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

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*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 1862–1999 and have two main findings. First, while there has been a downward trend in real commodity prices of about 1 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. Although there is a downward trend in real commodity prices, this is of little practical policy relevance, since it is small and completely dominated by the variability of prices.* [JEL E32, Q11]

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

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 issue, since about 25 percent of world merchandise trade consists of primary commodities, and many developing countries depend on one or a few commodities for the majority of their export

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earnings (Cashin and Pattillo, 2000).[1] Both sharp fluctuations and long-run trend movements in commodity prices present serious challenges for many developing countries, owing 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.

The Prebisch-Singer hypothesis states that owing to the low income elasticity of demand for commodities and because total factor productivity increases have been smaller for manufactured goods than for primary commodities, the price of commodities relative to manufactured goods should decrease over time. 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.[2] While later work by Grilli and Yang (1988), using data from 1900–88, found that this downward trend accelerated in 1921, several papers failed 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 or a stochastic trend, or whether there are structural breaks in the trend. Although 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]

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, with prices sometimes changing 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, since they are subject to large and unpredictable movements that may have persistent effects. Because 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, since improving our understanding of the duration and amplitude of commodity-price cycles will be a key input in efforts to stabilize the macroeconomic effects of

[1]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 accounted for about 75 percent of total exports in 1997.

[2]Lipsey (1994) finds that once allowance is made for quality changes in manufactures (estimated to be an improvement of at least 0.5 percent per year), the case is weak that there has been a long-term deterioration in relative commodity prices.

[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.

movements in prices, particularly for commodity-dependent countries. Information on the average duration and amplitude of commodity-price cycles 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, understanding the cyclical behavior of commodity prices is as important as understanding their underlying long-run trends. Although 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. Variances lump together both the amplitude and duration of price movements, however, and it might also be useful to distinguish between these effects in examining cycles 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. To our knowledge, however, 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 of the series from trend (also known as the growth cycle). However, the definition of the cycle we use here follows Watson (1994) and Cashin, McDermott, and Scott (forthcoming) by dealing with the data in levels (also known as the classical cycle), hence avoiding the somewhat subjective choice of which detrending method to use. Price slumps (booms) are then described as periods of absolute decreases (increases) in the level of the series, not as a period of below-trend (above-trend) growth in the series.[4]

While it is easy to imagine what a booming or slumping market is, and despite such terms being used frequently to describe the state of commodity markets, there is no formal definition in the literature. One definition would describe a boom

[4]For a discussion on the relative merits of measuring cycles as either classical cycles (with turning points as peaks and troughs in the *level* of any given series) or growth cycles (with turning points as peaks and troughs in the *deviations from trend* of any given series), see Pagan (1997) and Harding and Pagan (2002). Canova (1998) also discusses the distortions introduced to the measurement of growth cycles by the various commonly used trend-removal filters (such as first-differencing and the Hodrick-Prescott filter). As our definition of the classical cycle relates to turning points in the level of the series, the decomposition of the series into its "trend" and cycle components is not relevant.

(slump) in commodity markets as a period of generally rising (falling) commodity prices. Accordingly, we work with a definition of booms and slumps in commodity prices that emphasizes movements in the level of commodity prices between local peaks and troughs. This approach is in line with the business-cycle literature going back to Burns and Mitchell (1946). The definition implies that commodity prices have shifted from a boom phase to a slump phase if prices have declined from their previous (local) peak (Cashin, McDermott, and Scott, forthcoming).

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

The issues raised previously can be resolved by comparing real 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 changes in both the amplitude and duration of price cycles) 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.

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 $\{X_1,..., X_N\}$ split into two subperiods, $\{X_1,..., X_{N_1}\}$ and $\{X_{N_1+1},..., X_N\}$. Thus the number of elements in the first subperiod is N_1 and the number of elements in the second subperiod is $N_2 = N - N_1$.

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

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

where the average annual growth rate in the i^{th} subperiod (i = 1, 2) is denoted by $\overline{X}_i = (1/N_i)\Sigma_i \Delta \ln X_t$, and we define $\Sigma_1 = \Sigma_{t=1}^{N_i}$; $\Sigma_2 = \Sigma_{t=1+N_i}^{N}$. We also define the autocorrelation-robust variance error estimate as

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

where the variance estimators of the pre-whitened residuals ($\varepsilon_t = u_t - \hat{\rho}_1 u_{t-1}$ for $t = 2,\ldots, N_1$ and $\varepsilon_t = u_t - \hat{\rho}_2 u_{t-1}$ for $t = N_1 + 2,\ldots, N$) are defined as $s_i^2 = (1/N_i - 2)\Sigma_i \varepsilon_t^2$; the autocorrelation coefficient is defined as $\hat{\rho}_i = \Sigma_i u_{t-1} u_t$; and the raw residuals are $u_t = X_t - \overline{X}_1$ for $t = 1,\ldots, N_1$ and $u_t = X_t - \overline{X}_2$ for $t = N_1 + 1,\ldots, N$. When the breakpoint is known, the t-statistic for testing the null hypothesis 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 commodity-price changes, the test statistic we use is defined as

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

where the variance of price changes in each subperiod is denoted by s_i^2 (for $i = 1,2$), as defined for equation (2). When the breakpoint is known and the two samples are independent, this test has an F-distribution with degrees of freedom $(N - N_1)$ and 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 time series, a boom phase is naturally defined as a period when the growth rate is positive; a slump phase is obviously when the growth rate is negative. For future reference, we introduce the following definitions:

Definition 1: For annual data, a *boom* is defined as a sequence of absolute increases in the level of prices, and a *slump* is defined as a sequence of absolute decrease in the level of prices.

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) that assumes the distributions of durations (of booms or slumps) differ only in their means. Their test is a version of the Wilcoxon rank-sum statistic, which is defined as[5]

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

where n_i denotes the number of boom phases 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 duration of boom phases in the first and second subperiods;[6] and the standard error is defined as

[5]The test is described in this section for a change in the duration of boom phases—the test for a change in the duration of slump phases is the same.

[6]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.

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

When the breakpoint is known, the statistic for testing the null of no change in the duration of booms can be used like a *t*-test that converges in distribution to a standard normal.[7]

Hypothesis Testing with Structural Change at an Unknown Break

The above-mentioned tests are conditional on structural change occurring at a known point in time. However, many researchers select the breakpoint after examining plots of the data. This procedure implies that the breakpoint is a function of the data, rather than being truly exogenous. Importantly, this endogenization of the breakpoint-selection procedure (however informal) will alter the distributions of the above test statistics. This means that searching for the breakpoint increases the probability of rejecting the null hypotheses of no change in trend growth, no change in volatility, and no change in the duration of booms or slumps. For example, suppose one tests for the date of a break in the duration of booms at the start of the First World War (with our test having a nominal size of 5 percent), and one repeats 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 both tests have a correct size of 5 percent, the probability that at least one of these points has a break (assuming the two events are independent) must be higher than 5 percent if we use the same critical values as the individual tests (Cashin and McDermott, 2002). As a result, if we allow all points in time to be potential breaks, then the probability of a break occurring 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. Accordingly, if one takes the view that the breakpoints are unknown, then our tests for change in trend, duration, and volatility will need to account for the fact that the breakpoints are dependent on the data. Our testing procedure is to search over all possible breakpoints in the sample and select the largest of these tests, which we denote as the sup Q test. That is, we are interested in test statistics of the form

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

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]. Suppose that for a

[7]See Lehmann (1975) for a general discussion of the Wilcoxon rank-sum statistic. In addition to the standardized form of the statistic having an asymptotically standard normal distribution, it has an exact sampling distribution, which can also be deduced. In large samples, it is not necessary to use the Wilcoxon rank-sum statistic, since the standard *t*-test can be used to compare average durations. We prefer to use the rank-sum statistic, however, to ensure the comparability of our results with earlier work.

given π the test statistic $Q(\pi)$ has a distribution that can be represented by the probability density function $f(x)$ and that has a cumulative density function $F(x)$. Then the cumulative distribution of sup $Q(\pi)$, if the Q statistics are independently drawn, is

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

where m is such that Pr $(\sup(x) < m) = 1 - \alpha$ and α is the size of our test.[8] 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 a change in the duration of booms and slumps, then sup $Q(\pi)$ will follow the distribution given by the probability density function

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

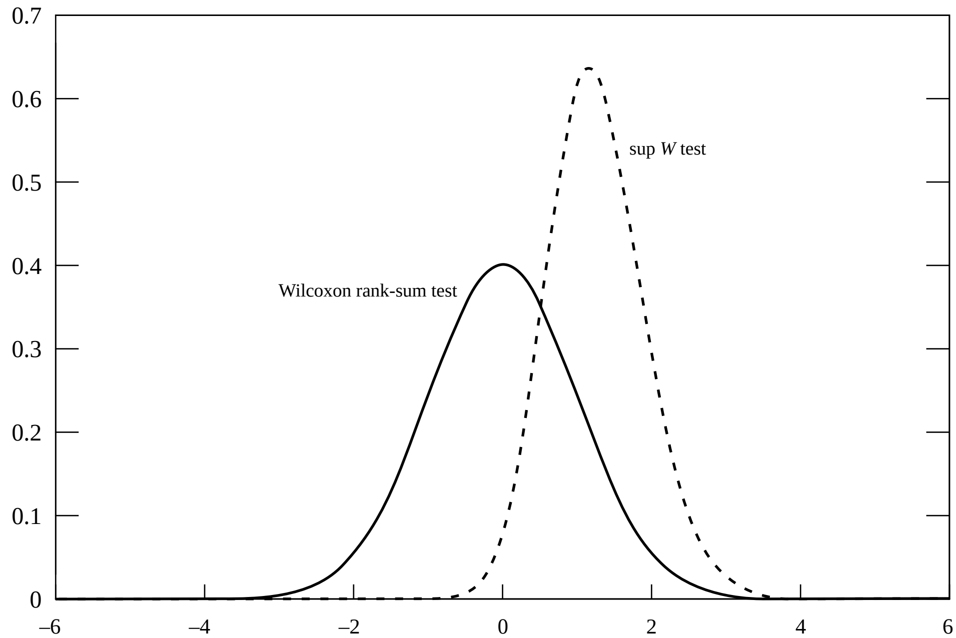
The distribution shown in equation (9), for the case when $n = 7$, is compared with the standard normal distribution in Figure 1. It can clearly be seen that the effect of searching over all possible breakpoints is to shift the distribution of the Wilcoxon rank-sum statistic to the right. Without making this correction, searching over breakpoints would increase the probability of falsely finding a break in the duration of price booms or slumps. 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 breakpoint. Intuitively, the location of the maximum $Q(\pi)$ statistic (denoted as $\hat{\pi}$) will be a reasonable estimator for the location of the breakpoint. That is, our estimator of the breakpoint is

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

[8]If the Q-statistics are dependent, then the true size of our test will be too small, and we will have an overly conservative test. In this case, any rejection of the null hypotheses (of no change in the trend rate of growth of commodity prices, no change in the duration of commodity price cycles, or no change in the variance of commodity prices) will be even stronger evidence of the existence of a change. To examine the sensitivity of our assumption of the independence of the Q-statistics, we conducted a bootstrap experiment (of 500 iterations) to compute the 5 percent critical values for the null hypotheses of no change in the duration of booms and slumps. We find that the bootstrap (5 percent) critical values are about 2.6 for changes in the duration of both booms and slumps; these are slightly lower than those derived from the approximate critical values of equation (9), both of which are about 2.8.

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



Source: Authors' calculations.

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 with 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 breakpoints located arbitrarily close to the ends of the sample.[9]

The principle behind the nonparametric tests that we have used is similar to the parametric tests used in Andrews (1993), Hansen (2000), and Khan and Senhadji (2001). That is, both the parametric and nonparametric methods search over all possible breakpoints to determine the optimal location to use in splitting the sample. In choosing between these approaches, however, there are the usual trade-offs between parametric and nonparametric methods. That is, the nonparametric method has the relative disadvantage of slower rates of convergence and less-efficient estimation, but the relative advantage of less need to specify the functional

[9]Let $\Pi \subset [0, 1]$ be the set of all possible breakpoints searched over. When searching for breakpoints in the duration of booms and slumps, no trimming of the cycle data is undertaken, that is $\Pi = [0, 1]$. When searching for breakpoints in growth rates or volatility, $\Pi = [0.1, 0.9]$. For our data series (1862–1999), this means that the first 15 and the last 15 observations will be excluded from being possible breakpoints in growth rates or volatility. Testing whether a sample has a breakpoint at its extremities is not desirable, because statistics tend to diverge to infinity since they cannot discriminate between true breakpoints and boundary conditions.

form of the model. This advantage of nonparametric methods would seem to be particularly important when testing for changes in the duration of cycles.

II. Data

In analyzing commodity-price movements, we use the longest dataset publicly available—the industrial commodity-price index of *The Economist*. The real annual data consist of the nominal industrial commodity-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. For 1862–1910, the annual data are formed as an average of price observations in January and July of each year; from 1911 onward, the annual data are an average of observations for all months. The industrial commodity-price index consists of the prices of textiles, metals, and nonfood industrial commodities. For details on the composition and construction of the index, see the appendix.[10]

Figure 2 sets out both the nominal and real industrial commodity-price indices, with the latter showing the potential command over U.S. resources of a fixed quantity of industrial commodities. Over the period 1862–1999, nominal prices of industrial commodities have increased about threefold. In the four decades prior to the First World War, and in the two decades prior to 1972, nominal prices of industrial commodities were remarkably stable. Movements in the nominal price index reflect several historical events—the rise in cotton prices in the wake of the shortfall in supplies caused by the U.S. Civil War of the 1860s, the rapid rise in commodity prices during and shortly after the First World War and their sharp decline in the Great Depression years of the 1930s, the steep rise in prices accompanying the Second World and Korean Wars, the upturn in prices during the commodity-price booms of the 1970s, and the sharp boom and slump periods in nominal commodity prices during the 1980s and 1990s.

The index of real industrial prices, after being somewhat stable over its first four decades, fell by four-fifths between 1900 and 1999, ending the century at a record low (Figure 2). The purchasing power over American resources of a basket of industrial commodities in 1999 is one-fifth the purchasing power yielded by that same basket of commodities in 1862. While the purchasing power of industrial commodities was slightly lower at the turn of the century than 40 years earlier, and by the early 1950s was about 20 percent lower than in the early 1860s, in the latter half of the twentieth century real industrial prices steadily declined. The largest annual declines in commodity prices were 37 percent from 1930 to

[10]The 2000 version of *The Economist's* dollar-based industrial commodity-price index (base 1995 = 100, weighted by the value of world imports during 1994–96) comprises 16 commodities, with weights 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 (fine wool), wool 48s (coarse wool), palm oil, coconut oil, soybeans, and soybean oil.

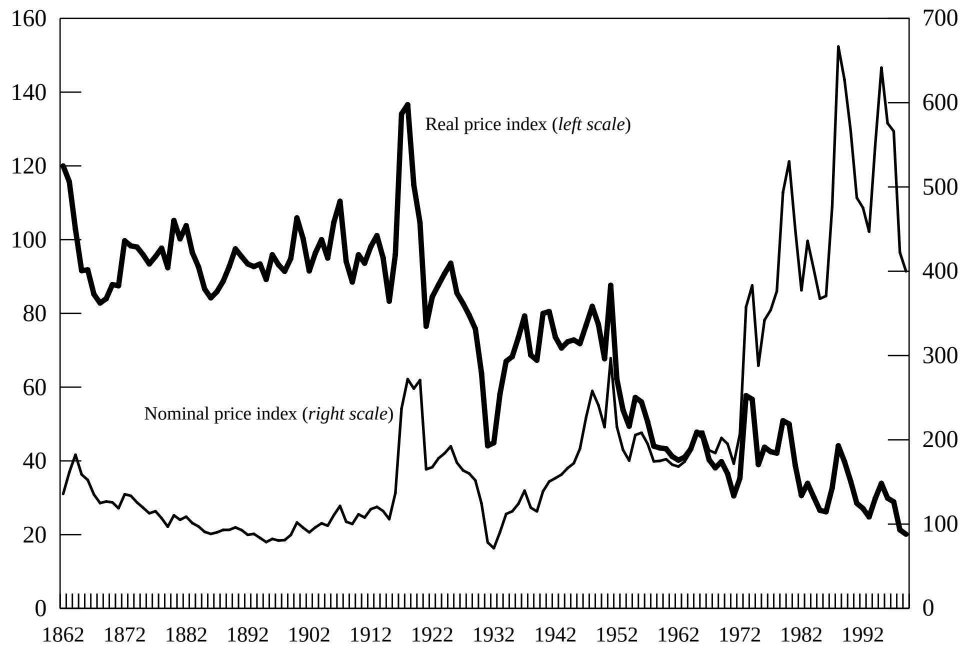
Table 1. Industrial Commodity Price Facts, 1862–1999

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]

Source: Authors' calculations.

Notes: The underlying data are an (annual) index number of *The Economist's* real industrial commodity prices, calculated as *The Economist's* nominal industrial commodity-price index (dollar-based with base of 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 (see Section II and the appendix for details). 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, $\ln(1/2)/\ln(\alpha)$ where α is the autoregressive parameter from an AR(1) model for the levels (for the third column) and from an ARCH(1) model for the variances (for the seventh column). For $\alpha \geq 0$, 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 the third and seventh columns. 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 denotes the r^{th} (central) moment. The skewness of a symmetrical distribution, such as the normal distribution, is zero; similarly, the kurtosis of the normal distribution is 3.

Figure 2. Nominal and Real Price Indexes for Industrial Commodities, 1862–1999



1931 and from 1974 to 1975, and 34 percent between 1951 and 1952. Conversely, the largest annual rises in commodity prices were 49 percent from 1972 to 1973, and 33 percent between 1916 and 1917.

Unlike most previous studies of long-run commodity prices, ours does not use the Grilli-Yang (1988) index of real commodity prices. The Grilli-Yang index runs from 1900 through 1988, while *The Economist's* industrials index covers four more decades of data and provides 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's* industrials index is very high. Over the period used by Grilli and Yang the correlation between the two series is 0.81, and is even higher (0.85) when the Grilli and Yang data are extended until 1999.[11]

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's* real industrial commodity-price series (in levels) for both its full history and for three subperiods. Following Eichengreen (1994), Reinhart and Wickham

[11]One key difference between the Grilli-Yang index and *The Economist's* industrials index is that the former trends downward more slowly than the latter. This is largely because in addition to industrial commodities, the Grilli-Yang index includes food commodities (especially beverages), which have displayed a substantially smaller trend decline than non-food commodities over the sample period (see Grilli and Yang, 1988).

(1994), and Cuddington and Liang (1999), the subperiods have been selected to analyze the behavior of commodity prices across exchange rate regimes, since fixed and flexible exchange rate regimes have been found to yield differing volatility of real commodity prices.[12] 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 that followed the Bretton Woods period (1972–99).

For the full sample, both the first- and second-order autocorrelation coefficients are quite high, indicating that prices tend to revert to their means 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 commodity-price index—that is, the length of time it takes for half of any initial shock to dissipate (see Cashin, Liang, and McDermott, 2000). The half-life of reversion 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 persistence of shocks in the flexible exchange rate regime appears to be less than for the fixed exchange rate regimes.

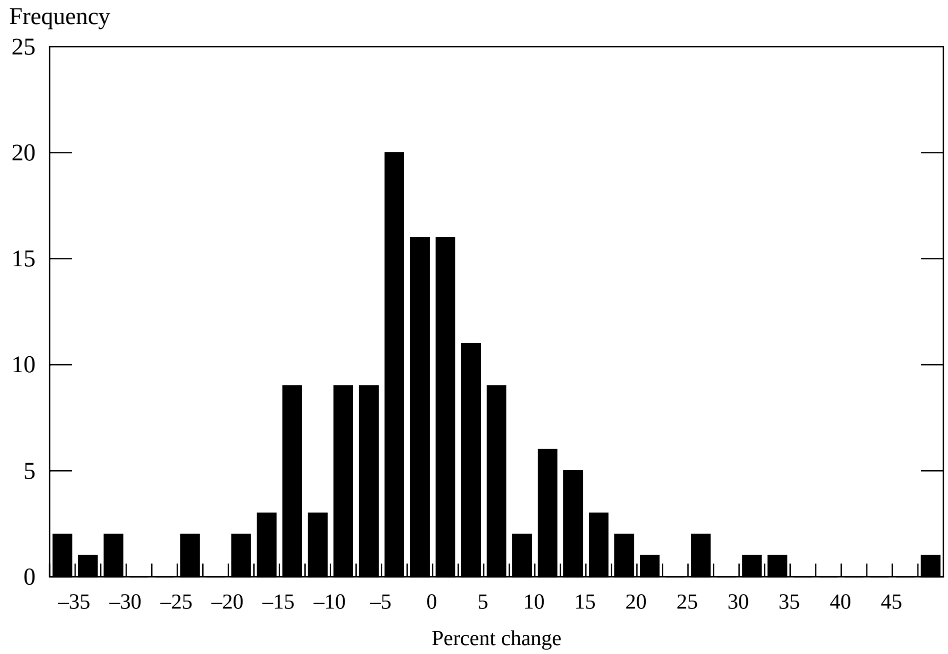
Consistent with earlier findings by Deaton and Laroque (1992), the coefficient of variation reveals that there is substantial annual variability in the real price of industrial commodities, particularly in the flexible exchange rate regime. There is negative skewness in commodity prices for the full sample period, since 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 its 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 (tails thinner than the normal distribution) of prices during the gold-standard period, indicating that large price movements were relatively uncommon. The final (seventh) 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 in less than a year. Accordingly, there is little evidence of clustering of large or small variances in commodity prices.

Figure 3 presents a histogram (frequency distribution) of the (absolute) change in the real price of industrial commodities. The distribution is centered at a little less than zero but has a high variance (standard deviation of 12.6 percent per year) around its mean trend decline (–1.3 percent per year), a slightly larger median decline (–2.0 percent per year), and fat tails (especially for price declines). Clearly, the large dispersion in price changes dominates the relatively small secular decline in real commodity prices.[13]

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

[13]While this point estimate of the annual average rate of decline of relative commodity prices is larger than those typically derived for the ratio of commodity prices to manufactured goods (estimated at about 0.7 percent per annum) using the Grilli-Yang (1988) data, it is close to those obtained using alternate relative commodity-price data (see Sapsford, 1985, and Lutz, 1999).

Figure 3. Histogram of Percentage Change in Real Commodity Prices



III. Empirical Results

In this section, we present the results from our empirical analysis of the properties of the index of the real price of industrial commodities. Specifically, we will test the null hypotheses of no change in the trend growth of commodity prices, no change in the variance of commodity prices, and no change in the duration of commodity-price cycles.

Trends

The salient feature of the real price of industrial commodities is that it has a downward trend over 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 break-points 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 (1989). 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 and 1999, more than 5 out of every 100 annual price movements exceeded 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.

Volatility

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

In examining changes in the volatility of the real 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 nineteenth century to the start of the First World War. In fact, the *F*-statistic as a function of breakpoints has a local peak in 1914 (with a value of 6.23).

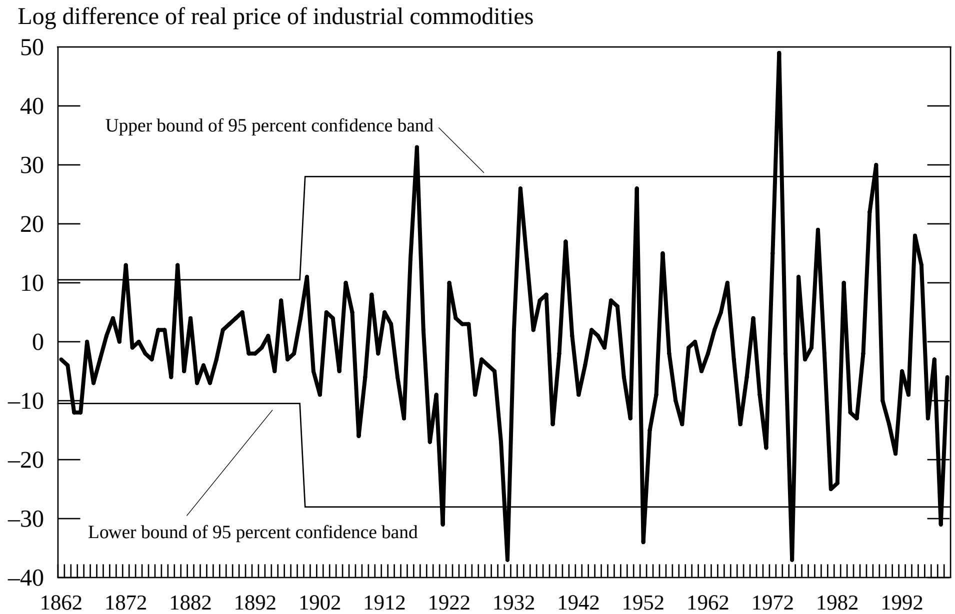
Figure 4 also shows the 95 percent confidence bands around the growth rate movements. For industrial commodities, the bands widen sharply after the change-point in 1899. Importantly, the frequency with which price movements exceed two standard deviations appears to increase after 1971. In the roughly 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.[14]

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 real commodity prices, two features of the commodity-price data need to be verified. First, the duration of booms and slumps needs 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

[14]Part of the change in the volatility of the price of industrial commodities could be due to the manner in which the industrial price index has been constructed (in particular, to changes in the commodity composition of the index), as argued by Romer (1986) in the context of U.S. employment data for 1900 through 1980. However, Cuddington and Liang's (1999) analysis of the Grilli-Yang data on relative commodity prices—which are measured using a uniform composition between 1900 and 1992—also shows an increase in volatility after the Bretton Woods period.

Figure 4. Volatility of Real Price of Industrial Commodities, 1862–1999



cycles are delineated in this paper on a non-trend-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 duration assumption appears to be confirmed, as the correlations 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.[15] As to the assumption of no difference in trend growth, we have demonstrated previously 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.

In this section, we examine the stability of real industrial commodity prices in terms of the relative *duration*, rather than relative volatility, of the commodity-price cycle. This duration perspective focuses on the *lengths* of boom and slump phases of the commodity-price cycle, while the (preceding) volatility analysis focuses on the amplitude of commodity-price movements.[16]

[15]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 much less than the 15 percent critical value, where the 15 percent critical values are 0.35 for booms and 0.34 for slumps, calculated as $1.44/T^{1/2}$, where $T = 17$ ($T = 18$) is the number of successive booms (slumps).

[16]This idea is developed more fully in Diebold and Rudebusch (1992).

Table 2. Duration and Rank of Booms and Slumps in Industrial Commodity Prices, 1862–1999

			Duration (*in years*)		Rank		sup *W* test	
Cycle	Peak	Trough	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		

Source: Authors' calculations.

Notes: The number of possible breakpoints in the nature of cycles in real industrial commodity prices is 18 for booms and 19 for slumps. A boom (slump) is defined as a sequence of absolute increases (decreases) in the level of prices. Values for the sup *W* statistic exceeding 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.

Using our definition of booms and slumps, we find there are 18 completed cycles in industrial commodity prices, where a cycle includes one boom and one slump, over the full sample period. 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*-statistics for duration change in booms (2.1) and slumps (2.5) were 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, while the unshaded portions denote the slump phase, of each cycle. Consistent with earlier work on commodity-price cycles (Cashin, McDermott, and Scott, forthcoming), on average price slumps (4.2 years) last longer than price booms (3.6 years).

Large Cycles

Clearly, there has been an increase in the last 30 years in the frequency with which large price movements have occurred. 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, which are formally defined as follows:

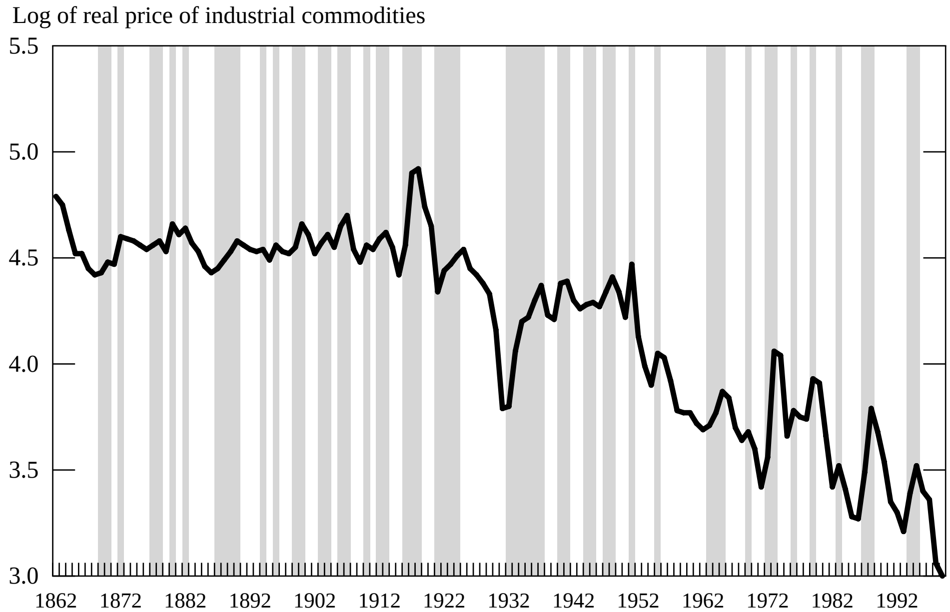
Definition 3: For annual data, a *large boom* is defined as a sequence of generally increasing prices that have had a price movement of at least 25 percent over the phase, and a *large slump* is defined as a sequence of generally decreasing prices that have had a price movement of at least 25 percent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

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 real price of industrial commodities (Figure 6).[17] 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 covering 1951–71 and the sharpest being the 39 percent decline between 1973 and 1975.

[17]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 values (0.52 for large booms and 0.48 for large slumps), calculated as $1.28/T^{1/2}$, where $T = 6$ ($T = 7$) is the number of successive large booms (large slumps).

Figure 5. Cycles in Real Price of Industrial Commodities, 1862–1999



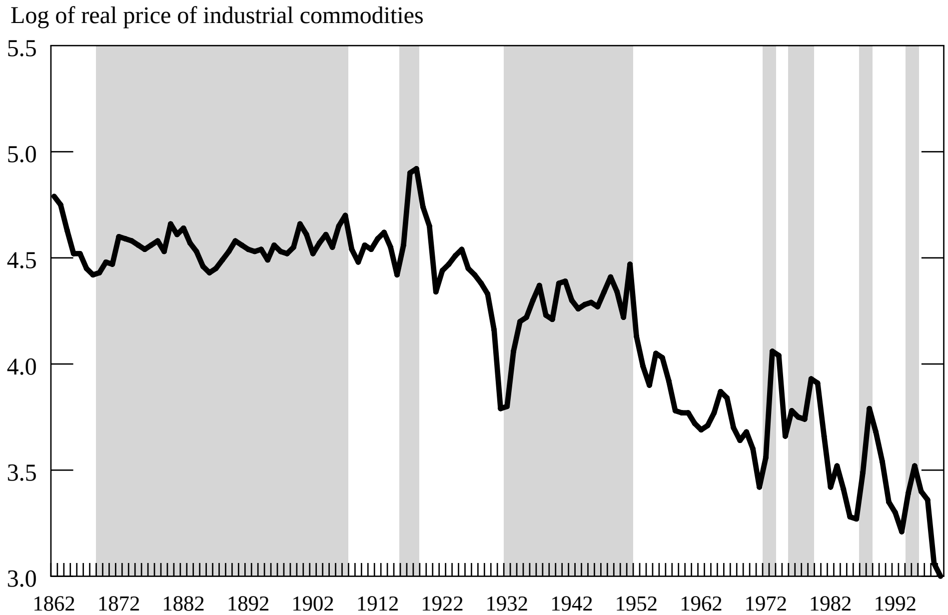
Notes: Shaded portions denote the boom phase, and unshaded portions the slump phase, of each commodity-price cycle. A boom (slump) is defined as a sequence of absolute increases (decreases) in the real price of industrial commodities.

We find significant breaks in the duration of large booms and large slumps in the price of industrial commodities (Table 3). The date of the changepoint in large booms of the price of industrial commodities lies somewhere between the booms of 1931–51 and 1971–73; that is, during the period before the dissolution of the Bretton Woods system of fixed exchange rates.[18] The date of the changepoint in large slumps of the industrial commodity price index lies between the slumps of 1951–71 and 1973–75; that is, in 1972. The Wilcoxon rank-sum test suggests that the duration of cycles in industrial commodity prices shortened significantly after the Bretton Woods period. Clearly, the duration of large booms and large slumps has been 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 although 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

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

Figure 6. Large Cycles in Real Price of Industrial Commodities, 1862–1999



Notes: Shaded portions denote the boom phase, and unshaded portions the slump phase, of each large commodity-price cycle. A large boom (large slump) is defined as a period during which the level of real industrial commodity prices increases (decreases) by at least 25 percent.

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 *one-half* the mean duration of its pre-1970s counterpart.

Three conclusions emerge from our analysis of the long-run data. First, trends in real 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, real 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 real 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.

Table 3. Duration and Rank of Large Booms and Large Slumps in Industrial Commodity Prices, 1862–1999

			Duration (*in years*)		Rank		sup *W* test	
Cycle	Peak	Trough	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		

Source: Authors' calculations.

Notes: The number of possible breakpoints in the nature of large cycles in real industrial commodity prices is 7 for large booms and 8 for large slumps. A large boom (large slump) is defined as a period during which the level of industrial commodity prices increases (decreases) by at least 25 percent. Values for the sup *W* statistic exceeding 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 large booms and large 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.

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 1862–1999 (including the gold-standard, interwar, Bretton Woods, and post-Bretton Woods periods), reveals that it is very strong (0.99).[19]

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. It is also likely, however, that movements in the dollar exchange rate vis-à-vis other currencies affect the variability of dollar commodity prices, since such movements influence both the supply of, and the demand for, commodities.

A possible relationship among 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 owing to slowly adjusting national price levels, real exchange rates are more volatile in periods of nominal exchange rate flexibility than in periods of nominal fixity. This increase in real exchange rate variability has real consequences, in that it affects the 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 led to greater-than-proportionate real depreciations of developing-country currencies, boosted the supply of commodities (owing 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. Although demand conditions in importing countries are also a key determinant of commodity prices, recent explanations of the rising variability in nominal commodity prices have

[19]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. Taylor also finds a very high correlation between real exchange rate volatility and nominal exchange rate volatility across all four exchange rate regimes.

focused 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 stabilization schemes (see Reinhart and Wickham, 1994).[20] In the presence of sluggishness in the adjustment of national price levels (owing in particular to slow adjustment of the price of nontradables), nominal exchange rate flexibility heightens the variance of real commodity prices.

IV. Conclusion

This paper has examined whether there have been changes in the long-run behavior of world commodity prices. Specifically, we looked at trends in real commodity prices, the duration of price booms and slumps, and the volatility of price movements. Although there has been a downward trend in real commodity prices of about 1 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 beginning of the twentieth century and after 1971. While earlier studies have noted the increased volatility of commodity prices in the twentieth century, what was not explained was the nature of this change in volatility. Our results indicate that although 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, to a fall in the *duration* of large price cycles).

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, 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 indicate any major change in long-run prices. Although there is a downward trend in real commodity prices, this is of little practical policy relevance, because 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 their behavior. Such movements can have serious consequences for the terms of trade, real incomes, and fiscal positions of commodity-dependent countries, and have profound implications for the achievement of macroeconomic stabilization.

[20]While the 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; and 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 a given supply) induce a fall in the commodity's real market-clearing price (expressed in U.S. dollars).

APPENDIX

The Economist's Industrial Commodity-Price Index

The Economist's U.S. dollar-based industrial commodity-price index covers 1862–1999, and has been calculated by the compiler as a linked version of price indices first published in 1864 (see *The Economist*, "A Raw Deal for Commodities," April 17, 1999).

The commodity coverage of this index has been revised several times over this long period. Up to 1911, the index comprised the unweighted average (arithmetic mean) of 16 industrial commodities traded on British and U.S. wholesale markets, with prices converted (where necessary) into dollars at the current rate of exchange. It consisted of the following nonfood commodities: fibers (two types of raw cotton, yarn, cloth, raw silk, flax and hemp, wool, and indigo); metals (pig iron, copper, lead, and tin); and other raw materials (oils, leather, tallow (candle wax), and timber). The 1911 revision broadened the composition of the index to include 28 industrial commodities; the additions were steam coal, house coal, iron and steel bars, jute, petroleum, oilseeds, rubber, U.S. timber, Australian wool, flax, hemp, and soda crystals. A historical comparison of the two commodity-price indices indicated that the revised composition altered the level of the nominal index only slightly (see *The Economist*, "Our New Index Number," November 18, 1911).

From the 1960s onward, commodities included in the index were weighted by their relative shares of developed-country (that is, members of the Organization for Economic Cooperation and Development's) imports of industrial commodities and reweighted annually according to the most recent annual pattern of trade. Prior to the First World War and after the Second World War, the index was typically calculated on an arithmetic-mean basis; between the wars, it was calculated on a geometric-mean basis.

The last major revision to the composition of the industrial commodity-price index occurred in 1984, when copra and groundnuts were dropped and aluminum, nickel, and timber were added to the index. The 1984 version of *The Economist's* industrial commodity-price index comprised 6 metals (aluminum, copper, nickel, zinc, tin, and lead) and 12 nonfood agricultural goods (wool 64s (fine wool), wool 48s (coarse wool), cotton, jute, sisal, timber, hides, rubber, palm oil, coconut oil, soybeans, and soybean oil).

For additional details, see *The Economist* ("Our New Index Number," November 18, 1911; "Our Index Number of Wholesale Prices," December 15, 1928; "A Commodity Price Indicator," July 19, 1952; "Commodity Prices," March 2, 1974; "Commodities Brief: *The Economist* Indicators," February 18, 1984; "A Raw Deal for Commodities," April 17, 1999; and "A Changed Commodity," January 15, 2000).

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