Trends and Persistence 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 and the nature of their persistence. Employing a dataset of 24 relative commodity prices for the 1900-98 period, the pervasiveness of the Prebisch-Singer hypothesis is shown to be a function of a priori selected decision criteria, providing an explanation of conflicting findings in the recent literature. Moreover, much less persistence is found in the relative commodity prices than previously reported, since 23 out of the 24 commodities can be classified as trend-stationary. This implies there may well be more room for stabilization and price support mechanisms than previously advocated.

JEL classification: O13, C22.

Keywords: Primary commodities, unit root tests, structural breaks.

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. This has become known as the Prebisch-Singer (PS) hypothesis. Some of the explanations that have been offered for this decline 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.

Recent empirical studies, using more advanced econometric techniques which permit commodity prices to contain a stochastic trend, have found evidence against the PS hypothesis. Notably, Cuddington (1992) examines the 24 commodities that comprise the Grilli-Yang index (plus oil and coal) and found that 13 of these 26 commodities can be modeled as difference-stationary (DS) processes for the period 1900-1983, with the remainder being modeled as trend-stationary (TS) processes. Just five of the TS models had the negative trend predicted by the PS hypothesis, while the other TS models had zero or positive trends. Overall, 21 of the 26 commodity prices exhibited a zero or positive trend, implying a strong rejection of the PS hypothesis in most cases[1].

Another issue in modeling commodity prices as stochastic trends relates to the persistence of shocks. Knowing whether shocks to commodity prices leave permanent or transitory imprints is important for the design of both short-run and long-run policies. The design of structural adjustment programs (to improve depressed terms of trade) will be different depending on whether export prices are expected to remain low for a short or long period of time. Furthermore, optimal management of stabilization policies depends, to an important degree, on the nature of the shock to commodity prices and the speed with which the shocks dissipate (Engel and Meller, 1993).

A number of studies have investigated the possibility of shifting deterministic trends in international commodity prices[2]. Specifically, Leon and Soto (1997) extend the methodology of Cuddington (1992) by applying formal tests 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, 20 are classified as TS models for the 1900-92 period, implying that shocks to commodity prices are, in several cases, less persistent than suggested by Cuddington (1992). 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 Cuddington (1992) and Leon and Soto (1997), this study provides new evidence by applying the Lumsdaine and Papell (1997) unit root test to individual commodity prices in an extended Grilli-Yang index. Allowing for the possibility of two endogenously determined break dates, even less persistence is found in the 24 relative commodity prices, over the period 1900-1998, than previous studies. We also reconcile the different results found in the literature. For example, Cuddington (1992), not allowing for the possibility of trend breaks, noted that only five commodity price series exhibited a negative trend over the entire sample period (evidence against the PS hypothesis). In contrast, Leon and Soto (1997) allow for the possibility of one trend break and find 17 of their series contain a negative trend and thus, claim support for the PS hypothesis. In this paper, when we allow for two breaks in deterministic trend components we find that 16 of the 24 commodity prices present a significant negative trend; twelve for at least 50% of the sample period; eight for at least 75% of the sample period; and 5 for at least 85% of the sample period. These results suggest that the pervasiveness of the PS hypothesis is a function of a priori selected decision criteria and goes a long way to explain the conflicting results in the literature.

The remainder of the paper is organized as follows: Section 2 describes the empirical estimation methodology. Section 3 presents the data, the empirical results and a novel measure of negative trend persistence. Finally, section 4 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 structural breaks are ignored and/or if the incorrect number of breaks is considered. Perron (1989) allows for an 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) extended 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 structural breaks in both the intercept and trend terms determined endogenously. The null and alternative hypotheses are given as follows:

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

where y_t is the logarithm of the relative commodity price, t is a linear trend, β and 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 assumed to satisfy the following conditions:

$$\delta T \leq TB1,\ TB2 \leq (1-\delta)T \ \ and \ \ |TB1 - TB2| \geq 2, \tag{2}$$

where T is the length of the data series and δ is a trimming parameter set at 0.05. 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:

$$\begin{aligned} \Delta y_t &= \mu_0 + \beta t + \rho y_{t-1} + \mu_1 D_{L1,t} + \mu_2 D_{T1,t} + \mu_3 D_{L2,t} + \mu_4 D_{T2,t} \\ &\quad + \sum_{j=1}^{k} d_j \Delta y_{t-j} + \varepsilon_t \end{aligned} \tag{3}$$

where ρ and d_j are coefficient parameters[3]. 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.

The optimal lag length k is determined by estimating equation (3) without the four dummy variables. A general-to-specific method is employed starting with k_{max} equal to 5. 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[4]. After determining the optimal lag length, equation (3) is estimated for all combinations of two breaks. The selection of break dates *TB1* and *TB2* correspond to the equation that yields the largest t-statistic (in absolute value) associated with the coefficient ρ.

As asymptotic critical values are often misleading in small samples, critical values are computed for these test statistics using a bootstrap procedure. Five hundred pseudo-samples are generated from a random walk model with drift. For each pseudo-sample, the procedure outlined above is carried out. The largest t-statistic for ρ for each pseudo-sample is tabulated. The 1%, 5%, and 10% critical values are obtained from the empirical distribution of these *t*-statistics. These values are similar to those computed by Lumsdaine and Papell (1997); see their Table 3.

As previously noted, the most general model considered allows for two breaks in both intercept and trend. If the unit root cannot be rejected, a sequential trend reduction methodology is conducted, in which time trend and level shift dummy variables that are insignificant are eliminated (beginning with time trend dummies) from the tests and the unit root tests are repeated with the more restricted model. We continue in this fashion until the unit root null is rejected. This procedure is employed as an over-parameterized model (i.e. one which includes a trend when it is actually not present) will result in low power (smaller probability of rejecting the unit root null). Thus, a general-to-specific methodology allows a more appropriate characterization of the data series.

After assessing the level of integration of each price series, it is then possible to model the relevant data generating process. The first generating process is represented by the TS model

$$y_t = \alpha + \beta t + u_t \tag{4}$$

where the random variable u_t is stationary with mean zero. The focal point of interest in equation (5) is in the slope parameter β. The alternative generating process for the data is represented by the DS model

$$\Delta y_t = \beta + v_t \tag{5}$$

where the generating process for v_t is stationary and invertible. In this framework, interest is in the drift parameter β. For both the TS model (4) and DS model (5), the error term is permitted to follow an ARMA (p,q) process

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

$$v_t - \phi_1 v_{t-1} - \ldots\ldots - \phi_p v_{t-p} = \varepsilon_t - \theta_1 \varepsilon_{t-1} - \ldots\ldots - \theta_q \varepsilon_{t-q} \tag{7}$$

where ε_t is zero-mean white noise. For each break date *TB*, the previously defined dummy variables were added to the right-hand side of (4) and (5). The parameters of (4) and (6), and (5) 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)[5], allowing all possible models with $p + q \leq 6$.

3. Data and Analysis of 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. It can be observed that some of the series trend downward while others have movements around a shifting mean.

Two of the most commonly used unit root tests in the literature are the augmented Dickey-Fuller test (ADF *r*-test) of Said and Dickey (1984) and Phillips-Perron test (PP *Z*-test) developed in Phillips and Perron (1988). It is well known that the ADF and PP tests have low power against local stationary alternatives. Elliot, Rothenberg and Stock (1996) (ERS *DF^{GLS})* develop a feasible point optimal test that relies on local GLS detrending to increase the power of the unit root tests.

A second serious problem associated with unit root testing is that the above named tests all suffer from serious size distortions when the data generating process (DGP) has a negative moving average terms. Schwert (1987, 1989), Phillips and Perron (1988), Pantula (1991), Ng and Perron (1995, 2001) and Perron and Ng (1996) demonstrate that the empirical size of conventional ADF and PP tests approach unity as the sum of the MA parameters in a univariate process approach negative one. Perron and Ng (1996) extend the work of Elliot, Rothenberg, and Stock (1996) by developing modified versions of the PP tests that have much better size properties than the conventional PP tests but also retain the power of the (ERS *DF^{GLS}).* These unit root tests are based on the local GLS detrending method and in addition use an autoregressive spectral density estimator of the long-run variance. The two tests are labeled the MZ_{ρ} and the MZ_t test. Ng and Perron (2001) suggest that these two tests have similar power to the DF^{GLS} test of Elliot, Rothenberg and Stock (1996) but also have superior size properties in the presence of MA disturbances. The decrease in size and increase in power are enhanced when one chooses the lag length based on the modified AIC criteria (MIC) developed in Ng and Perron (2001).

While the current paper is concerned primarily with unit root tests allowing for shifts in the deterministic components, for completeness, we report the MZ_{ρ} and the MZ_t statistics as well as the lag length *k*, selected using MIC. These results (which employ GLS detrending) are reported in Table 1. In an effort to examine the degree of persistence of these series, we report the value of one plus the coefficient on the lag level in the ADF regression ($1+\rho$). The OLS point estimates of $(1+\rho)$ are greater than or equal to 0.80 in 19 of the 24 commodities examined. Only hide, lead, sugar, timber, and zinc have values less than 0.80. Of course, these point estimates are biased downward. The unit root tests indicate that for all series, with the exception of hide, lead, robber, timber and zinc, the unit root hypothesis cannot be rejected. For the 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 shifting deterministic trend. We investigate this issue next.

Table 2 reports the unit root test results allowing for shifts in the deterministic trends. The first column reports the commodity. The second column contains the two break dates which refer to the end of the first and second regimes, respectively. The third column contains the coefficient estimates for lagged level of the data series, ρ. The coefficient on the time trend is reported in column four. The coefficient estimates for the intercept level shifts are reported in columns five and seven, μ_1 and μ_3, respectively. Columns six and eight report coefficient estimates for the trend slope coefficients, μ_2 and μ_4, respectively. Finally, the last column reports the number of lagged difference terms included.

For eight commodity prices (hide, lead, rubber, sugar, tea, timber, wool, and zinc) the null hypothesis of a unit root is rejected 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 in this fashion as with the idea that we can further refine the trend component, or 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 of the trend reduction method are reported for these remaining commodities. We find that the null is rejected in favor of the alternative of two breaks in intercept and one break in trend for cocoa, jute, and lamb. We find that for three commodities (aluminum, rice and tobacco), the null can be rejected in favor of the alternative of two breaks in the intercept only. Finally, we find that the null of unit root can be rejected in favor of the alternative of one break in intercept only for banana. Finally, for nine commodities (cocoa, cotton, jute, silver, maize, copper, wheat, lamb and tin) the unit root null cannot be rejected regardless of the specification of the alternative hypothesis.

To summarize, fifteen commodities are classified as TS, adopting a general-to-specific unit root testing procedure. Of these, fourteen are characterized as TS with two breaks (either trend or intercept) , emphasizing the importance of multiple break tests to commodity price series analysis. The recent literature is also divided as to the time series properties of commodity prices. Cuddington (1992), analyzing the Grilli-Yang index from 1900-1983 and employing unit root tests with no breaks, found a similar proportion of commodities were TS. However, Leon and Soto (1997), analyzing the Grilli-Yang index from 1900-1992 and employing unit root tests with one break, claim that 20 from 24 commodity prices are TS. Using a specific-to-general unit root testing procedure, eight of these are characterized as TS with one break.

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, cocoa, and lamb where the break date is found in the mid-1940s. With respect to the second break date, breaks in the deterministic trend occur in the 1930s (rubber, timber and aluminum), late 1940s or early to mid-1950s (hide, tea, wool, jute, and lamb), early 1970s (sugar) and the 1980s (lead, rice and tobacco).

In Table 3, relative commodity prices that are TS with two breaks in the intercept and trend are modeled by estimating equations (4) and (6). Tables 4, 5 and 6 conduct similar estimations for the other specifications. 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 15 TS models, 6 (lead, rubber, sugar, wool, aluminum, cocoa) contain a negative and significant trend (although in some cases not for the entire series), 2 (rice, banana) are trendless and 3 (tobacco, hide, lamb) contain a positive and significant trend. Interestingly, 4 series contain both positive and negative significant trends (tea, timber, zinc, jute).

Table 7 estimates *I(1)* models for those relative commodity prices which were found to exhibit unit root behavior. In 8 of the 9 cases, the estimated model strongly suggests over–differencing with the estimated moving average coefficients practically summing to one[6]. Consequently, Table 8 estimates the relevant TS models for completeness. Of the 9 commodities, only 1 (wheat) display a negative and significant stochastic trend, 2 (copper, coffee) are trendless and 1 (beef) contains a significantly positive trend. Five series contain both positive and negative significant trends (tin, maize, silver, cotton, palmoil).

To aid in gauging the relevance of the results to the PS hypothesis, a novel relative measure, ψ, of negative trend persistence is constructed for each commodity

$$\psi = \frac{\lambda}{N} \tag{8}$$

where $\lambda =$ number of years that a statistically significant negative trend exists and $N =$ total number of sample years. Table 9 displays the relative measure results for all commodities. The derived measure of negative trend persistence demonstrates that 16 of the 24 commodity prices present a significant negative trend; twelve for at least 50% of the sample period; eight for at least 75% of the sample period; and 5 for at least 85% of the sample period. Clearly, the pervasiveness of the Prebisch-Singer hypothesis is a function of a priori selected decision criteria and this may help to explain the conflicting results in the literature. For example, Cuddington (1992), who did not consider the possibility trend breaks, noted that only five price series contain a negative trend *for all* the 1900-83 period and concludes that the PS hypothesis, "…should certainly not be considered a universal phenomenon or stylized fact." However, Leon and Soto (1997), who allow for the possibility of one trend break, note that 17 of their series contain a negative trend *for all or most* of the 1900-92 period and thus claim that the PS hypothesis is, "…the case for most commodities."

The imprint of such shocks to commodity prices can be either permanent or temporary. Relative commodity prices that have unit roots imply that such shocks leave a permanent imprint on the series, while TS prices will have shocks that are temporary and dissipate over time. Persistence in the DS case is commonly defined as the ratio of the long-run effect of an innovation to its immediate effect (Campbell and Mankiw, 1987). If persistence is less than unity then the influence of a contemporary shock has a smaller impact on the long-run forecast than on the short-run forecast. Therefore, for all TS processes, persistence is zero, and in that sense very low persistence is taken to indicate behaviour that is "almost stationary".

The results from this paper suggest overwhelmingly that relative commodity prices have zero or close to zero persistence. Initially, fifteen commodities are classified as TS, adopting a general-to-specific unit root testing procedure. A further eight commodities have moving average parameters which suggest over-differencing when they are modeled as DS processes. Again this provides a contrast with the recent literature. Cuddington (1992) noted that 13 of 26 commodities contained a permanent component; estimating a gain function[7], these permanent components ranged from 0.34 of the innovation (wool) to one (beef, copper and rubber). Leon and Soto (1997) suggest that only four commodities contain a permanent component; ranged from 0.66 to 0.99 (silver, cocoa, bananas and beef). This paper indicates that only one commodity has a permanent component, measuring 0.12 (palmoil).

4. Conclusion

The purpose of this paper is to investigate trends and persistence in the behaviour 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 PS 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*, Cuddington,1992, and 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.

Our results indicate that 14 commodities are characterized as trend-stationary with two breaks, emphasizing the importance of multiple break tests to commodity price series analysis. Adopting a novel general-to-specific approach to unit root testing, overall, fifteen commodities are classified as trend-stationary. The long-run trend is then estimated by adopting the relevant ARIMA specification and appropriate dummies as indicated by the unit root analysis. Some 12 commodities have a negative time trend for 50% or more of the time, providing modest support for the PS hypothesis. However this result is sensitive to the decision criterion adopted and one should caution against any quick judgements as to the robustness of the PS hypothesis.

Finally, the level of persistence is examined. A relatively low level would indicate that forms of stabilization may be appropriate while higher levels of persistence motivate the need to adjust levels of consumption and investment. Apart from the 15 trend stationary commodities, eight of the nine remaining commodities indicated strong evidence of over-differencing in the ARIMA estimation. Overall, therefore, 23 of 24 commodities exhibit trend-stationary behaviour. Given that persistence is zero for trend-stationary processes, this indicates there is perhaps more room for stabilization and price support mechanisms than previously advocated. This is an issue for future research.

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). In contrast, 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), Newbold, Rayner, and Kellard (2000) and Kim *et al.*(2001). A good summary of this literature can be found in Greenaway and Morgan (1999).

[2] 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.

[3] Equation (3) is Lumsdaine and Papell's (1997) model CC which is based on the sequential Dickey-Fuller test procedure of Zivot and Andrews (1992).

[4] 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.

[5] 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.

[6] 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.

[7] Gain function, G, can be represented by $G = \dfrac{1-\theta_1-\theta_2-...-\theta_q}{1-\phi_1-\phi_2-...\phi_p}$.

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

Commodity	$(1+\rho)$	MZ_{ρ}	MZ_t	k
Aluminum	0.887	-11.65	-2.37	2
Banana	0.939	-5.60	-1.59	0
Beef	0.865	-12.33	-2.40	0
Cocoa	0.901	-7.86	-1.98	2
Coffee	0.856	-12.16	-2.41	2
Copper	0.889	-7.05	-1.85	5
Cotton	0.938	-3.66	-1.19	3
Hide	0.700	-19.33**	-3.11**	2
Jute	0.851	-8.37	-1.90	4
Lamb	0.838	-14.56	-2.68	0
Lead	0.797	-17.80**	-2.92**	0
Maize	0.906	-3.55	-1.19	5
Palmoil	0.834	-10.43	-2.28	5
Rice	0.819	-13.77	-2.60	4
Rubber	0.808	-18.79**	-2.92**	0
Silver	0.921	-6.27	-1.77	2
Sugar	0.693	-17.02	-2.90	5
Tea	0.891	-9.10	-2.11	2
Timber	0.785	-18.77**	-3.06**	0
Tin	0.817	-16.26	-2.76	0
Tobacco	0.971	-2.34	-0.92	4
Wheat	0.814	-12.84	-2.47	4
Wool	0.945	-2.32	-0.88	4
Zinc	0.645	-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) MZ_{ρ} and MZ_t are -17.3 and -2.91 respectively. $(1+\rho)$ is equal to one plus the coefficient on the lagged level in the ADF test. 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 **.

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

Commodity	*TB1* *TB2*	ρ	β	μ_1	μ_2	μ_3	μ_4	k
2 Breaks in Intercept and 2 Breaks in Trend								
Hide	1919	-0.687**	0.027*	-0.022*	-0.598*	-0.004	-0.404*	0
	1951	(-7.90)	(3.12)	(2.35)	(-4.89)	(-0.86)	(-4.34)	
Lead	1946	-0.572**	-0.002	0.281*	-0.004	-0.292*	0.010	0
	1981	(-6.58)	(-1.47)	(3.80)	(-1.50)	(-3.13)	(1.36)	
Rubber	1924	-0.409*	-0.036*	0.870*	-0.019*	0.985*	0.212*	0
	1932	(-6.32)	(-4.42)	(4.21)	(-4.99)	(6.03)	(5.60)	
Sugar	1923	-0.640**	0.010	-0.500*	-0.011	0.406*	-0.029*	1
	1971	(-6.65)	(0.98)	(-3.12)	(-1.06)	(2.59)	(-3.39)	
Tea	1921	-0.635**	-0.016*	0.318*	0.014*	0.349*	-0.019*	2
	1952	(-6.67)	(-2.70)	(3.98)	(2.11)	(5.70)	(-4.13)	
Timber	1914	-0.629**	0.015	0.320*	-0.027*	0.286*	0.017*	3
	1938	(-6.70)	(1.38)	(3.41)	(-2.41)	(4.50)	(4.15)	
Wool	1916	-1.319**	-0.004	0.406*	-0.009	0.349*	-0.034*	4
	1949	(-7.19)	(-0.29)	(3.72)	(-0.69)	(4.29)	(-5.79)	
Zinc	1914	-0.730**	0.002	0.767*	-0.201*	0.458*	0.201*	1
	1921	(-8.02)	(0.15)	(5.06)	(-6.24)	(4.34)	(6.71)	
Cotton	1929	-0.630	0.007	-0.410*	0.009	0.143	-0.030*	3
	1948	(-5.00)	(1.72)	(4.04)	(1.38)	(1.49)	(-4.78)	
Silver	1939	-0.633	-0.011*	-0.251*	0.031*	0.446*	-0.072*	2
	1978	(-6.08)	(-3.76)	(-3.07)	(5.51)	(3.71)	(-5.75)	
Maize	1919	-0.920	0.035*	-0.288*	-0.041*	0.114	-0.029*	4
	1972	(-5.28)	(2.90)	(-2.69)	(-3.21)	(1.24)	(-4.03)	
Copper	1917	-0.412	0.008	-0.321*	-0.008	0.245*	-0.005**	0
	1952	(-5.41)	(1.15)	(-3.58)	(-1.06)	(3.07)	(-1.88)	
Wheat	1913	-0.767	0.008	0.211*	-0.026	0.327*	0.006*	4
	1945	(-5.89)	(0.45)	(2.06)	(-1.37)	(4.20)	(1.95)	
Tin	1918	-0.457	0.018*	-0.301*	-0.011	0.269*	-0.034*	0
	1975	(-5.79)	(2.34)	(-3.27)	(-1.49)	(2.83)	(-4.70)	
Coffee	1948	-0.395	0.000	0.290*	-0.003	-0.611*	0.039*	0
	1986	(-5.29)	(0.12)	(2.61)	(-0.65)	(-3.72)	(2.01)	
Beef	1948	-0.380	0.005*	-0.486*	0.047*	0.398*	-0.069*	0
	1958	(-5.25)	(2.21)	(-3.11)	(2.17)	(2.47)	(-2.71)	
Palmoil	1918	-0.557	0.039*	-0.452*	-0.041*	-0.688*	0.410*	2
	1985	(-6.02)	(4.04)	(-4.82)	(-4.15)	(-5.59)	(3.17)	

Commodity	*TB1* *TB2*	ρ	β	μ_1	μ_2	μ_3	μ_4	*k*
2 Breaks in Intercept and 1 Break in Trend								
Cocoa	1946	-0.467**	-0.014*	0.708*		0.538*	-0.016*	2
	1972	(-5.77)	(-4.21)	(4.82)		(4.00)	(-2.44)	
Jute	1929	-0.634**	0.01*	-0.384*		0.363*	-0.023*	2
	1946	(-5.74)	(2.21)	(-3.16)		(3.74)	(-4.05)	
Lamb	1946	-0.476**	0.017*	-0.602*		0.337*	-0.011*	4
	1958	(-5.83)	(4.45)	(-4.40)		(3.18)	(-2.70)	
2 Breaks in Intercept Only								
Aluminum	1916	-0.362**	0.002	-0.227*		-0.306*		2
	1939	(-5.77)	(1.57)	(-4.20)		(-4.51)		
Rice	1929	-0.495**	0.000	-0.169*		-0.368*		2
	1981	(-5.84)	(0.09)	(-2.71)		(-4.75)		
Tobacco	1916	-0.464**	0.003*	0.273*		-0.174*		4
	1989	(-5.79)	(3.42)	(5.41)		(-3.90)		
1 Break in Intercept Only								
Banana	1924	-0.321*	-0.003*	0.195*				0
		(-4.99)	(-4.48)	(4.14)				

Notes: TB1 and TB2 correspond to the first and second break dates. The coefficients in the first row correspond to those coefficients in equation (3) in the text.

The bootstrapped critical values for 2 breaks in intercept and 2 breaks in trend are: -7.00 (1%), -6.41 (5%) and –6.19 (10%).

The bootstrapped critical values for 2 breaks in intercept and 1 break in 2nd trend are: -6.56 (1%), -6.04 (5%) and –5.73 (10%).

The bootstrapped critical values for 1 breaks in intercept and 1 break in trend are: -5.98 (1%), -5.31 (5%) and –4.90 (10%).

The bootstrapped critical values for 2 breaks in intercept only are: -6.51 (1%), -6.07 (5%) and –5.69 (10%).

The bootstrapped critical values for 1 break in intercept are: -5.67 (1%), -5.08 (5%) and –4.82 (10%).

The numbers in parentheses are t-statistics.

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

Table 3: Estimated trend-stationary models with two breaks (intercept and trend)

	HIDE	LEAD	RUBBER	SUGAR	TEA	TIMBER	WOOL	ZINC
TB1	1919	1946	1924	1923	1921	1914	1916	1914
TB2	1951	1981	1932	1971	1952	1938	1949	1921
$\hat{\alpha}$	4.75 (60.0)	4.42 (243.0)	7.15 (27.6)	5.29 (29.2)	4.91 (57.8)	3.46 (26.5)	5.32 (48.5)	4.57 (34.7)
100 $\hat{\beta}$	3.82 (5.40)	-0.283 (-3.98)	-6.44 (-3.91)	0.446 (0.36)	-2.54 (-4.01)	1.61 (1.23)	0.743 (0.71)	1.11 (0.80)
D_{L1}	-0.797 (-7.80)	0.396 (11.0)	1.46 (7.90)	-0.559 (-2.80)	0.427 (4.29)	0.405 (3.47)	0.307 (2.60)	1.07 (6.77)
100 D_{T2}	-3.44 (-4.93)	-0.745 (-5.96)	-24.8 (-4.76)	-0.899 (-0.67)	2.60 (3.42)	-2.74 (-1.68)	-1.87 (-1.64)	-26.43 (-6.82)
D_{L2}	-0.524 (-7.98)	-0.320 (-4.01)	0.791 (4.73)	0.670 (3.37)	0.460 (5.37)	0.283 (2.78)	0.260 (2.74)	0.444 (3.43)
100 D_{T2}	-0.234 (-0.89)	-0.059 (-0.09)	29.3 (6.83)	-4.28 (-3.67)	-3.21 (-7.17)	1.91 (2.53)	-2.44 (-5.25)	25.60 (7.73)
$\hat{\phi}_1$	0.756 (5.85)	1.36 (19.3)	-0.014 (-0.14)			0.622 (6.93)		0.540 (6.00)
$\hat{\phi}_2$		-0.608 (-8.44)	0.461 (4.81)					
$\hat{\theta}_1$	0.506 (2.35)	1.00 (37.9)	-1.00 (-19.7)	-0.594 (-8.53)	-0.540 (-6.64)		-0.540 (-5.82)	
$\hat{\theta}_2$	0.494 (2.54)							
$\hat{\theta}_3$								

Note: Figures in brackets are estimated t-statistics.

Table 4: Estimated trend-stationary models with two breaks in intercept and one break in trend

	COCOA	JUTE	LAMB
TB1	1946	1929	1946
TB2	1972	1946	1958
$\hat{\alpha}$	4.03 (27.7)	4.60 (47.8)	2.75 (14.3)
100 $\hat{\beta}$	-2.41 (-5.30)	1.77 (3.51)	2.24 (3.64)
D_{L1}	1.22 (6.54)	-0.585 (-4.20)	-0.598 (-2.94)
D_{L2}	0.798 (4.14)	0.476 (4.39)	0.554 (2.75)
100 D_{T2}	-1.92 (-1.42)	-3.75 (-6.52)	-1.13 (-1.01)
$\hat{\phi}_1$	0.644 (7.02)		0.723 (8.12)
$\hat{\phi}_2$			
$\hat{\theta}_1$		0.623 (6.67)	
$\hat{\theta}_2$			
$\hat{\theta}_3$			

Table 5: Estimated trend-stationary models with two breaks (intercept only)

	ALUMINUM	RICE	TOBACCO
TB1	1916	1929	1916
TB2	1939	1981	1989
$\hat{\alpha}$	5.88 (45.4)	5.11 (32.2)	3.60 (112.0)
100 $\hat{\beta}$	-0.738 (-2.53)	-0.091 (-0.95)	0.693 (9.93)
D_{L1}	-0.233 (-1.93)	-0.231 (-4.80)	0.546 (11.7)
D_{L2}	-0.325 (-2.60)	-0.712 (-13.1)	-0.373 (-6.91)
$\hat{\phi}_1$	0.679 (6.85)	1.37 (14.3)	0.751 (7.38)
$\hat{\phi}_2$		-0.559 (-5.89)	-0.167 (-1.36)
$\hat{\phi}_3$			-0.287 (-2.87)
$\hat{\theta}_1$	-0.468 (-4.27)	0.597 (5.59)	
$\hat{\theta}_2$		0.403 (3.85)	
$\hat{\theta}_3$			

Table 6: Estimated trend-stationary models with one break (intercept only)

	BANANA
TB	1924
$\hat{\alpha}$	4.82 (41.2)
100 $\hat{\beta}$	-0.330 (-1.52)
D_{L1}	0.199 (2.29)
$\hat{\phi}_1$	0.869 (13.5)
$\hat{\theta}_1$	
$\hat{\theta}_2$	
$\hat{\theta}_3$	

Table 7: Estimated *I(1)* models with two breaks (trend and level)

	COFFEE	COTTON	BEEF	SILVER	MAIZE	COPPER	WHEAT	PALMOIL	TIN
TB1	1948	1929	1948	1939	1919	1917	1913	1918	1918
TB2	1986	1948	1958	1978	1972	1952	1945	1985	1975
$100\ \hat{\beta}$	0.418 (0.73)	1.05 (2.48)	1.03 (1.65)	-1.68 (-4.12)	3.02 (3.04)	-0.105 (-0.08)	1.55 (1.03)	4.84 (4.53)	2.30 (1.99)
$\hat{\phi}_1$	0.666 (8.16)	0.670 (6.39)	0.737 (9.01)	0.547 (5.89)		0.640 (6.74)	0.844 (8.97)		
$\hat{\phi}_2$		-0.304 (-3.15)					-0.351 (-3.75)		
$\hat{\phi}_3$									
$\hat{\theta}_1$	1.00 (37.6)	0.999 (35.2)	1.00 (37.0)	0.999 (37.3)	0.458 (4.93)	1.00 (34.7)	1.00 (36.3)	0.277 (2.85)	0.278 (2.38)
$\hat{\theta}_2$					0.542 (6.05)			0.600 (7.25)	0.350 (3.63)
$\hat{\theta}_3$									0.373 (3.16)
$\hat{\theta}_4$									
ΔD_{L1}	0.456 (2.31)	-0.550 (-5.01)	-0.365 (-1.70)	-0.355 (-2.97)	-0.265 (-2.07)	-0.420 (-3.00)	0.246 (1.97)	-0.661 (-5.49)	-0.454 (-3.44)
$100\Delta D_{T1}$	-0.389 (-0.36)	1.20 (1.28)	3.22 (0.83)	4.60 (7.47)	-3.70 (-3.59)	-0.20 (-0.16)	-4.00 (-2.41)	-5.35 (-4.76)	-0.9 (-0.74)
ΔD_{L2}	-0.916 (-3.84)	0.200 (2.03)	0.951 (4.78)	0.667 (4.86)	0.132 (1.07)	0.452 (3.57)	0.469 (4.67)	-0.921 (-6.59)	0.433 (3.28)
$100\Delta D_{T2}$	0.963 (0.26)	-5.00 (-6.02)	-5.15 (-1.29)	-10.4 (-8.74)	-3.10 (-4.20)	-0.50 (-0.65)	0.80 (1.53)	4.63 (2.62)	-5.90 (-6.21)

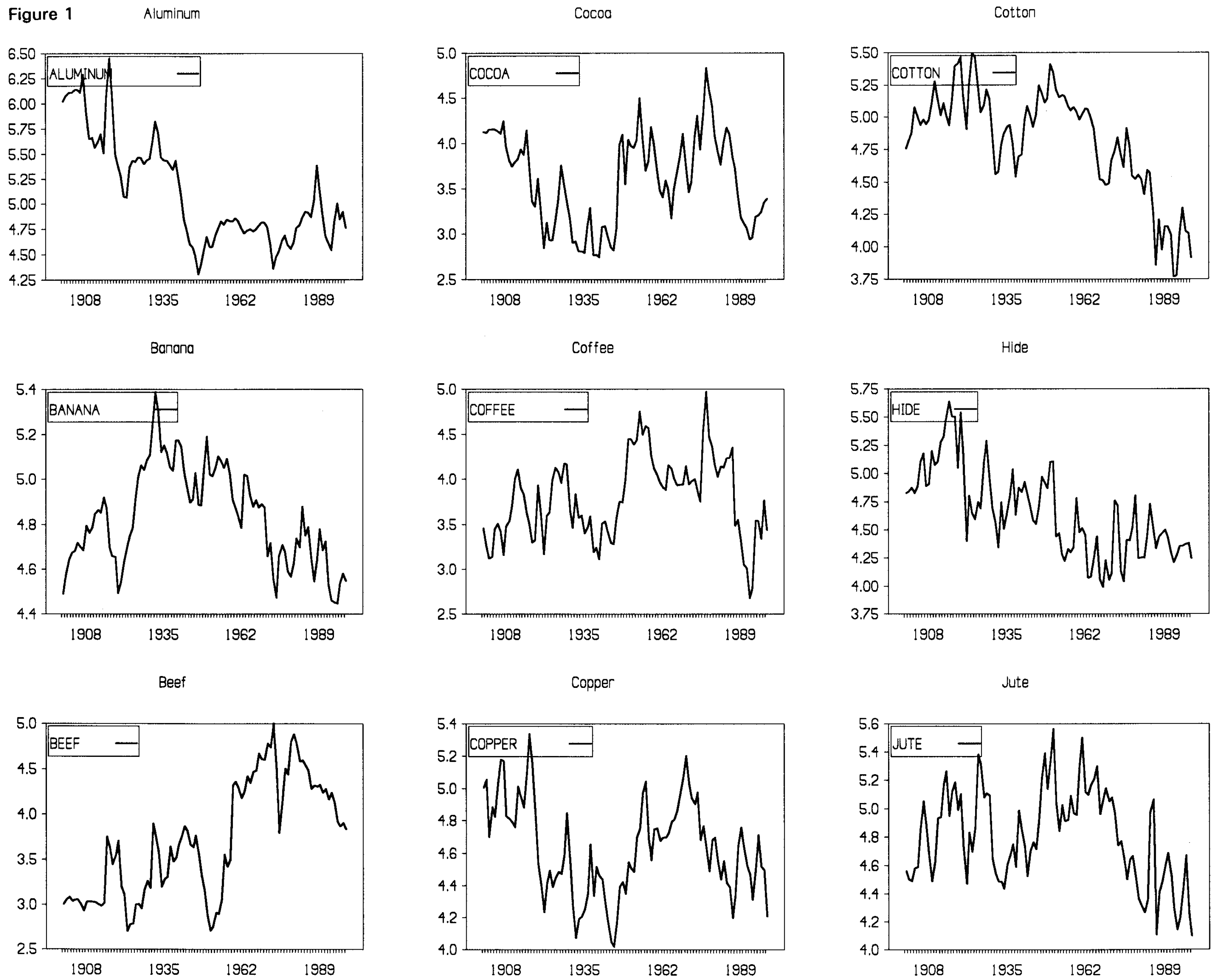
Table 8: Estimated trend-stationary models with two breaks (intercept and trend)

	COFFEE	COTTON	BEEF	SILVER	MAIZE	COPPER	WHEAT	TIN
TB1	1948	1929	1948	1939	1919	1917	1913	1918
TB2	1986	1948	1958	1978	1972	1952	1945	1975
$\hat{\alpha}$	3.50 (21.2)	4.93 (66.6)	2.99 (17.3)	4.18 (43.9)	4.89 (51.1)	4.94 (32.9)	5.00 (40.7)	3.41 (24.7)
100 $\hat{\beta}$	0.398 (0.73)	1.06 (2.58)	1.05 (1.80)	-1.67 (-4.26)	2.77 (3.28)	-0.04 (-0.03)	2.19 (1.63)	2.32 (2.04)
D_{L1}	0.471 (2.45)	-0.557 (-5.17)	-0.390 (-1.83)	-0.359 (-3.08)	-0.212 (-2.01)	-0.430 (-3.13)	0.135 (1.21)	-0.456 (-3.48)
100 D_{T1}	-0.399 (-0.39)	1.19 (1.32)	3.26 (0.87)	4.62 (7.78)	-3.43 (-4.11)	-0.333 (-0.22)	-4.32 (-2.94)	-0.922 (-0.77)
D_{L1}	-0.920 (-3.9)	0.203 (2.09)	0.967 (4.91)	0.670 (4.98)	0.053 (0.55)	0.466 (3.79)	0.397 (4.38)	0.437 (3.34)
100 D_{T2}	1.06 (0.29)	-5.01 (-6.20)	-5.20 (-1.36)	-10.4 (-9.09)	-2.96 (-5.65)	-0.472 (-0.68)	0.578 (1.21)	-5.93 (-6.31)
$\hat{\phi}_1$	0.648 (8.18)	0.660 (6.39)	0.714 (8.94)	0.530 (5.84)	0.849 (14.9)	0.617 (6.76)		
$\hat{\phi}_2$		-0.311 (-3.25)						
$\hat{\theta}_1$					-0.423 (-5.15)		0.765 (12.5)	0.712 (6.32)
$\hat{\theta}_2$					-0.577 (-7.21)			0.362 (3.09)
$\hat{\theta}_3$								

Table 9: Relative trend measure ψ

	ALUMINUM	BANANA	BEEF	COCOA	COFFEE	COPPER	COTTON	HIDES
ψ	1.000	0.000	0.000	1.000	0.000	0.000	0.505	0.000
	JUTE	LAMB	LEAD	MAIZE	PALMOIL	RICE	RUBBER	SILVER
ψ	0.535	0.000	1.000	0.800	0.677	0.000	1.000	0.800
	SUGAR	TEA	TIMBER	TIN	TOBACCO	WHEAT	WOOL	ZINC
ψ	0.273	0.687	0.242	0.242	0.000	0.869	0.838	0.071

Figure 1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Banana
BANANA
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5.0
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4.4
Coffee
COFFEE
Hide
HIDE
5.75
5.50
5.25
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4.00
3.75
Beef
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Figure 1
(Con't)

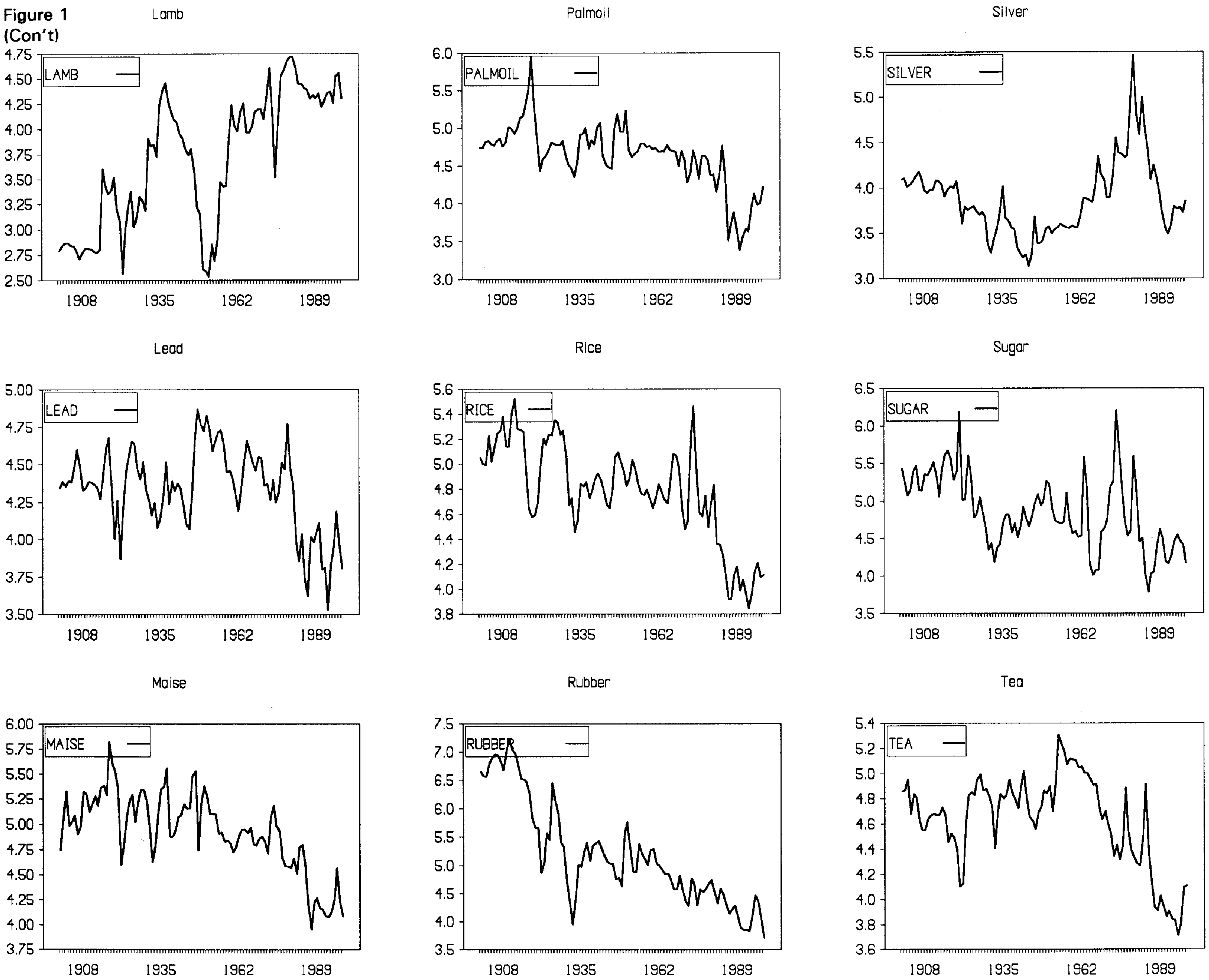

Figure 1
(Con't)