smallest t-statistic (largest in absolute value) associated with ρ is calculated for each of the pseudo series generated. The 1%, 5% and 10% critical values for the ADF statistic are then constructed from this empirical distribution.

With conventional unit root tests that allow shifts in the trend function under the alternative, one is normally only concerned with bootstrapping the t-statistic associated with the lagged level term (i.e. the ADF term). The significance of the intercept and trend dummy variables is not an issue because if one allows for infrequent shifts in the intercept and/or trend under the alternative, and such shifts do not actually exist, this will only decrease the power of the test. Thus, if one rejects the unit root null with superfluous structural breaks, one will also reject the unit root null not allowing for such breaks. The same is true with respect to the number of breaks that one allows for. If two breaks are allowed for and only one break is the true model, then the power of the test will be decreased. The critical values rise (in absolute value) causing a loss of power if too many breaks are included. Alternatively, if the true data generating process were trend stationary, failing to include a time trend result in a reduction in the power of the test (Papell, 2003). In addition, the loss of power from excluding a time trend when it should be present is more severe than the reduction in power associated with including a time trend when it is extraneous; see West (1987).

However, when investigating the PS hypothesis, the sign of the trend coefficient and how it may change over time is of crucial importance. It is interesting to investigate whether the trend term has undergone any change in its mean level or its slope. As previously noted, the most general model considered allows for two breaks in both intercept and trend under the alternative hypothesis. If the unit root with drift null hypothesis cannot be rejected in favor of the alternative of two breaks in intercept and trend, we next consider a model that allows only one break in intercept and trend term. This procedure is employed as an over-parameterized model (i.e. one which includes more breaks when they are not actually present) will result in a decrease in power of the test.

After determining which series are trend stationary (allowing for shifts in intercept and trend terms) it is then possible to model the relevant data generating process. For example, for those series for which we are able to reject the null hypothesis in favor of the trend stationary alternative with either two or one break we next conduct analysis whereby the (log) commodity series is regressed on a constant and time trend and two (or one if the case arises) intercept and trend dummy variables corresponding to the break dates determined from the unit root tests and model the error structure as an ARMA (p,q) process. This is because while the unit root tests are not tests for structural change, once one is able to reject the null hypothesis one can conduct analysis *under the specification of the alternative* with the understanding that it is possible that there may in fact be fewer breaks than that specified under the alternative. We proceed in this fashion in order to determine how the trend has shifted over time. Thus, for each TS commodity series in the 2 break case we conduct the following regression:

$$y_t = \alpha + \beta_1 t + \beta_2 D_{L1,t} + \beta_3 D_{T1,t} + \beta_4 D_{L2,t} + \beta_5 D_{T2,t} + u_t \tag{5}$$

$$u_t - \phi_1 u_{t-1} - \ldots\ldots - \phi_p u_{t-p} = \varepsilon_t - \theta_1 \varepsilon_{t-1} - \ldots\ldots - \theta_q \varepsilon_{t-q} \tag{6}$$

where ε_t is zero-mean white noise. Of course the prior identification of break dates allows the defining of three regimes where:

regime 1 = start date to *TB1*;

regime 2 = (*TB1*+1) to *TB2*;

regime 3 = (*TB2*+1) to end date.

As we are interested primarily in the trend coefficient and its value in the 3 different regimes equation (5) can be reparameterised in the following way:

$$\begin{aligned} y_t^* = \alpha R_{L1,t} + \beta_1 R_{T1,t} + (\alpha + \beta_2) R_{L2,t} + (\beta_1 + \beta_3) R_{T2,t} \\ + (\alpha + \beta_2 + \beta_4) R_{L3,t} + (\beta_1 + \beta_3 + \beta_5) R_{T3,t} + u_t \end{aligned} \tag{7}$$

where $R_{Li,t}$ is an intercept dummy variable for a level shift in regime *i* (*i* = 1, 2, 3) , $R_{Ti,t}$ is a trend dummy variable for a trend shift in regime[8] *i* (*i* = 1, 2, 3) and y_t^* is a modified y_t , adjusted[9] to allow for the level effects of $R_{Ti,t}$. The parameters of (6) and (7) were estimated jointly through exact maximum likelihood, assuming a Gaussian error distribution, using the OX package. The autoregressive-moving average order (*p,q*) was selected through the Schwarz Bayesian Criterion (SBC)[10], allowing all possible models with $p + q \leq 6$.

3. Data and Results

To facilitate comparison of our results with previous studies, an extended series of the original Grilli and Yang (1988) index is employed, where each nominal commodity price (in US dollars) is deflated by the United Nations Manufactures Unit Value (MUV) index. The data set covers the period 1900-1998, comprises 24 commodities and uses annual values in natural logarithms. Figure 1 plots the natural logarithm of 24 commodity prices relative to the MUV index.

Two recent unit root tests, labeled the MZ_ρ and the MZ_t test, and developed by Perron and Ng (1996) are modified versions of the PP tests that have much better size and power properties than the conventional tests. While the current paper is concerned primarily with the test of the unit root hypothesis that allows for the possibility of breaks at unknown dates under the alternative of no unit root, for completeness, we report the MZ_ρ and the MZ_t statistics as well as the lag length *k*, selected using the modified information criteria (MIC); see Ng and Perron (2001). These results (which employ GLS detrending) are reported in Table 1. The unit root tests indicate that for all series, with the exception of hide, lead, rubber, timber and zinc, the unit root hypothesis cannot be rejected. For these other series the null of unit root is rejected in favor of the trend stationary alternative. These results are in accordance with the early literature that subjected commodity prices to unit root tests. As discussed earlier, the non-rejection of the null of unit root may be the result of infrequent shifts in deterministic trends. We investigate this issue next.

Table 2 reports the unit root test results allowing for shifts in the deterministic trends. The results reported in our table 2 are presented in a similar format to papers such as Papell (2003) who also conduct unit root tests allowing for 2 endogenously determined breaks. For 10 commodity prices (hide, lead, maize, rubber, silver, tea, timber, tin, wool and zinc) the null hypothesis of a unit root with drift is rejected (at least at the 10% level) in favor of the alternative of two breaks in both intercept and trend. While five of these series were found to be trend-stationary using the M-tests in Table 1, we proceed with the idea that we can further refine the trend component, on the grounds that such analysis can lead to greater power advantages, as we do not know the true data generating process.

Consider now the series for which the unit root null was not rejected. In the second half of Table 2, the results allowing for one break in intercept and one break in trend under the alternative are reported for 4 series (aluminum, jute, palmoil and rice). We find that we can reject the null of unit root in favor of the alternative of trend stationarity around one mean and one trend shift. Finally, for 10 commodities (banana, beef, cocoa, coffee, copper, cotton, lamb, sugar, tobacco and wheat) the unit root null cannot be rejected regardless of the specification of the alternative hypothesis.

To summarize, 14 commodities are classified as TS after allowance for (up to) two breaks. Of these, 10 are characterized as TS with two breaks in trend and intercept and 4 TS with one break in intercept and trend. For 10 of the 24 commodities we cannot reject the unit root null. Clearly, given that only five commodities were found to be TS using the M-tests, allowing for the possibility of two structural breaks under the alternative hypothesis greatly affects the conclusions of unit root tests.

In Table 3, relative commodity prices that are TS with two breaks in the intercept and trend (and one break in intercept and trend when appropriate) are modeled by estimating equations (6) and (7). Notably, when the fitted models contain an autoregressive component, the estimated root is not close to one, suggesting little evidence of under-differencing. Of the 14 TS models, 7 (aluminum, jute, lead, rice, rubber, tea and wool) contain a negative and significant trend (although in some cases not for the entire series) and 1 (hide) contains a positive and significant trend. Interestingly, 6 series contain both positive and negative significant trends (maize, palmoil, silver, tin, timber and zinc). Finally, Table 4 shows estimated *I(1)* models for those relative commodity prices which were found to exhibit unit root behavior. In other words, we estimate DS model (2) with the error term permitted to follow an ARMA (p,q) process. Some 8 commodities (banana, beef, cocoa, coffee, copper, cotton, lamb and tobacco) are shown to be trendless with the remaining two (sugar and wheat) possessing negative and significant trends. However, it is important at this stage to note some possible shortcomings in the analysis. Firstly, of course, that the null hypothesis of the unit root test does not allow the possibility of commodity prices being a DS process *with* structural breaks. Secondly, the estimation of TS models in small samples has been shown to produce spurious rejections of a zero trend, in the presence of strong autocorrelation (see, *inter alios*, Kim *et al*., 2003). Given these it seems a useful exercise to re-estimate the TS models assuming they are *I(1)* series. Thus, we conduct the following regression:

$$\Delta y_t = \beta_1 + \beta_2 \Delta D_{L1,t} + \beta_3 \Delta D_{T1,t} + \beta_4 \Delta D_{L2,t} + \beta_5 \Delta D_{T2,t} + v_t \tag{8}$$

$$v_t - \phi_1 v_{t-1} - \ldots\ldots - \phi_p v_{t-p} = \varepsilon_t - \theta_1 \varepsilon_{t-1} - \ldots\ldots - \theta_q \varepsilon_{t-q} \tag{9}$$

which can be reparameterised in the following way:

$$\Delta y_t = \beta_1 \Delta R_{T1,t} + \beta_2 \Delta R_{L2,t} + (\beta_1 + \beta_3)\Delta R_{T2,t} + \beta_4 \Delta R_{L3,t} + (\beta_1 + \beta_3 + \beta_5)\Delta R_{T3,t} + v_t \tag{10}$$

In Table 5 the results from (10) are presented. Notably almost all models have a moving average root equal or close to unity[11] as would be expected if the series are *I(0)* and the (*p,q*) order has been correctly specified[12]. The coefficients and t-statistics on the trend and intercept terms are very similar across the TS and DS estimation suggesting that the strong autocorrelation observed in the unit root tests without structural breaks is removed by the addition of structural breaks and that the conclusions of the previous sections are robust to specification. This extends the work of Cuddington and Urzua (1989) and Kim *et al.* (2003) who note that evidence for the PS hypothesis, without allowing for the possibility of structural breaks, is greatly reduced when you look at DS specification as compared to TS specifications.

4. Discussion

To aid in gauging the relevance of the results to the PS hypothesis, a simple relative measure, $\psi(-)$, of the prevalence of a negative trend is constructed for each commodity:

$$\psi(-) = \frac{\lambda(-)}{N} \tag{11}$$

where $\lambda(-) =$ number of years that a statistically significant (at the 10% level) negative trend exists and $N =$ total number of sample years. To understand what this measure shows consider the case of maize. Estimating a DS specification using the same ARMA *(p,q)* methodology as outlined earlier without structural breaks gives the following estimation:

$$\Delta y_t = \underset{(-1.37)}{-0.88(0.01)} + \varepsilon_t + 0.23\varepsilon_{t-1} + 0.46\varepsilon_{t-2}$$

where t-statistics are written in parentheses. Including structural breaks gives the regression indicated in Table 3:

$$y_t = 4.86R_{L1,t} + \underset{(3.76)}{3.34(0.01)}R_{T1,t} + 4.44R_{L2,t} \underset{(-1.98)}{-0.42(0.01)}R_{T2,t} + 4.21R_{L3,t} \underset{(-3.27)}{-2.72(0.01)}R_{T3,t} + \varepsilon_{t\,t} - 0.54\varepsilon_{t-1}$$

The estimation without structural breaks clearly indicates an *insignificant* negative trend of less than 1% per year over the duration of the sample from 1900 to 1998. Previous studies have taken this as evidence against the PS hypothesis. However, including structural breaks, a significant negative trend is apparent from the first break point in 1920. Clearly any evidence against the PS hypothesis comes only from before this break point. The measure in (11) conveys this by simply showing the proportion of the sample period that a significant negative trend exists and in the case of maize this is 0.79. A summary measure of this form also allows for quick comparisons between commodity series.

For completeness it is useful to consider two complementary measures[13] to $\psi(-)$. Firstly, $\psi(+)$, a measure of the prevalence of a positive trend:

$$\psi(+) = \frac{\lambda(+)}{N} \tag{12}$$

where $\lambda(+) =$ number of years that a statistically significant (at the 10% level) positive trend exists. Secondly, $\psi(.)$, a measure of the prevalence of trendless behavior:

$$\psi(.) = 1 - \psi(+) - \psi(-) \tag{13}$$

Table 6 displays the relative measure results for all commodities. The derived measure of negative trend prevalence $\psi(-)$ demonstrates that 15 of the 24 commodity prices present at least one significant negative trend and therefore, superficially at least, there is strong evidence for the PS hypothesis. However, of the 15 commodity prices that exhibit at least one significant negative trend, 12 of these last for less than 90% of the sample period; 7 last for less than 70% of the sample period; and 4 for less than 50%. When one defines as evidence for the PS hypothesis that a negative trend lasts for at least 70% of the sample period, the evidence is clearly less than overwhelming in favor of the PS hypothesis. Moreover, the measures of the prevalence of a positive trend $\psi(+)$ and trendless behavior $\psi(.)$ are greater than 0 in the case of 7 and 17 price series respectively. Of those series that present trendless segments, only 9 last for less than 90% of the sample period; 8 last for less than 60%; and 7 last for less than 50%. The number and prevalence of these trendless segments in commodity price series provides additional evidence to weaken any case favoring the PS hypothesis.

In summary, for the vast majority of commodities, the trend is not well represented by a single downward slope; strikingly, the most common value of the trend in a given segment is zero; and the PS hypothesis is clearly not as pervasive as previous work (see Leon and Soto, 1997) has presented. The variability of the trend suggests that forecasting of commodity prices should not typically occur about a single trend.

The above analysis suggests a refocusing of effort in the literature away from assessing the PS hypothesis to asking the less discussed question – what's generating the local trends? Consistent with the earlier literature, the first break date for most of the commodities in which a break in either intercept or trend cannot be rejected, is found just prior to or just after 1920. The exceptions are lead, silver, aluminum and jute (where the break date is found from the late 1930s to the mid-1940s); and palmoil and rice (break dates in the 1980s). With respect to the second break date, breaks in the deterministic trend occur in the 1920s (zinc), 1930s (rubber and timber), mid-1940s to mid-1950s (hide, tea and wool), 1970s (maize, silver and tin) and the 1980s (lead). For the sake of illumination consider those commodities that experienced a second break in the deterministic trend during the 1970s. There are some interesting joint characteristics of these series. Firstly the trend slope is significantly negative after the second break indicating a steep trend decline in prices after the second break. For 2 of the 3 commodities (silver and tin) this steep decline is accompanied by an initial ratcheting up of relative prices at the break point indicated by significantly positive constant dummy. In fact for these 2 commodities their highest relative price occurs on or around this second break point. In some sense then the steep falls in price after the break point appear as a reaction to the high levels obtained beforehand. Further analysis is of course warranted but this is illustrative for our general purposes.

Of course the variability of commodity prices has just as much policy relevance as the trend. Many developing countries depend on one or two primary commodity products for the majority of their export earnings. Movements in price, whether transitory or permanent, can seriously effect income and consumption levels of commodity exporting countries. The success of stabilization policies designed to smooth income flows depends upon the time series properties of commodity price series. For example it is very difficult to stabilize income around a declining trend, likewise it is also difficult to stabilize income when movements in price are large and/or continue for a long time.

In part, variability is related to the characterization of shocks as TS or DS. Relative commodity prices that are TS will have shocks that are transitory, dissipating completely over time. DS processes, on the other hand, allow the decomposition of shocks into both permanent and transitory components (see Beveridge and Nelson, 1981). Clearly, given the finite horizon of shocks, TS commodities may be amenable to stabilization. In contrast, the permanent component in the price innovations of DS commodities may motivate adjustment to long-run levels of income and consumption.

Given that 14 commodities are found to be TS after allowance for (up to) two breaks, the results of this paper suggest that the majority of relative commodity prices experience innovations that are entirely transitory. Interestingly, Table 7 gives the application of bootstrapped unit root test critical values that *do not* account for temporal dependence, wherein 5 more series are adjudged, at the 10% level at least, TS after allowing for breaks (cocoa, cotton, sugar, tobacco and wheat). Previous studies (for example, Leon and Soto, 1997) have not allowed for temporal dependence and concluded that commodity series are overwhelmingly TS after allowing for breaks. Our results suggest that accounting for temporal dependence implies that stabilization policies may not be as applicable as previously suggested (see Leon and Soto, 1997). In any case, even if a series is designated TS, this does not mean that stabilization policy is feasible. As already noted a transitory shock may still last a very long time and/or be very large, meaning that income smoothing is impractical as a policy objective.

5. Conclusion

The purpose of this paper is to investigate the prevalence of trends in the behavior of relative commodity prices. These issues have clear policy implications for the developing countries which produce primary commodities. For instance the finding of a long-run negative trend in prices predicted by the Prebisch-Singer hypothesis has often motivated diversification into manufactures. The recent literature has adopted unit root testing procedures and time series modeling to assess the trend behaviour of commodity prices (see, *inter alia*, Leon and Soto, 1997). In a direct extension to that work, this paper applies the Lumsdaine and Papell (1997) unit root testing methodology where the alternative hypothesis is a trend-stationary process with two endogenously chosen break dates. In an effort to compare results with previous studies, an extended series of the original Grilli and Yang (1988) index is employed, where each nominal commodity price (in US dollars) is deflated by the United Nations Manufactures Unit Value (MUV) index. The data set covers the period 1900-1998 and comprises 24 commodities.

As a further extension of previous work we note that there is substantial serial correlation observed in many commodity prices. We thus employ a parametric bootstrap to generate appropriate unit root test critical values which allow for temporal dependence under the null hypothesis of a unit root. Allowing for temporal dependence under the unit root null generally results in larger critical values. Thus, previous studies that employ either asymptotic critical values or use bootstrap critical values that do not account for temporal dependence under the unit root null will tend to reject a true unit root null too often

Using the Lumsdaine and Papell (1997) unit root testing methodology where the alternative hypothesis is a trend-stationary process with two endogenously chosen break dates we find that 10 of the commodity price series analyzed are characterized as trend-stationary with two structural breaks in intercept and trend, with an additional 4 being characterized as trend stationary with one structural break in intercept and trend. Series are then modeled using the appropriate ARMA methodology augmented by breaks in both the trend and the constant. This segmentation of the trend allows us clarify whether any downward trend is a long-run phenomenon or more local event. The results are striking; although 15 commodity prices are shown to exhibit at least one negative trend, a measure of the prevalence of a negative trend reveals that in only 8 cases does the deterioration exists for at least 70% of the sample period. Thus, in contrast to previous work that has allowed for structural breaks (see Leon and Soto, 1997), this paper provides little support for the Prebisch-Singer hypothesis. The trend variability shown suggests that forecasters of commodity prices should not typically attempt analysis around a single long-run trend. Importantly, these results are shown to be robust to recent criticism (see Kim *et al.,* 2003) that estimated trend values are dependent on specifying series as difference or trend-stationary.

While this paper is primarily concerned with an examination of a shifting trend in commodity prices, some information regarding the variance can also be ascertained. The trend-stationary nature of commodity prices of the majority of series examined indicates that shocks are finite raising the possibility of successful stabilization policies. Of course the success of such policies depends not only on the finite nature of shocks but on their amplitude and the length of time taken to revert to equilibrium. Conventional tests of the latter (e.g. half-lives), such as the approximately median-unbiased method of Andrews and Chen (1994), are not robust to series with structural breaks and we suggest this as an avenue of future work.

Finally, it is generally noted that a single trend fitted to a commodity price series is dominated by the variance. Given the notion of trend variability presented in this paper we suggest it might be useful for future work to move away from examining the Prebisch-Singer hypotheses. Instead theoretical and empirical work should concentrate on examining the extent and causes of local trends. This may provide a better picture of the forces that shape commodity prices over the medium-run to emerge and provide much needed input into policy decisions.

Endnotes

[1] Spraos (1980), Sapsford (1985), Thirwall and Bergevin (1985), Grilli and Yang (1988) and Powell (1991) report results that suggests that there has been a deterioration in the terms of trade of commodity exporting developing countries, although not to the extent emphasized in Prebisch (1950) and Lewis (1952).

[2] Similarly Cuddington and Urzua (1989) found no deterioration in the terms of trade, but instead found that commodity prices fluctuated secularly around a stable trend. For other studies on the long-run trends in commodity prices see Powell (1991), Bleaney and Greenaway (1993), Labys (1993), Gafer (1995), Bloch and Sapsford (1997), Newbold and Vougas (1996) and Newbold, Rayner, and Kellard (2000). A good summary of this literature can be found in Greenaway and Morgan (1999).

[3] For example Sapsford (1985), Cuddington and Urzua (1989), Ardeni and Wright (1992), Sapsford, Sarkar, and Singer (1992), and Reinhart and Wickham (1994) examine trends in aggregate commodity price indexes, while Cuddington (1992), Leon and Soto (1997) and Badillo, Labys, and Wu (1999) examine trends in individual commodity prices.

[4] Zanias (2004) applies the Lumsdaine and Papell (1997) unit root test to the Grilli-Yang index, but not the underlying 24 commodities. After accounting for two structural breaks in the level of the series, the index is found to be stationary and thus follows neither a negative deterministic or stochastic trend.

[5] Cuddington *et al.* (2002, 2003), examining the PS hypothesis, also allows for (up to) two breaks in primary commodity prices. The calculation of bootstrapped unit root test critical values in our paper provides an extension to such work. We thank an anonymous referee for this point.

[6] Longer lag lengths of *k* lagged differences did not change the results. We selected 5 lagged differences because 4 lagged differences was the largest selected for any of the commodity prices. We use a lag of 5 for all combinations of lag length and break dates so as to compare over equal sample periods for each specification. After the lag length has been selected for a particular commodity, that lag length was then used in the final stage of estimation so as to maximize the number of observations.

[7] Ng and Perron (1995) demonstrate that an under-parameterized model can have large size distortions, while an over-parameterized model may have low power. But the size problem is more severe than power loss. They show that methods based on sequential tests have an advantage over both the Said and Dickey (1984) fixed-rule and information-based rules such as the Akaike information criterion and the Schwarz information criterion, because the former have less size distortions and have comparable power. The procedure adopted in this paper falls into this category of the general-to-specific sequential procedures.

[8] There is a long running debate in the literature surrounding the calculation of critical values for structural breaks. Given a stationary DGP, standard critical values can be adopted when breaks are considered exogenous or *known*. On the other hand, when employing a structural change test for stationary series, such as the Bai and Perron (1998) methodology, to identify *unknown* breaks,

bootstrapped critical values are required to allow for the possibility of data mining (although there is still some debate here too; see, for example, Hendry (1999) on the difficulty of detecting structural breaks in small samples). Our case is complicated because we do not know a priori whether commodity series contain a unit root. During the unit root tests, structural breaks are estimated from the data and thus if the null is rejected it appears sensible to undertake further analysis under the specification of the alternative. However, the alternative we suggest, treats the breaks as *known* and therefore that the significance of each regime can be acceptably examined by comparing the relevant t-statistic against standard critical values. Additionally and following Hendry (1999), the use of such critical values delivers the highest possible power for the test used, imperative considering the small sample size employed. We thank David Harvey and Christian Murray for their help on this issue.

[9] Specifically, in a two break model:

$$y_t^* = y_t; \quad if\ 1 \le t \le TB1$$
$$y_t^* = y_t - \beta_1(TB1 - 1899); \quad if\ TB1 + 1 \le t \le TB2$$
$$y_t^* = y_t - \beta_1(TB1 - 1899) - \beta_3(TB2 - TB1); \quad if\ t > TB2$$

[10] SBC is the most commonly applied model selection criterion for ARMA processes. It is known to yield consistent estimators of (*p,q*) if the true model is in the set considered.

[11] The exception is rubber where the moving average coefficients sum to –0.085.

[12] This cannot be taken as conclusive evidence of trend-stationarity, since it is well known that maximum likelihood can often yield estimates on the boundary of the invertibility region even

when the true parameter values are well within that region (see, for example, Cryer and Ledolter, 1981, and Shephard and Harvey, 1990). Nevertheless, it is difficult in these circumstances to see what the analyst can do other than proceed with the TS model.

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[13] We thank an anonymous referee for this useful suggestion.

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Table 1: Ng and Perron (2001) unit root tests for the logarithm of relative commodity prices, 1900-1998.

Commodity	MZ_ρ	MZ_t	k
Aluminum	-11.65	-2.37	2
Banana	-5.60	-1.59	0
Beef	-12.33	-2.40	0
Cocoa	-7.86	-1.98	2
Coffee	-12.16	-2.41	2
Copper	-7.05	-1.85	5
Cotton	-3.66	-1.19	3
Hide	-19.33^{**}	-3.11^{**}	2
Jute	-8.37	-1.90	4
Lamb	-14.56	-2.68	0
Lead	-17.80^{**}	-2.92^{**}	0
Maize	-3.55	-1.19	5
Palmoil	-10.43	-2.28	5
Rice	-13.77	-2.60	4
Rubber	-18.79^{**}	-2.92^{**}	0
Silver	-6.27	-1.77	2
Sugar	-17.02	-2.90	5
Tea	-9.10	-2.11	2
Timber	-18.77^{**}	-3.06^{**}	0
Tin	-16.26	-2.76	0
Tobacco	-2.34	-0.92	4
Wheat	-12.84	-2.47	4
Wool	-2.32	-0.88	4
Zinc	-29.24^{**}	-3.81^{**}	0

Notes: The above statistics are derived on the basis of GLS detrending, thus, the alternative hypothesis is trend stationarity. The 5% critical values for the Ng and Perron (2001) $\underline{MZ_\rho}$ and $\underline{MZ_t}$ are -17.3 and -2.91 respectively. k is the augmented lag length chosen by the Ng and Perron (2001) modified AIC. Rejection of the null of unit root at the 5% significance level is designated with **. It should be further noted that the unit root tests reported in Table 1 ignore the possibility of breaks in either the level or trend of the series.

Table 2: Unit root tests (allowing for shifts in the deterministic trends) for the logarithm of relative commodity prices, 1900-1998.

Commodity	*TB1*	*ρ*	*k*	*Bootstrapped Critical Values for ρ*		
	TB2	*(t-stat)*		*1%*	*5%*	*10%*
2 Breaks in Intercept and 2 Breaks in Trend						
Hide	1919	-0.800^{*}	1(2)			
	1951	(-7.93)		-8.95	-8.12	-7.77
Lead	1946	-0.649^{**}	1(0)			
	1981	(-6.86)		-7.41	-6.83	-6.52
Maize	1920	-0.752^{*}	1(0)			
	1976	(-6.85)		-7.51	-6.91	-6.58
Rubber	1924	-0.428^{*}	0(0)			
	1932	(-6.58)		-7.43	-6.78	-6.49
Silver	1939	-0.656^{**}	1(0)			
	1978	(-7.15)		-7.30	-6.81	-6.46
Tea	1921	-0.633^{***}	1(0)			
	1952	(-7.51)		-7.32	-6.86	-6.57
Timber	1914	-0.629^{*}	3(0)			
	1938	(-6.70)		-7.25	-6.78	-6.51
Tin	1918	-0.526^{*}	1(0)			
	1975	(-6.52)		-7.54	-6.78	-6.50
Wool	1916	-1.264^{*}	3(2)			
	1949	(-7.99)		-9.08	-8.19	-7.86
Zinc	1914	-0.730^{***}	1(0)			
	1921	(-8.02)		-7.51	-6.76	-6.47

Commodity	*TB1*	ρ *(t-stat)*	*k*	*Bootstrapped Critical Values for* ρ *1%*	 *5%*	 *10%*
1 Break in Intercept and 1 Break in Trend						
Aluminum	1941	-0.395*	1(2)			
		(-5.75)		-6.97	-6.07	-5.52
Jute	1945	-0.536*	1(2)			
		(-5.90)		-7.00	-5.98	-5.53
Palmoil	1985	-0.502*	1(2)			
		(-5.95)		-7.02	-6.00	-5.53
Rice	1981	-0.464**	1(2)			
		(-6.11)		-7.00	-5.98	-5.53

Notes: TB1 and TB2 correspond to the first and second break dates. TB1 represents the end date of the first regime while TB2 represents the end date of the second regime. Figures in parentheses are t-statistics. The number of lagged differences in the ADF regression is given in the column designated as *k*. The subsequent number in parentheses is the number of lagged differences used under the unit root null to generate a unit root series with temporal dependence. This lag was selected using the SBC criterion.

*** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level.

Table 3: Estimated trend-stationary models with breaks

	HIDE	LEAD	MAIZE	RUBBER	SILVER
TB1	1919	1946	1920	1924	1939
TB2	1952	1981	1976	1932	1978
$\hat{\alpha}$	4.77 (41.0)	4.42 (243.0)	4.86 (42.5)	7.15 (27.6)	4.18 (43.9)
$100\hat{\beta}_1$	3.56 (3.71)	-0.28 (-3.98)	3.34 (3.76)	-6.44 (-3.91)	-1.67 (-4.26)
$\hat{\alpha}+\hat{\beta}_2$	4.08 (19.6)	4.82 (97.3)	4.44 (22.8)	8.61 (26.4)	3.82 (22.5)
$100(\hat{\beta}_1+\hat{\beta}_3)$	-0.07 (-0.16)	-1.03 (-7.70)	-0.42 (-1.98)	-31.2 (-7.63)	2.95 (7.40)
$\hat{\alpha}+\hat{\beta}_4$	3.66 (14.7)	4.50 (41.9)	4.21 (17.7)	9.40 (30.2)	4.49 (19.9)
$100(\hat{\beta}_1+\hat{\beta}_5)$	0.03 (0.10)	-1.09 (-1.52)	-2.72 (-3.27)	-1.84 (-4.13)	-7.47 (-7.37)
p, q	0, 1	2, 1	0, 1	2,1	1, 0

Note: Figures in parentheses are t-statistics.

Table 3 continued: Estimated trend-stationary models with breaks

	TEA	TIMBER	TIN	WOOL	ZINC
TB1	1921	1914	1918	1916	1914
TB2	1952	1938	1975	1949	1921
$\hat{\alpha}$	4.91 (57.8)	3.43 (28.9)	3.41 (24.7)	5.32 (48.5)	4.57 (34.7)
$100\hat{\beta}_1$	-2.54 (-4.01)	1.92 (1.59)	2.32 (2.04)	0.74 (0.71)	1.11 (0.80)
$\hat{\alpha}+\hat{\beta}_2$	5.34 (35.5)	3.81 (21.2)	2.95 (13.5)	5.62 (30.1)	5.64 (25.8)
$100(\hat{\beta}_1+\hat{\beta}_3)$	0.06 (0.16)	-1.22 (-1.99)	1.40 (5.81)	-1.13 (-2.85)	-25.3 (-7.70)
$\hat{\alpha}+\hat{\beta}_4$	5.80 (31.7)	4.13 (19.7)	3.40 (10.0)	5.88 (26.9)	6.08 (22.7)
$100(\hat{\beta}_1+\hat{\beta}_5)$	-3.15 (-14.5)	0.74 (4.17)	-4.53 (-5.16)	-3.56 (-15.9)	0.27 (1.92)
p, q	0, 1	1, 1	0, 2	0, 1	1, 0

Note: Figures in parentheses are t-statistics.

Table 3 continued: Estimated trend-stationary models with breaks

	ALUMINUM	JUTE	PALMOIL	RICE
TB1	1941	1945	1985	1981
TB2	N/A	N/A	N/A	N/A
$\hat{\alpha}$	6.13 (49.0)	4.80 (51.0)	4.99 (61.1)	5.14 (61.9)
$100\hat{\beta}_1$	-2.48 (-5.37)	0.02 (0.07)	-0.56 (-3.47)	-0.54 (-3.16)
$\hat{\alpha}+\hat{\beta}_2$	5.77 (31.2)	5.30 (30.5)	4.01 (19.9)	4.67 (28.9)
$100(\hat{\beta}_1+\hat{\beta}_3)$	0.09 (0.31)	-1.90 (-6.83)	4.28 (1.93)	-1.28 (-0.93)
$\hat{\alpha}+\hat{\beta}_4$	N/A	N/A	N/A	N/A
$100(\hat{\beta}_1+\hat{\beta}_5)$	N/A	N/A	N/A	N/A
p, q	1, 1	0, 1	2, 0	1, 1

Note: Figures in parentheses are t-statistics.

Table 4: Estimated difference-stationary models

	BANANA	BEEF	COCOA	COFFEE	COPPER
$100\hat{\beta}$	0.058 (0.06)	0.850 (0.38)	-0.801 (-0.40)	-0.015 (-0.01)	-0.811 (-0.49)
p, q	0, 0	0, 0	2, 0	0, 0	0, 0

Note: Figures in parentheses are t-statistics.

Table 4 continued : Estimated difference-stationary models

	COTTON	LAMB	SUGAR	TOBACCO	WHEAT
$100\hat{\beta}$	-0.859 (-0.85)	1.55 (0.647)	-0.973 (-3.80)	0.980 (1.41)	-0.879 (-6.36)
p, q	3, 0	0, 0	1, 2	4, 0	0, 4

Note: Figures in parentheses are t-statistics.

Table 5: Estimated difference-stationary models with breaks

	HIDE	LEAD	MAIZE	RUBBER	SILVER
TB1	1919	1946	1920	1924	1939
TB2	1952	1981	1976	1932	1978
$100\hat{\beta}_1$	3.55 (3.66)	-0.28 (-4.68)	3.34 (3.73)	-5.55 (-1.63)	-1.68 (-4.12)
$\hat{\beta}_2$	-0.69 (-5.03)	0.40 (12.7)	-0.42 (-3.46)	1.59 (9.05)	-0.36 (-2.97)
$100(\hat{\beta}_1+\hat{\beta}_3)$	-0.08 (-0.18)	-1.04 (-8.87)	-0.42 (-1.97)	-36.6 (-5.77)	2.95 (7.13)
$\hat{\beta}_4$	-0.41 (-3.61)	-0.30 (-3.94)	-0.23 (-1.82)	0.68 (4.19)	0.67 (4.86)
$100(\hat{\beta}_1+\hat{\beta}_5)$	0.03 (0.09)	-1.26 (-1.79)	-2.73 (-3.25)	-1.26 (-0.62)	-7.42 (-7.11)
p, q	0, 2	3, 2	0, 2	1, 2	1, 1

Note: Figures in parentheses are t-statistics.

Table 5 continued: Estimated difference-stationary models with breaks

	TEA	TIMBER	TIN	WOOL	ZINC
TB1	1921	1914	1918	1916	1914
TB2	1952	1938	1975	1949	1921
$100\hat{\beta}_1$	-2.41 (-5.99)	2.34 (2.06)	2.30 (1.99)	1.09 (1.60)	1.10 (0.77)
$\hat{\beta}_2$	0.43 (6.84)	0.30 (2.78)	-0.45 (-3.44)	0.20 (2.69)	1.08 (6.77)
$100(\hat{\beta}_1 + \hat{\beta}_3)$	-0.13 (-0.78)	-1.13 (-2.86)	1.40 (5.74)	-0.91 (-5.74)	-25.3 (-7.58)
$\hat{\beta}_4$	0.54 (14.5)	0.35 (6.25)	0.43 (3.28)	0.23 (6.13)	0.44 (3.35)
$100(\hat{\beta}_1 + \hat{\beta}_5)$	-3.29 (-47.9)	0.68 (13.3)	-4.51 (-5.07)	-3.55 (-53.8)	0.27 (1.84)
p, q	2, 2	4, 2	0, 3	1, 3	1, 1

Note: Figures in parentheses are t-statistics.

Table 5 continued: Estimated difference-stationary models with breaks

	ALUMINUM		JUTE		PALMOIL		RICE	
TB1	1941		1945		1985		1981	
TB2	N/A		N/A		N/A		N/A	
$100\hat{\beta}_1$	-2.50	(-5.20)	0.24	(0.31)	-0.57	(-3.17)	-0.54	(-3.06)
$\hat{\beta}_2$	-0.34	(-3.05)	0.42	(2.74)	-0.92	(-5.23)	-0.47	(-3.61)
$100(\hat{\beta}_1 + \hat{\beta}_3)$	0.08	(0.25)	-1.74	(-2.44)	3.87	(1.69)	-1.27	(-0.91)
$\hat{\beta}_4$	N/A		N/A		N/A		N/A	
$100(\hat{\beta}_1 + \hat{\beta}_5)$	N/A		N/A		N/A		N/A	
p, q	1, 2		0, 2		1, 2		1, 2	

Note: Figures in brackets are estimated t-statistics.

Table 6: Relative measures of the prevalence of a trend

	ALUMINUM	BANANA	BEEF	COCOA	COFFEE	COPPER	COTTON	HIDE
$\psi(-)$	0.424	0.000	0.000	0.000	0.000	0.000	0.000	0.000
$\psi(+)$	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.202
$\psi(.)$	0.576	1.000	1.000	1.000	1.000	1.000	1.000	0.798

	JUTE	LAMB	LEAD	MAIZE	PALMOIL	RICE	RUBBER	SILVER
$\psi(-)$	0.535	0.000	0.828	0.788	0.869	0.828	1.000	0.606
$\psi(+)$	0.000	0.000	0.000	0.212	0.131	0.000	0.000	0.394
$\psi(.)$	0.464	1.000	0.172	0.000	0.000	0.172	0.000	0.000

	SUGAR	TEA	TIMBER	TIN	TOBACCO	WHEAT	WOOL	ZINC
$\psi(-)$	1.000	0.687	0.253	0.232	0.000	1.000	0.828	0.071
$\psi(+)$	0.000	0.000	0.596	0.768	0.000	0.000	0.000	0.778
$\psi(.)$	0.000	0.313	0.151	0.000	1.000	0.000	0.172	0.151

Table 7: Unit root tests (allowing for shifts in the deterministic trends) for the logarithm of relative commodity prices, 1900-1998.

Commodity	*TB1* *TB2*	ρ *(t-stat)*	*k*	*Bootstrapped Critical Values for* ρ *1%*	 *5%*	 *10%*
2 Breaks in Intercept and 2 Breaks in Trend						
Cocoa	1946	-0.584*	1			
	1975	(-6.68)		-7.47	-6.80	-6.46
Cotton	1929	-0.668***	2			
	1949	(-7.52)		-7.39	-6.79	-6.42
Sugar	1923	-0.640*	0			
	1971	(-6.65)		-7.47	-6.86	-6.59
Tobacco	1919	-0.659**	4			
	1980	(-6.95)		-7.32	-6.76	-6.50
Wheat	1913	-0.798**	4			
	1945	(-7.23)		-7.50	-6.90	-6.57

Notes: TB1 and TB2 correspond to the first and second break dates. TB1 represents the end date of the first regime while TB2 represents the end date of the second regime. Figures in parentheses are t-statistics. The number of lagged differences in the ADF regression is given in the column designated as *k*. The bootstrapped unit root test critical values *do not* account for temporal dependence

*** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level.

Figure 1

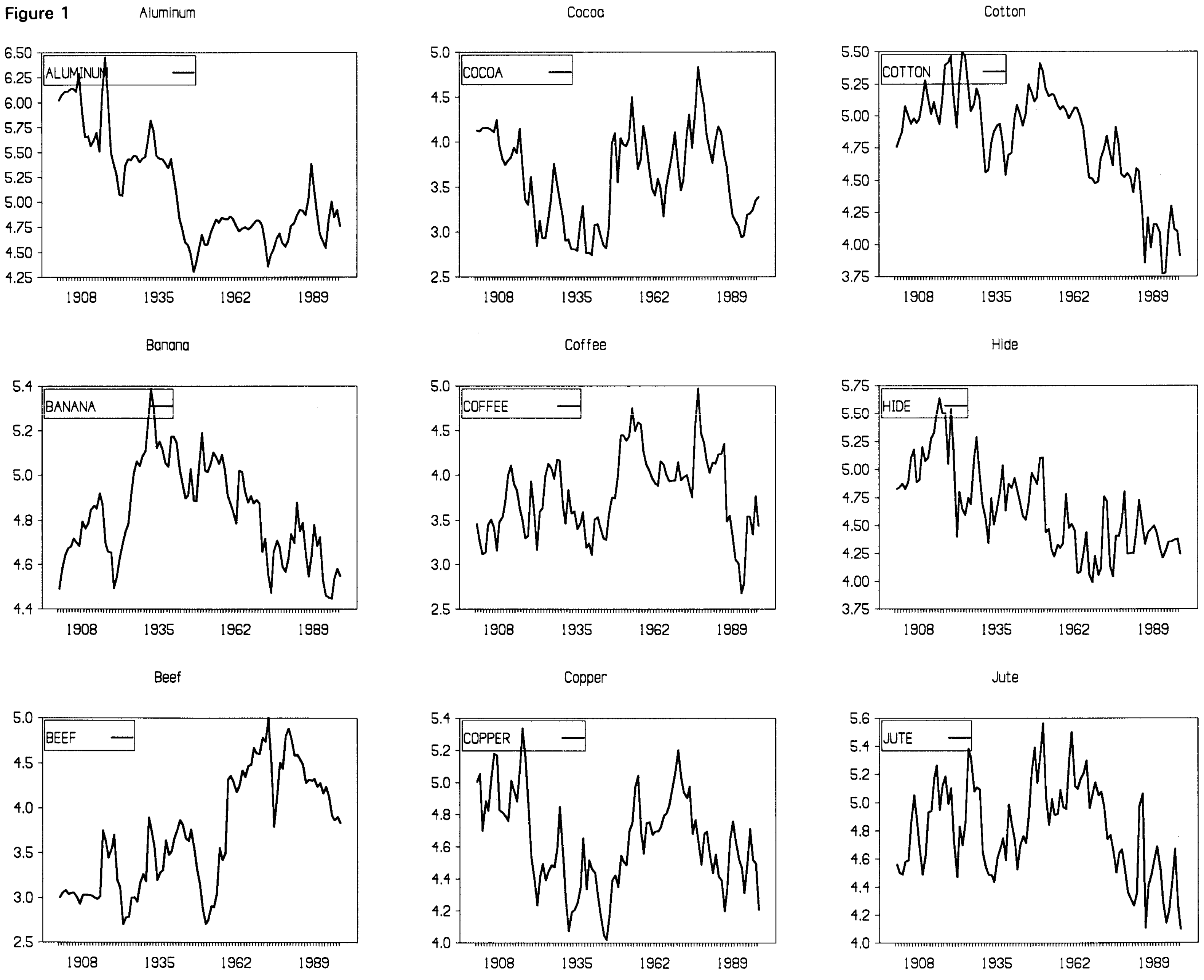
Aluminum
ALUMINUM
6.50
6.25
6.00
5.75
5.50
5.25
5.00
4.75
4.50
4.25
1908
1935
1962
1989
Cocoa
COCOA
5.0
4.5
4.0
3.5
3.0
2.5
1908
1935
1962
1989
Cotton
COTTON
5.50
5.25
5.00
4.75
4.50
4.25
4.00
3.75
1908
1935
1962
1989
Banana
BANANA
5.4
5.2
5.0
4.8
4.6
4.4
1908
1935
1962
1989
Coffee
COFFEE
5.0
4.5
4.0
3.5
3.0
2.5
1908
1935
1962
1989
Hide
HIDE
5.75
5.50
5.25
5.00
4.75
4.50
4.25
4.00
3.75
1908
1935
1962
1989
Beef
BEEF
5.0
4.5
4.0
3.5
3.0
2.5
1908
1935
1962
1989
Copper
COPPER
5.4
5.2
5.0
4.8
4.6
4.4
4.2
4.0
1908
1935
1962
1989
Jute
JUTE
5.6
5.4
5.2
5.0
4.8
4.6
4.4
4.2
4.0
1908
1935
1962
1989

Figure 1
(Con't)

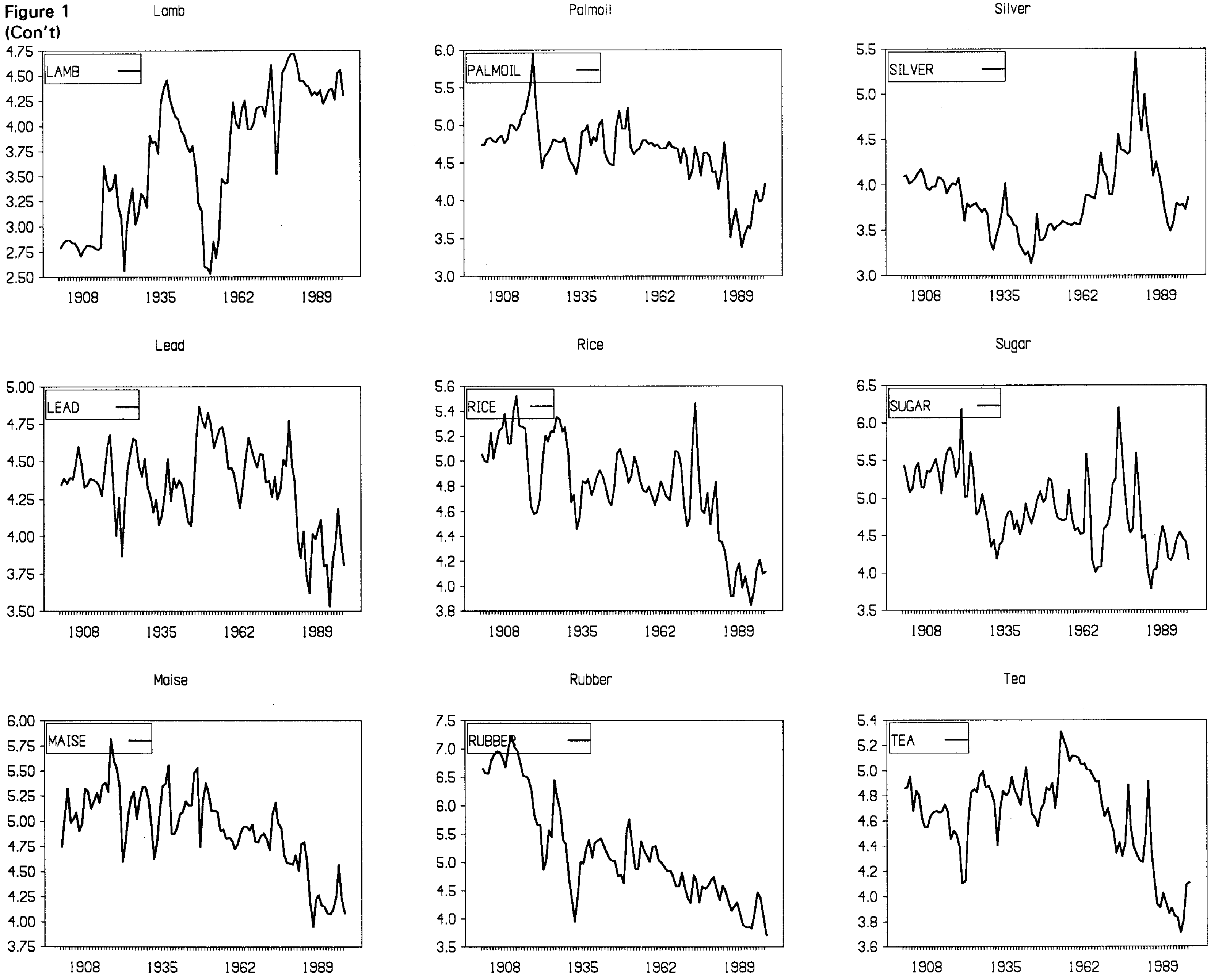

Figure 1
(Con't)