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The Long-Run Behavior of Commodity Prices: Small Trends and Big Variability

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Abstract

Using the longest dataset publicly available (*The Economist*'s index of industrial commodity prices), we analyze the behavior of real commodity prices over the period 1862-99, and have two main findings. First, while there has been a downward trend in real commodity prices of 1.3 percent per year over the last 140 years, little support is found for a break in the long-run trend decline in commodity prices. Second, there is evidence of a ratcheting up in the variability of price movements. The amplitude of price movements increased in the early 1900s, while the frequency of large price movements increased after the collapse of the Bretton Woods regime of fixed exchange rates in the early 1970s. While there is a downward trend in real commodity prices, this is of little practical policy relevance as it is small and completely dominated by the variability of prices.

JEL Classification Numbers: E32, Q11

Keywords: Commodity prices, trends, cycles, variability

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[1] John McDermott is an advisor in the Reserve Bank of New Zealand; this paper was completed while he was a Visiting Scholar in the Research Department. The authors would like to thank Ximena Cheetham, Hong Liang, Sam Ouliaris, Blair Rourke and Peter Wickham for helpful discussion and comments. The views expressed in the paper are those of the authors, and do not necessarily represent the views of either the IMF or the Reserve Bank of New Zealand.

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What commodity prices lack in trend, they make up for in variance (Deaton (1999), p.27).

I. INTRODUCTION

This paper is concerned with the empirical behavior of commodity prices, in particular, changes in the variability of commodity prices and in the trend growth of prices over time. This is an important question, as about 25 percent of world merchandise trade consists of primary commodities, and many developing countries are dependent on one or a few commodities for the majority of their export earnings (Cashin and Pattillo (2000)).[2] Both sharp fluctuations and long-run trend movements in commodity prices present serious challenges for many developing countries, due to their large impacts on real output, the balance of payments and government budgetary positions, and because of the consequent difficult problems they pose for the conduct of macroeconomic policy.

It has been postulated that owing to the low income elasticity of demand for commodities and because productivity increases have been larger for manufactured goods than for primary commodities, the price of commodities relative to manufactured goods should decrease over time—the Prebisch-Singer hypothesis. If this hypothesis were true, then the long-term outlook for commodity-exporting countries would be quite unfavorable. Few ideas in development economics have been studied more intensively, yet remain so controversial. Prebisch (1950) and Singer (1950) originally found a downward trend in real commodity prices. While later work by Grilli and Yang (1988) found that this downward trend accelerated in 1921, there are several papers which fail to detect such an adverse trend movement (see Cuddington and Urzúa (1989)). Accordingly, much of the empirical debate about the Prebisch-Singer hypothesis has centered on whether the nonstationarity of real commodity prices takes the form of a deterministic trend, a stochastic trend, or whether there are structural breaks in the trend. While many researchers have imposed a break around 1920 based on visual inspection of the data, León and Soto (1997) have criticized such ad hoc procedures for identifying breaks.[3]

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[2] According to the World Bank's *World Development Indicators*, primary commodities accounted for 42 percent of developing (low- and middle-income) countries' total merchandise exports in 1997, compared with 19 percent for developed (high-income) countries. Commodity dependence is even greater in Sub-Saharan Africa, where primary commodities account for about 75 percent of total exports in 1997.

[3] Cuddington and Urzúa (1989) and Perron (1990) found support for a structural break in 1920/21. Powell (1991) prefers the hypothesis of three downward jumps in real commodity prices (in 1921, 1938 and 1975), rather than a long-run trend decline. See Lutz (1999) for an overview of empirical tests of the Prebisch-Singer hypothesis.

The controversy over the Prebisch-Singer hypothesis probably remains unresolved because there is a low signal to noise ratio in the commodity price data. Over the last 140 years, real commodity prices have declined by about 1 percent per year, but this has not been a smooth process, as prices can change by as much as 50 percent in a single year. Clearly, price variability is large relative to trend. This makes forecasting future commodity prices a difficult exercise, as they are subject to large and unpredictable movements that may have persistent effects. As volatility is a key feature of commodity prices, information on the nature of this volatility can be useful to policymakers.

In addition to examining long-run price trends, this paper is concerned with the length and size of commodity-price cycles. The topic is important, as improving our understanding of the duration and amplitude of commodity-price cycles will be a key input to efforts to stabilize the macroeconomic effects of movements in prices, particularly for commodity-dependent countries. Information on the average duration and amplitude of booms and slumps can be used in designing domestic countercyclical policies, in examining whether it is useful to borrow externally in the presence of a temporary adverse shock, or in deciding on the efficacy of national commodity stabilization funds and international market-sharing agreements (Deaton (1999), Deaton and Miller (1996)). In addition, the paper will examine whether the key features of commodity-price cycles have changed over time. If so, what is driving the movements in price volatility—is it a change in the frequency of price shocks, or a change in the amplitude of price shocks?

For policy purposes, a good understanding of the cyclical behavior of commodity prices is as important as an understanding of their underlying long-run trends. While it is generally agreed that the volatility of primary commodity prices is a significant policy issue for commodity-exporting countries, there has been little previous work examining this issue. A key exception is Cuddington and Liang (1999). These authors study conditional variances and how these shift over time. However, variances lump together both the amplitude and duration of price movements, and it might also be useful to separate out these effects to examine the duration of booms and slumps in commodity prices.

Cuddington (1992), Reinhart and Wickham (1994), and Cashin, Liang and McDermott (2000) find that shocks, which drive cycles in commodity prices, can exhibit differing degrees of persistence across commodities. However, to our knowledge there has been little previous work that has studied the duration of commodity-price cycles. Cuddington (1992) and Reinhart and Wickham (1994) both study the duration and persistence of commodity-price cycles using the Beveridge-Nelson decomposition technique, which gives a different meaning to the cycle than the one we wish to adopt. The cycle derived from use of the Beveridge-Nelson decomposition is defined in terms of deviation from trend. However, the definition of the cycle we use here follows Watson (1994) and Cashin, McDermott and Scott (1999) by dealing with the data in levels, hence avoiding the somewhat subjective choice as to which detrending method to use. Price slumps (booms) are then described as periods of absolute declines (increases) in the series, not as a period of below-trend (above-trend) growth in the series.

The paper is organized as follows. Section II sets out several nonparametric tests to detect a change in the long run behavior of commodity prices, particularly changes in the trend growth of prices, and changes in the duration and volatility of commodity-price cycles. Section III gives an overview of the long-run data used, while Section IV presents the empirical findings. Section V contains some concluding comments.

II. Tests for Change in Trend Growth, Volatility and Cycle Duration

The issues raised above can be resolved by comparing commodity prices over two different time periods within our sample. We will examine the differences from the early part of the sample in comparison with the later part of the sample over two dimensions. First, we will analyze whether there have been changes in trend growth rates of prices, and second, whether there have been changes in the volatility of price movements. When analyzing the volatility of price movements we will examine changes in variances (which include both changes in the amplitude and duration of cycles in prices) and changes in the duration of cycles alone. Our focus on the duration perspective considers the relative *frequency* of commodity-price cycles, while the volatility perspective also considers changes in the amplitude of cycles. Specifically, we will test the null hypotheses of no change in the trend growth of commodity prices, no change in the duration of commodity price cycles, and no change in the variance of commodity prices.

A. Null Hypotheses To Be Tested

To form these tests, we need to split our time series into two subperiods. That is, consider a sample period $\{X_1,\dots, X_N\}$ split into two subperiods, $\{X_1,\dots, X_{N_1}\}$ and $\{X_{N_1+1},\dots,X_N\}$. Thus the number of elements in the first subperiod is N_1 and the number of elements in the second subperiod is $N_2 = N\text{-}N_1$.

First, we examine whether there is any change in the trend rate of growth of commodity prices. Following Watson (1994), the t statistic for testing the null hypothesis of no change in the trend rate of growth of prices is

$$t_{\bar{X}} = \frac{\bar{X}_1 - \bar{X}_2}{(\sigma_1^2 + \sigma_2^2)^{1/2}}\ , \tag{1}$$

where the average annual growth rate in the i^{th} subperiod (i=1,2) is $\bar{X}_i = (1/N_i)\sum_i \Delta \ln X_t$, and we define $\sum_1 = \sum_{t=1}^{N_i}$; $\sum_2 = \sum_{t=1+N_i}^{N}$. We also define the autocorrelation-robust variance error estimate as

$$\sigma_i^2 = (1/N_i)\, s_i^2 \left[1 + 2\,\hat{\rho}_i/((1-\hat{\rho}_i)(1-\hat{\rho}_i^{\,2}))\right], \tag{2}$$

where the variance estimators of the pre-whitened residuals ($\varepsilon_t = u_t - \hat{\rho}_1\ u_{t-1}$ for t=2,…,N_1 and $\varepsilon_t = u_t - \hat{\rho}_2\ u_{t-1}$ for t=N_1+2,…,N) are defined as $s_i^2 = (1/N_i - 2)\sum_i \varepsilon_t^2$; the autocorrelation coefficient is defined as $\hat{\rho}_i = \sum_i u_{t-1} u_t$; and the raw residuals are $u_t = X_t - \overline{X}_1$ for t=1,…,N_1 and $u_t = X_t - \overline{X}_2$ for t=N_1+1,…,N. When the breakpoint is known, the t statistic for testing the null of no change in growth rates follows a t-distribution.

The second feature of the data we want to examine is whether the variance of price changes in two subperiods are different. To do so, we consider the ratio of the sample variances. To test the null hypothesis of no change in the variance of price changes, the test statistic we use is defined as

$$F = \frac{s_1^2}{s_2^2}\ , \tag{3}$$

where the variance of price changes in each subperiod is s_i^2 (for i=1,2), as defined for equation (2) above. When the breakpoint is known and the two samples are independent, this test has an F-distribution with degrees of freedom N_1 and (N- N_1).

Finally, we examine whether the duration of booms and slumps in commodity prices has changed. Identifying specific cycles in economic time series requires precise definitions of a boom and a slump. For annual data we can simply use the change in price over a given year as an indicator of when booms and slumps occur. For future reference we introduce the following definitions:

Definition 1: For annual data *booms* are defined as sequences of absolute increases in prices, and *slumps* are defined as sequences of absolute decreases.

Definition 2: A *cycle* includes one boom and one slump.

To test the null hypothesis of no change in the duration of booms and slumps in commodity prices we adapt a nonparametric test suggested by Diebold and Rudebusch (1992), which has as its maintained assumption that the distribution of durations (of booms or slumps) differ only by a location shift. Their test is a version of the Wilcoxon rank sum statistic, which is defined as

$$W^* = (n_1\, n_2 / n)^{1/2} \left\lfloor (\overline{R}_2 - \overline{R}_1) / s^* \right\rfloor , \tag{4}$$

where n_i is the number of cycles in the i^{th} subperiod (i=1,2); $n=n_1+n_2$; $\overline{R}_1$ and $\overline{R}_2$ denote the mean ranks of the first and second subperiods[4], and the standard error is defined as

$$s^* = (n-2)^{-1/2}\left[\sum_{i=1}^{n_1}(R_i - \overline{R}_1)^2 + \sum_{i=n_1+1}^{n}(R_i - \overline{R}_2)^2\right]^{1/2}. \tag{5}$$

When the breakpoint is known, the statistic for testing the null of no change in the duration of booms and slumps can be used like a t test which converges in distribution to a standard normal.

B. Hypothesis Testing with Structural Change at an Unknown Break

The above tests are conditional on structural change at a known point in time. However, many researchers select the breakpoint after first examining plots of the data. This procedure implies that the breakpoint is a function of the data, rather than being a truly exogenous process. Importantly, this endogenization of the breakpoint selection procedure (however informal) will alter the distributions of the above test statistics. This means that by searching for the breakpoint, there is an increase in the probability of rejecting the null hypothesis of no change. For example, suppose one tests for the date of a break in a time series at the start of the First World War (with our test having a nominal size of 5 percent), and then we repeat the test assuming the break is at the start of the Second World War (with that test also having a nominal size of 5 percent). While individually both tests each have a correct size of 5 percent, the probability that at least one of these points has a break must be higher than 5 percent if we use the same critical values as the individual tests. As a result, if we allow all points in time to be a potential break, then the probability of a break at some point in time is much greater than 5 percent (given standard critical values).

Clearly, for all three tests we cannot assume that all breakpoints are known. If one takes the view that the breakpoint is unknown, then our tests for change in trend, duration and volatility will need to account for the fact that the breakpoints are data-dependent. Our testing procedure is to search over all possible breakpoints in the sample and select the largest of these tests. That is, we are interested in test statistics of the form

$$\sup_{\pi\in\Pi} Q(\pi), \tag{6}$$

[4] Following Diebold and Rudebusch (1992), in the case of a tie the relevant ranks are replaced by the average of the ranks of tied observations.

where $Q(\pi)$ is the test statistic of interest, conditioned on the value of the breakpoint $\pi=N_1/N$ (for change in trend and volatility) and $\pi=n_1/n$ (for change in cycle duration); and where Π is some prespecified subset of [0,1].[5] Suppose that for a given π the test statistic $Q(\pi)$ has a distribution which can be represented by the probability density function $f(x)$, and which has a cumulative density function $F(x)$. Then the cumulative distribution of sup $Q(\pi)$, if they are independently drawn, is

$$\Pr\left(\sup(x) < m\right) = F(m)^n, \tag{7}$$

where m is such that $\Pr(\sup(x) < m) = 1-\alpha$ and α is the size of our test. Thus the probability density function is given by the derivative

$$pdf(m) = nF(m)^{n-1} f(m). \tag{8}$$

For example, if $Q(\pi)$ follows a normal distribution, as in the tests for changing duration, then sup $Q(\pi)$ will follow the distribution given by the probability density function

$$pdf(m) = n\left(\int_{-\infty}^{m}\frac{1}{\sqrt{2\pi}}\exp\left(-x^2/2\right)dx\right)^{n-1}\frac{1}{\sqrt{2\pi}}\exp\left(-m^2/2\right). \tag{9}$$

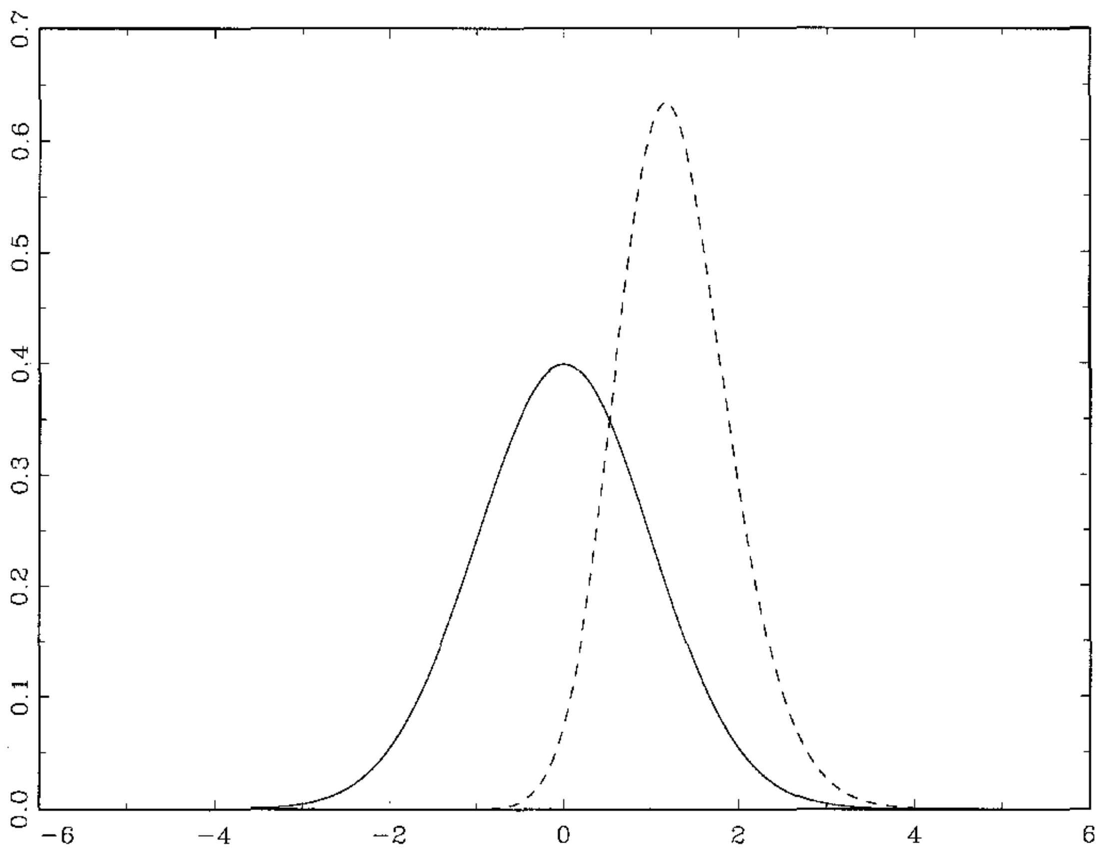
The distribution shown in equation (9), for the case when $n=7$, is compared to the standard normal distribution in Figure 1. It can clearly be seen that the effect of searching over all possible break points is to shift the distribution of the Wilcoxon rank sum statistic to the right. Without making this correction, searching over break points would increase the probability of falsely finding a break. The correction counters this effect, so that the probability of falsely accepting a break is maintained at the correct size (in our case, 5 percent). Similarly, we can derive the probability density function of the sup $Q(\pi)$ statistic if $Q(\pi)$ follows a t or F distribution, as in the test for a change in trend growth rates or change in variance, respectively.

We also use these statistics to estimate the timing of any change-point. Intuitively, the location of the maximum $Q(\pi)$ statistic will be a reasonable estimator for the location of the change-point. That is, our estimator of the change-point is

$$\hat{\pi} = \inf\left\{\pi \in \Pi : \sup_{x\in\Pi} Q(x)\right\}. \tag{10}$$

[5] The sup type of statistic has been analyzed in detail by Andrews (1993), where they are shown to have certain optimality properties.

Figure 1. Distribution of the Wilcoxon Rank Sum Test
for a Change in Duration of Cycles



Notes: The distribution of the Wilcoxon rank sum test (when there is a change in cycle durations of known timing) is shown by the solid line and is $N(0,1)$. The distribution of the sup W test (when there is a change in cycle durations of unknown timing) is shown by the dashed line. The distribution for a change in cycle durations with unknown timing is drawn for a possible seven breaks.

Defining $\hat{\pi}$ as a maximizer over Π instead of over [0,1] serves the sole purpose of excluding change-points located arbitrarily close to the ends of the sample.[6]

III. DATA

In analyzing commodity-price movements, we use the longest data set publicly available—the industrial commodities index of *The Economist.* The real annual data consists of the nominal industrial commodities price index (dollar-based with base 1845-50=100, weighted by the value of developed-country imports), deflated by the GDP deflator of the United States, over the period 1862-1999. While data are available on the nominal index back to 1851, data prior to 1857 are incomplete and data between 1857-61 are available only for January of each year. Between 1862-1910, the annual data are formed as an average of price observations in January and July of each year; from 1911 onwards, the annual data are an average of all months. The industrial commodities price index consists of the prices of metals and nonfood industrial commodities. Figure 2 sets out both the nominal and real industrial commodities indices—with the latter showing the potential command over American resources of a fixed quantity of industrial commodities.[7]

The index of real industrial prices, after being somewhat stable over its first four decades, fell by four-fifths between 1900-99, and ended the century at its lowest level on record (Figure 2). The purchasing power over American resources of a given basket of industrial commodities in 1999 is one-fifth the purchasing power yielded by that same basket of commodities in 1862. The largest annual declines in commodity prices were 37 percent from 1930 to 1931 and 1974 to 1975, followed by the 34 percent fall between 1951 and 1952. Conversely, the largest annual rise in commodity prices was 49 percent from 1972 to 1973, followed by the 33 percent rise between 1916 and 1917.

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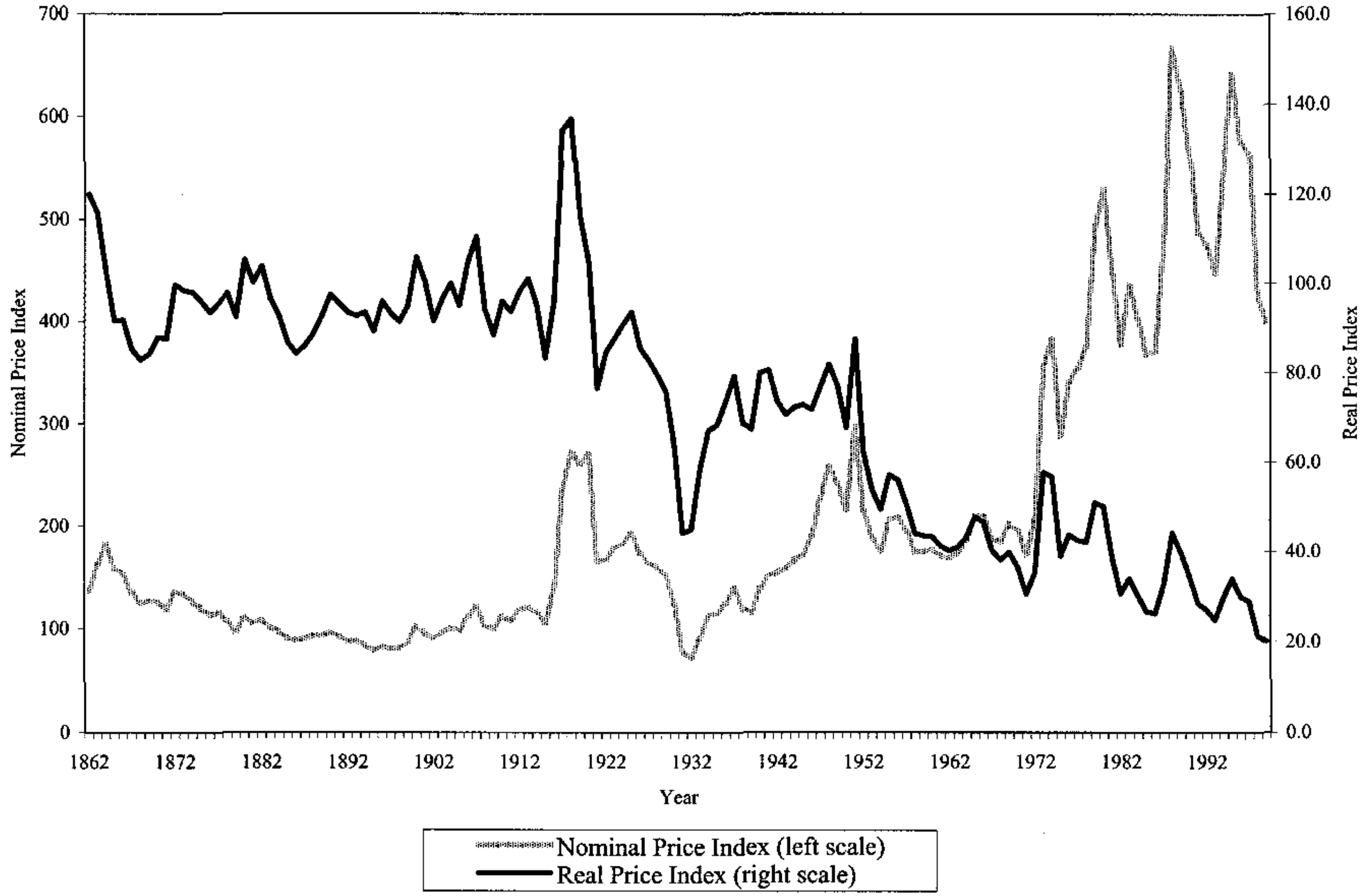
[6] Let $\Pi \subset [0,1]$ be the set of all possible breakpoints searched over, where in this case $\Pi=[0.1, 0.9]$. For our data series (1862-1999) this will means that the first 15 and last 15 observations will be excluded from being possible breakpoints. Testing whether a sample has a change point at its extremities is not desirable, because statistics tend to diverge to infinity since they cannot discriminate between true change points and boundary conditions.

[7] The 2000 version of *The Economist's* industrial commodities price index (base 1995=100, weighted by the value of world imports in 1994-96) has a weight of 45.9 percent for nonfood agricultural commodities and 54.1 percent for metals. The *metals* index consists of aluminum, copper, nickel, zinc, tin, and lead; the *nonfood agricultural* index consists of: cotton, timber, hides, rubber, wool 64s, wool 48s, palm oil, coconut oil, soybeans, and soybean oil. The commodity coverage of this index has changed over time—for details see *The Economist* ("A Raw Deal for Commodities", April 17, 1999 and "A Changed Commodity", January 15, 2000).

Figure 2. Nominal and Real Price of Industrial Commodities, 1862-1999
(1845-50=100)



Source: *The Economist.*

Unlike most previous studies of long-run commodity prices, we do not use the Grilli-Yang (1988) index. The Grilli-Yang index runs from 1900-88, while *The Economist* industrials index covers four more decades of data, and gives the best chance of understanding the long-run trends in the data (and thus allows for a more powerful test of the Prebisch-Singer hypothesis). In any event, the correlation of the Grilli-Yang index with *The Economist* industrials index is very high. Over the period used by Grilli-Yang (1900-88) the correlation between the two series is 0.81, and is even higher (0.85) when the Grilli-Yang data is extended until 1999.

Before proceeding further, it is useful to have in mind some of the salient features of this price index. Table 1 summarizes some of the facts about *The Economist* industrial commodities price series, for both its full history and for three subperiods. Following Eichengreen (1994), Reinhart and Wickham (1994) and Cuddington and Liang (1999), the subperiods have been selected to analyze the behavior of commodity prices across exchange rate regimes, as fixed- and flexible-exchange rate regimes have been found to yield differing volatility of real commodity prices.[8] The subperiods include the fixed exchange rate regime of the gold standard (1862-1913) and the Bretton Woods system (1946-71), and the flexible-exchange rate regime of the post-Bretton Woods period (1972-99).

Table 1. Industrial Commodity Price Facts, 1862-1999

Period	Autocorrelation Coefficient (One year)	Autocorrelation Coefficient (Two years)	Persistence (In years)	Coefficient of Variation	Skewness	Kurtosis	Volatility Persistence (In years)
1862-1999	0.93	0.85	3.29	0.38	-0.20	1.96	2.79
			[2.18, 6.12]				[1.86, 5.01]
1862-1913	0.57	0.19	1.26	0.08	0.96	4.59	0.85
			[0.79, 2.24]				[0.62, 1.19]
1946-71	0.81	0.67	0.90	0.30	0.83	2.52	0.62
			[0.38, 2.59]				[0.17, 1.57]
1972-99	0.67	0.32	0.75	0.28	0.64	2.70	0.72
			[0.34, 1.70]				[0.25, 1.99]

Notes: Autocorrelations of one and two years are the first- and second-order autocorrelation coefficients, respectively. Persistence is measured by the half-life of innovations, that is $\ln(½)/\ln(\alpha)$ where α is the autoregressive parameter from an AR(1) model for the levels (for column (4)) and an ARCH(1) model for the variances (for column (8)). For α 0, the half-life is the length of time (number of years) until the impulse response of a unit shock is half its original magnitude. The 90 percent confidence interval for the half-life is given in square brackets, for both columns (4) and (8). The coefficient of variation is the ratio of the standard deviation to the arithmetic mean. The skewness measure is $\mu_3/(\mu_2)^{1.5}$ and the kurtosis measure is $\mu_4/(\mu_2)^2$, where μ_r is the r^{th} (central) moment. The skewness of a symmetrical distribution, such as the normal, is zero; similarly, the kurtosis of the normal distribution is 3.

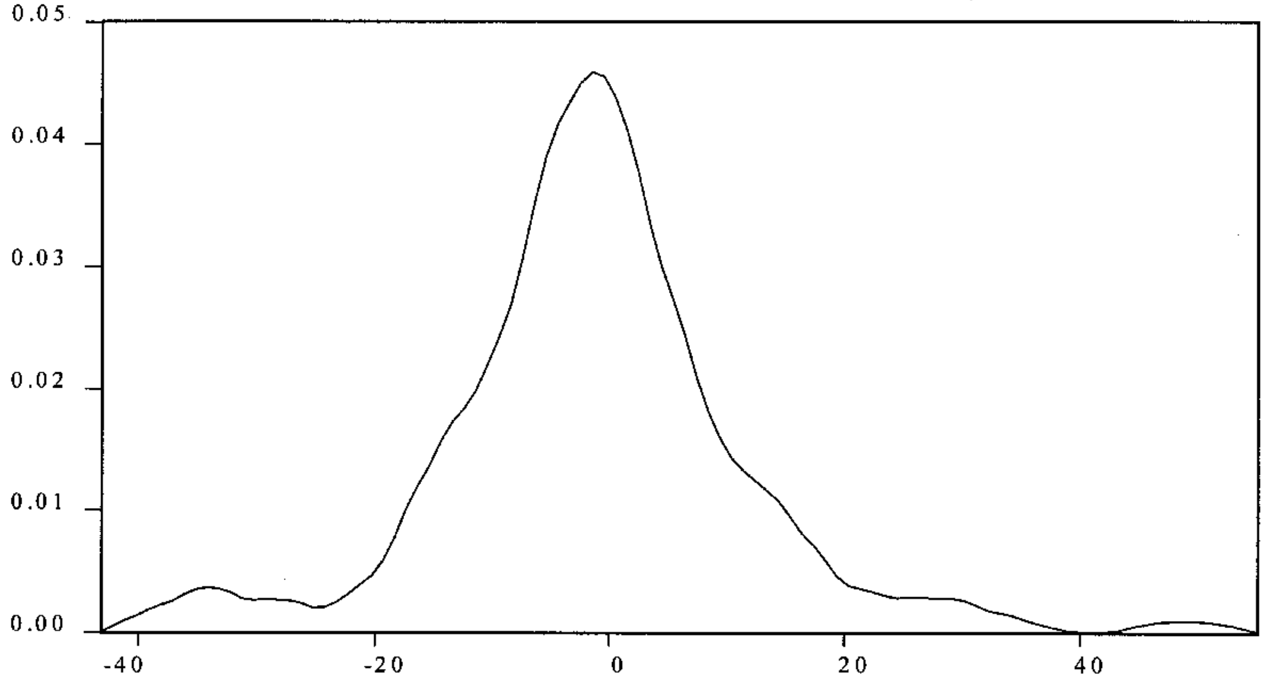
[8] The validity of this finding will be examined in Section IV below, when analyzing the presence or absence of a significant break in the variability of industrial commodity prices.

For the full sample, both the first- and second-order autocorrelation coefficients are quite high, indicating that prices tend to revert to their mean or to a deterministic trend at a rather slow rate. The persistence measure in the third column presents the half-life of shocks to the industrial commodities price index—the length of time it takes for half of any initial shock to dissipate (see Cashin, Liang and McDermott (2000)). The median half-life is found to be about 3 years, with 90 percent of the shocks to industrial commodity prices having a half-life ranging between 2.2 and just over 6 years. In addition, the median persistence of shocks in the flexible-exchange rate regime appears to be less than for the fixed-exchange rate regimes.

Consistent with earlier findings by Deaton and Laroque (1992), the coefficient of variation reveals that there is substantial annual variability in the prices of industrial commodities, particularly in the flexible-exchange rate regime. There is negative skewness in commodity prices for the full sample period, as there appear to be slightly more downward spikes in prices than upward spikes (see also Figure 2). During each of the subperiods there was substantial positive skewness, indicating that the upper tail of the distribution of prices is thicker than the lower tail. For the full sample and the latter two subperiods, commodity prices display kurtosis, with tails thicker than the normal distribution (platykurtic), indicating that large price movements are relatively common. The clear exception is the leptokurtic behavior of prices during the gold standard period. The final column of Table 1 examines the persistence of price volatility, finding that the volatility of prices tends to revert to its average relatively quickly. Over the various subperiods the persistence in volatility is relatively small, in that half the innovation in the variances has dissipated is less than a year. Accordingly, there is little evidence of clustering of large or small variances in commodity prices.

Figure 3 shows the empirical distribution of the (absolute) change in real commodity prices, which is centered a little less than zero, but has a high variance around its average trend decline (-1.3 percent per year) and fat tails (especially for price declines). Clearly, the large dispersion in price changes dominates the small secular decline in commodity prices.

Figure 3. Distribution of Percentage Changes in Real Commodity Prices



Note: The empirical distribution of price movements was estimated using the Epanechnikov kernel density with a bandwidth of 6.1.

IV. EMPIRICAL RESULTS

In this Section we present the results from our empirical analysis of the properties of the industrial commodities index. Specifically, we will test the null hypotheses of no change in the trend growth of commodity prices (Section IV.A), no change in the variance of commodity prices (Section IV.B), and no change in the duration of commodity price cycles (Sections IV.C and D).

A. Trends

The salient feature of the real price of industrial commodities is that it has a downward trend over the span of nearly 140 years, which seems to become marginally steeper around 1920 (Figure 2). The trend decline in the series is 1.3 percent per year. However, prices of industrial commodities showed no significant change in trend, as the t-statistic for our test of trend break ($t_{\bar{x}}$) over all possible breakpoints, was less than the 5 percent critical value (of the sup t distribution) of 3.30 in absolute value. Although not significant, the highest t statistic for changing trends (of 1.15) was located at 1917, which is close to the trend break imposed in earlier work by Grilli and Yang (1988) and Cuddington and Urzúa (1992). The average annual rate of decline in industrial commodity prices between 1917 and 1999 is 2.3 percent.

The trend lines over the 60-70 year periods before and after the First World War are a poor representation of the local (say decade-long) trends—the local trends vary dramatically from 2.7 percent in the 1910s to –6.9 percent in the 1990s. However, yearly price movements are often very large, in excess of 40 percent in some cases (see Figure 3). Between 1862-1999, more than 5 out of every 100 annual price movements were in excess of 20 percent. Accordingly, trends appear to be widely variable and largely uncertain, and cannot be relied upon as a basis for making forecasts of future commodity prices.

B. Volatility

Sharp price movements can be seen in Figure 4, which shows the rate of change of industrial commodity prices. Annual price changes of less than 20 percent are the norm prior to the abandonement of the gold standard in 1913. After that time, annual price changes of more than 20 percent occurred on 13 occasions.

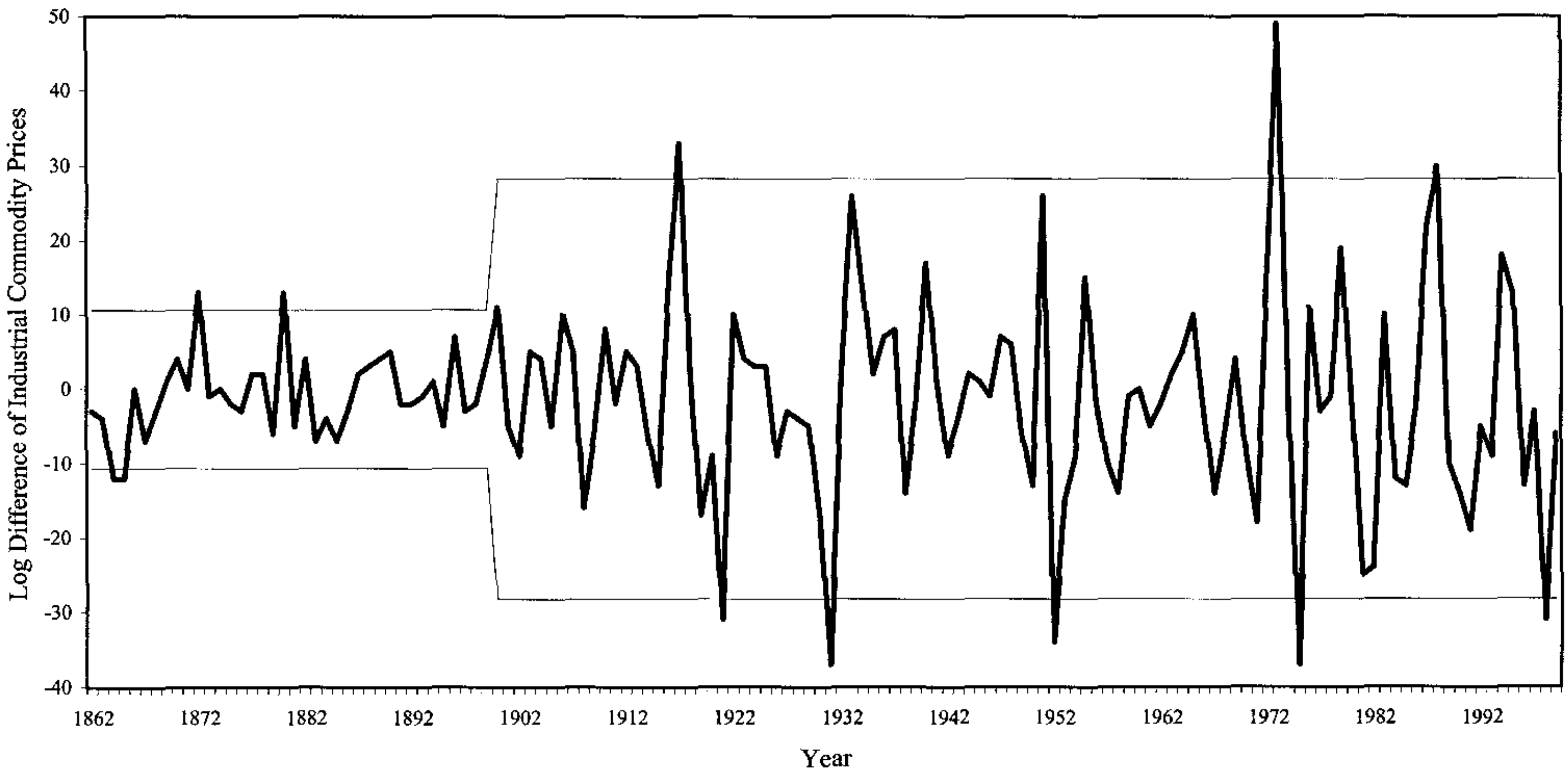
In examining changes in the volatility of the price of industrial commodities, the largest *F*-statistic for testing the null hypothesis of no change in the variance of price changes is 6.99, which is larger than the 5 percent critical value (of the sup F distribution) of 4.75. The largest *F*-statistic occurs in 1899. We have a run of very high *F*-statistics from the end of the 19th Century to the start of the First World War. In fact, the *F*-statistic as a function of break points has a local peak in 1914 (with a value of 6.23).

Figure 4 also shows the 95 percent confidence bands around the growth rate movements. For industrial commodities, the bands widen sharply after the change-point in 1899. Importantly, the frequency with which price movements exceed two standard deviations appears to increase after 1971. In the 70-year period from 1899 to 1971, movements in industrial commodity prices exceeded two standard deviations four times, while there were four price movements in excess of two standard deviations in less than 30 years from 1972 to 1999.

C. Cycles

Prior to applying the Wilcoxon rank sum statistic to examine whether there has been a change in the duration of booms and slumps in commodity prices, two features of the commodity-price data need to be verified. First, the duration of booms and slumps need to be independent (in order to derive appropriate critical values for the Wilcoxon test). Second, there should be no change in the trend rate of growth of prices. These pretests are necessary because commodity-price cycles are delineated in this paper on a nontrend-adjusted basis. As a result, any difference in trend growth rates in prices across subperiods would affect the comparability of durations of booms and slumps (Diebold and Rudebusch (1992)). The independence of

Figure 4. Percentage Change in Real Industrial Commodity Prices, 1862-1999



Note: The 95 percent standard error bands (derived as 1.96 times the mean standard deviation of the annual percentage change) are denoted by the thin line.

duration assumption appears to be confirmed, as the correlation between the length of successive booms or the length of successive slumps (over the entire sample) are only significantly different from zero at about the 15 percent level of significance.[9] As to the no difference in trend growth assumption, we have demonstrated above that the growth of prices exhibited no significant change in trend across all possible breakpoints. Accordingly, changes in trend growth are not influencing our findings on changes in the duration of commodity-price cycles.

Using our definition of booms and slumps, we find there are 18 completed cycles in industrial commodity prices over the full sample, where a cycle includes one boom and one slump. The dates of the peaks and troughs in industrial commodity prices are shown in Table 2, along with their duration, the rank number of that duration, and the sup W statistic testing the null hypothesis of no change in duration of booms and slumps in commodity prices.

Over these 18 cycles, we cannot reject the null hypothesis of no change in the duration of either booms or slumps, irrespective of the location of the breakpoint. That is, for all 18 possible breakpoints in the duration of each boom and slump, the largest t statistic for duration change in booms (2.1) and slumps (2.5) was less than the 5 percent critical values (of the sup W distribution) of 2.75 and 2.77, respectively. The shaded portions of Figure 5 denote the boom phase of the cycle, while the unshaded portions denote the slump phase of each cycle. Consistent with earlier work on commodity-price cycles (Cashin, McDermott and Scott (1999)), on average price slumps (4.2 years) last longer than price booms (3.6 years).

D. Large Cycles

Clearly, there has been an increase in the last 30 years in the frequency with which large price movements are occuring. To investigate further this feature of the price series, we examine whether there has been any change in the duration of large booms and large slumps. We formally define what we mean by large booms and large slumps below.

Definition 3: For annual data, *large booms* are defined as a sequence of generally increasing prices that have had a price movement of at least 25 percent over the phase, and *large slumps* are defined as sequences of generally decreasing prices that have had a price movement of at least 25 percent over the phase.

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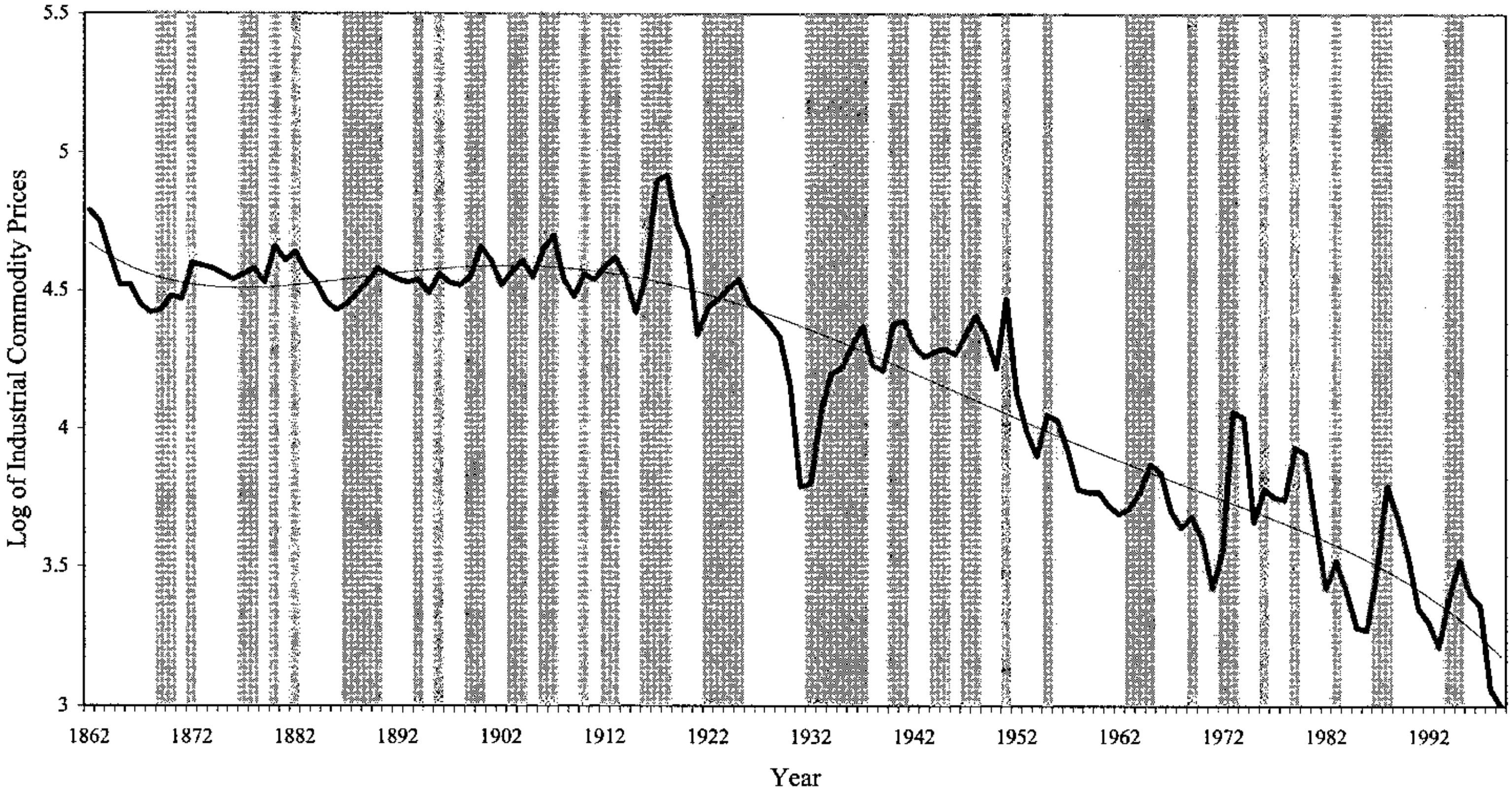
[9] The correlation between the length of successive booms (-0.36) is just significant at the 15 percent level, while the correlation between the length of successive slumps (0.24) is not significant at even the 20 percent level; the 15 percent and 20 critical values (0.35 and 0.31) were calculated as $1.44/T^{1/2}$ and $1.28/T^{1/2}$, where T=17 is the number of successive booms and slumps.

Table 2. Duration and Rank of Booms and Slumps in Industrial Commodity Prices

Cycle	Peak	Trough	Duration (In years)		Rank		sup W test	
			Boom	Slump	Boom	Slump	Boom	Slump
1	1862	1868		7		17.5		1.5
2	1872	1879	4	7	9.5	17.5	0.0	2.3
3	1880	1886	1	6	2	14	1.3	2.5
4	1890	1895	4	5	9.5	10	1.1	2.2
5	1896	1898	1	2	2	4	2.1	1.3
6	1900	1902	2	2	9.5	4	1.9	0.6
7	1907	1909	5	2	16	4	0.9	0.1
8	1913	1915	4	2	9.5	4	0.9	0.4
9	1918	1921	3	3	9.5	10	0.9	0.4
10	1925	1931	4	6	9.5	14	0.1	0.1
11	1937	1939	6	2	17	4	0.1	0.6
12	1941	1943	2	2	9.5	4	0.8	1.2
13	1951	1954	8	3	18	10	0.0	1.2
14	1955	1962	1	7	2	17.5	0.0	0.5
15	1965	1971	3	6	9.5	14	0.0	0.2
16	1973	1975	2	2	9.5	4	0.0	0.9
17	1979	1986	4	7	9.5	17.5	0.0	0.0
18	1988	1993	2	5	9.5	10	0.0	0.0
19	1995	1999	2	4	9.5	10		

Note: The number of possible break points in the nature of cycles in industrial commodity prices is 18 (for booms) and 19 (for slumps). A boom (slump) is defined as a sequence of absolute increases (decreases) in prices. Values for the sup W statistic in excess of the 5 percent critical value (2.75 for booms and 2.77 for slumps) indicate rejection of the null hypothesis of no change in the duration of booms and slumps in real commodity prices. In the case of a tie, the relevant ranks are replaced by the average of the ranks of tied observations.

Figure 5. Real Price of Industrial Commodities, 1862-1999
(Logarithm)



Note: The shaded portion of the series indicates when commodity prices are in a boom period; the white (unshaded) portion indicates when commodity prices are in a slump period. A boom (slump) is defined as a sequence of absolute increases (decreases) in prices. The trend in prices is denoted by the thin line, and is derived using a fifth-order polynomial trend.

Definition 4: A *large cycle* includes one large boom and one large slump.

Using our definition of large booms and large slumps we find seven completed large cycles in the industrial commodity price index (Figure 6).[10] The peaks and troughs of these large cycles in industrial commodity prices are reported in Table 3, together with the durations of the large booms and slumps, their ranks, and the sup *W* statistic that tests the null hypothesis of no change in duration of large booms and large slumps. Cumulative price declines of 25 percent or more have occurred eight times over the period 1862-1999—the longest-lasting being the 106 percent decline in the two decades between 1951-71, and the sharpest being the 39 percent decline in the two years between 1973-75.

Table 3. Duration and Rank of Large Booms and Large Slumps
in Industrial Commodity Prices

Cycle	Peak	Trough	Duration (In years)		Rank		Sup *W* test	
			Boom	Slump	Boom	Slump	Boom	Slump
1	1862	1868		7		4.5		0.0
2	1907	1915	39	8	7	6	1.8	0.5
3	1918	1931	3	13	4	7	1.3	1.2
4	1951	1971	20	20	6	8	**2.5**	**3.6**
5	1973	1975	2	2	2	1	1.2	1.2
6	1979	1986	4	7	5	4.5	1.9	1.4
7	1988	1993	2	5	2	3	1.0	1.1
8	1995	1999	2	4	2	2		

Notes: The number of possible break points in the nature of large cycles in industrial commodity prices is 7 (for large booms) and 8 (for large slumps). A large boom (large slump) period is defined as a period during which industrial commodity prices increase (decrease) by at least 25 percent. Values for the sup *W* statistic in excess of the 5 percent critical value (2.39 for large booms and 2.44 for large slumps) indicate rejection of the null hypothesis of no change in the duration of booms and slumps in real commodity prices. In the case of a tie, the relevant ranks are replaced by the average of the ranks of tied observations.

[10] The assumed independence of durations appears to be confirmed for large booms and slumps, as the correlation between the length of successive large booms or the length of successive large slumps (over the entire sample) are not significantly different from zero at even the 20 percent level of significance. The correlation between the length of successive large booms (-0.27) and length of successive large slumps (0.02) is much less than the 20 percent critical value (0.52), calculated as $1.28/T^{1/2}$, where T=6 is the number of successive large booms and large slumps.

We find significant breaks in the duration of large booms and large slumps in the industrial price index (Table 3). The date of the change-point in large booms of the price of industrial commodities lies somewhere between the booms of 1931-51 and 1971-73; that is, in the period prior to the dissolution of the Bretton Woods system of fixed exchange rates.[11] The date of the change-point in large slumps of the industrial commodity price index lies between the slumps of 1951-71 and 1973-75; that is, 1972. The Wilcoxon rank sum test suggests that the duration of cycles in industrial commodity prices shortened significantly in the post-Bretton Woods period. Clearly, the duration of large booms and large slumps was shortened in the post-Bretton Woods period because of the increased frequency of large movements in the price index.

The post-Bretton Woods shifts in the duration of large cycles in commodity prices are economically, as well as statistically, significant. Our results indicate that while real industrial commodity prices spent about 62 percent of the post-Bretton Woods period in a large slump, only about 43 percent of the period prior to the early 1970s was spent in a large slump. In addition, the mean duration of the post-Bretton Woods large boom is *one-tenth* the mean duration of its pre-1970s counterpart, while the mean duration of the post-Bretton Woods large slump is *half* the mean duration of its pre-1970s counterpart.

Three conclusions emerge from our analysis of the long-run data. First, trends in commodity prices are highly volatile, which implies that for various subperiods we cannot tell whether differences in trend rates of growth are statistically significant. Moreover, because trend rates of growth in commodity prices are unstable, knowing the historical trend growth rate (over the full period or any subperiod) is of no practical policy relevance. Second, commodity price movements have become more variable over time—volatility first increased around 1899 and then again in the early 1970s. The first increase was due to price movements with bigger amplitudes. The second increase was due to a rise in the frequency of large price movements, which consequently reduced the duration of large price cycles. Third, long-run trends in commodity prices are small in comparison with annual variability in prices, making short-run movements in commodity prices highly unpredictable. In terms of its economic (and statistical) significance, price variability completely dominates long-run trends.

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[11] Major world currencies began to float against one another in the last quarter of 1971, following the United States' abrogation of the Bretton Woods gold clause in August 1971, which suspended convertibility of official dollar reserves into gold.

Figure 6. Real Price of Industrial Commodities, 1862-1999, Large Booms and Slumps
(Logarithm)

Note: The shaded portion of the series indicates when commodity prices are in a large boom period (price increases of at least 25 percent); the white (unshaded) portion indicates when commodity prices are in a large slump period (price decreases of at least 25 percent).

E. What Explains the Post-Bretton Woods Increase in the Variability of Commodity Prices?

Our results indicate that since the breakdown of the Bretton Woods exchange regime, real commodity prices have exhibited increasing variability. This finding confirms previous work, which postulated an important link between the growing volatility of nominal and real exchange rates since the early 1970s and the increased instability of the nominal and real (dollar-denominated) commodity prices (Chu and Morrison (1984), Reinhart and Wickham (1994), Cuddington and Liang (1999)). Indeed, an examination of the correlation between the volatility of the nominal U.S. dollar-U.K. pound exchange rate (as measured by the standard deviation of the growth rate) and the volatility of our measure of real commodity prices, across four major exchange rate regimes covering the period 1862-1999 (gold standard, interwar, Bretton Woods and post-Bretton Woods), reveals that it is very strong (0.99).[12]

What might explain this increasing volatility of real commodity prices during the post-Bretton Woods flexible exchange-rate regime? Some of the increase in the variability of nominal commodity prices is simply due to the mechanical effect of the denomination of commodity prices in a single currency (the dollar) at a time of fluctuating exchange rates. However, it is also likely that movements in the dollar exchange rate vis-à-vis other currencies affect the variability of dollar commodity prices, as such movements influence both the supply of, and demand for, commodities.

A possible link between more flexible exchange rate regimes, external debt and the increased volatility of dollar commodity prices has been suggested by Gilbert (1989). Mussa (1986) found that due to slowly-adjusting national price levels, real exchange rates in periods of nominal exchange rate flexibility are more volatile than in periods of nominal fixity. This increase in real exchange rate variability has real consequences in that it affects supply curves of commodity-exporting countries, and hence affects the level and variability of world prices for commodities. Gilbert (1989) argues that the interaction between flexibility of the U.S. dollar and dollar-denominated debt has affected the level and variability of commodity prices. For example, he finds that the appreciation of the dollar in the first half of the 1980s

[12] Using data for a group of 20 countries over 100 years, recent work by Taylor (2000) confirms the dramatically increased variability in both nominal and real exchange rates in the post-Bretton Woods period, in comparison with the gold standard and Bretton Woods fixed exchange rate periods. He also finds a very high correlation between real exchange rate volatility and nominal exchange rate volatility across all four exchange rate regimes.

led to greater than proportionate real depreciations of developing-country currencies, boosted the supply of commodities (due to the need to maintain export revenues to service growing debt obligations in the presence of credit rationing) and accordingly induced greater than proportionate reductions in dollar commodity prices. While demand conditions in importing countries are also a key determinant of commodity prices, recent explanations of the rising variability in nominal commodity prices have focussed on increased instability of the supply of commodities, induced in part by increasingly volatile real exchange rates and the dismantling of domestic and international price stabilisation schemes (see Reinhart and Wickham (1994)).[13] In the presence of sluggishness of the adjustment of national price levels (due in particular to slow adjustment of the price of nontradables), nominal exchange rate flexibility heightens the variance of real commodity prices.

V. Conclusion

This paper examined whether there have been changes in the long-run behavior of world commodity prices. Specifically, we looked at trends in commodity prices, the duration of price booms and slumps, and the volatility of price movements. While there has been a downward trend in real commodity prices of 1.3 percent per year over the last 140 years, little support is found for a break in the long-run trend decline in commodity prices. In contrast, there was evidence of a ratcheting up in price volatility over this period. There have been two periods of increased volatility—at the turn of the century and after 1971. While earlier studies have noted the increased volatility of commodity prices in the 20th Century, what was not explained was the nature of this change in volatility. Our results indicate that while the rise in volatility in the early 1900s was due to greater *amplitude* of price movements, the further rise in volatility in the early 1970s was due to the increased frequency of large price movements (that is, a fall in the *duration* of large price cycles).

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[13] While foreign supply of commodities is increased by a real appreciation of the dollar, foreign demand for commodities should fall, although it appears that most of the variability in commodity prices is driven by the supply side (Reinhart and Wickham (1994), Deaton and Miller (1996)). Dornbusch (1985) argued that a real appreciation of the dollar would increase the real price of any given commodity in terms of the foreign currency, reduce demand by the rest of the world and (for given supply) induce a fall in the commodity's real market-clearing price (expressed in U.S. dollars).

Importantly, within-sample trends in commodity prices are completely overwhelmed by the observed variance of price movements. Accordingly, in attempting to draw policy conclusions from the analysis of the behavior of commodity prices, misplaced concerns about long-run declines in price trends (in particular the Prebisch-Singer hypothesis) are much less important than concerns over the implications of increasing price volatility. In predicting movements in commodity prices, it is important not to be confused by the presence of large, long-lived booms and slumps—such cycles are highly unlikely to be indicative of any major change in long-run prices. While there is a downward trend in real commodity prices, this is of little practical policy relevance as it is small when compared with the variability of prices. In contrast, rapid, unexpected and often large movements in commodity prices are an important feature of the behavior of commodity prices. Such movements can have serious consequences for the terms of trade, real incomes and fiscal position of commodity-dependent countries, and have profound implications for the achievement of macroeconomic stabilization.

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