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Commodity and Manufactures Prices in the Long Run

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Abstract

The low level of primary commodity prices since 1985 is examined in the context of the behavior of those prices relative to prices of manufactured goods since 1854. The Prebisch-Singer hypothesis of a secular decline in relative commodity prices is sustained, but the recent decline is shown to be well outside the realm of historical experience. Commodity and manufactures prices are found to be cointegrated, conditional on the negative trend and a number of unexplained short-term swings. The earlier finding of a Gibson paradox is explained in terms of the difference between short- and long-run relationships.

JEL Classification Numbers:

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Summary

Primary commodity prices have been depressed since the mid-1980s, with severe consequences for countries that depend heavily on commodity export revenues. The causes of this problem are manifold. At the global level, the depreciation of the U.S. dollar tended to raise the dollar prices of manufactured goods but probably not those of primary commodities. The debt crisis may have contributed to the problem by increasing the pressure on developing countries to raise exports. At the micro level, there have been many developments with adverse effects on the markets for specific commodities. These factors have included health and environmental concerns, the weakening of international marketing arrangements, bumper crops, and cyclical increases in metals supplies.

Rather than attempting to quantify the relative importance of these various disturbances, this paper examines recent developments in the context of longer-run movements in commodity prices. Based on data for the prices of primary commodities and manufactured goods in international trade since the middle of the nineteenth century, it is shown that there has been a stable negative trend in the relative price of commodities amounting to around 1/3 of one percent per annum, or about 40 percent over the past 120 years. The economic importance of this trend is difficult to assess, because of the impossibility of accurately measuring quality changes over such a long period. What seems more disturbing is that there has been a secular tendency for commodity prices to become more volatile, with adverse consequences for consuming as well as producing countries. This tendency first appeared around the time of the first World War, but it has become much more marked since the early 1970s.

The relative price of commodities in 1990 is estimated to be at the lowest level ever recorded: some 60 percent below the level that would have prevailed in the absence of the negative trend and major disturbances to market conditions. Taking the negative trend into account, it appears that the recent decline is not quite unprecedented. There have been three times in this century that the relative commodities price has fallen more than thirty percent below trend: in the aftermath of World War I, after which commodity prices rapidly recovered in the boom period of the 1920s; at the onset of the Great Depression, after which commodity prices remained relatively low until the end of the Second World War; and now. History thus seems to suggest that strong growth in industrial countries is required for a recovery in commodity prices. The estimates in this paper, however, must dampen even that degree of optimism, because the recent decline is not attributable to slow growth and is largely outside the realm of previous experience.

"... how long most people would look at the best book before they would give the price of a large turbot for it!"

Ruskin (1865)

I. Introduction

Over the decades since Ruskin lamented the limited demand for books, it would seem that fishmongers have fared rather less well than booksellers. In March 1991, large turbot were fetching around \$10 at the Washington, DC waterfront, while farther uptown, Ruskin's essays were going for \$35 and up in hardbound editions. But how general is this phenomenon: Has there been a secular tendency for the prices of primary commodities to decline relative to those of manufactured goods? That hypothesis was advanced some forty years ago by Raul Prebisch and Hans Singer, but empirical tests of it have been inconclusive. This paper takes yet another look and concludes that the decline has been significant, substantial, and prolonged. The massive and widespread decline in relative commodity prices since the mid-1980s, however, has been of an unprecedented scale and cannot be seen as a continuation of earlier tendencies.

The tests discussed below are based on index numbers for the prices of non-fuel primary commodities and of manufactured goods over virtually the whole of the industrial age (1854-1990). ^{1/} The data, which are described in some detail in Appendix I, are illustrated in Figure 1. Both series have had similar trends, and in most cases the broad cyclical swings have been similar as well. Nonetheless, there have been some very large movements in the relative price of primary commodities. When both the first two decades (the 1850s and 1860s) and the last two (the 1970s and 1980s) are taken into account, there appears to have been a substantial negative trend, averaging about 0.3 percent per annum over the full period. Most of the decline, however, occurred either in the 19th century or since the mid-1970s, and it is not obvious that the trend is significant, nor that there is an absence of a tendency toward mean reversion.

There are many obvious questions that could be raised about the quality of the data: the sources, composition, and weighting scheme all have changed over time, and quality shifts have been allowed for only implicitly and cursorily. To recall the initial example, if one were to treat paperback editions of Ruskin's (or anyone else's) essays as equivalent to his "best book," one might conclude that relative prices were relatively unchanged since the 1860s. These quality issues, however, are empirically

^{1/} For an even longer picture of primary commodity prices, see Commodity Research Bureau (1939, pp. 28-29), which shows a chart of an annual U.S. wholesale commodity price index covering the period 1720-1940. To my knowledge, there are, however, no estimates of prices of manufactured goods prior to 1854.

unresolvable, and the net direction of their impact is ambiguous. For a good discussion of the issues and for further references, see Grilli and Yang (1988).

In addition to tests of the significance of the secular trend in relative prices, tests are presented below for cointegration of the levels of commodities and manufactures prices, which would have implications for judging whether there has been a strong tendency for unusual movements in relative prices to be reversed rather than persisting. The paper also examines whether there has been a persistent relationship between the level of commodity prices and the rate of change in manufactures prices. Such a relationship is predicted by at least one popular theoretical model, resulting from the tendency of commodity prices to respond relatively quickly to shocks. This issue has implications for understanding both the theory and the observed tendencies in the data. Finally, a reduced-form model is estimated over the period 1921-1990, and these estimates suggest that the extent of the recent weakness of commodity prices cannot be explained as a continuation of previous relationships.

II. A Theoretical Model of Short- and Long-Run Relationships

To study the determinants of commodity prices in the aggregate requires acceptance of a reduced-form approach. The price indexes examined in this paper lump together more than thirty primary commodities as diverse as coffee and copper, and it would be futile to attempt to develop a structural model of the market forces influencing these indexes. An aggregated commodity-price model was developed by Frankel (1986), based on Dornbusch's (1976) model of the jumping behavior of what have come to be called "flex-prices." ^{1/} Frankel's model was then extended by Boughton and Branson (1990, 1991) to cover the possibility that primary commodities might serve either as inputs to the production of manufactures or as substitutes for manufactures in final consumption; and by Boughton, Branson, and Muttardy (1989) to analyze the relationship between commodity prices and exchange rates. The general flexprice model starts from macroeconomic relationships characterizing the markets for commodities and manufactures, from which may be derived (see Appendix II) a dynamic equation of the form

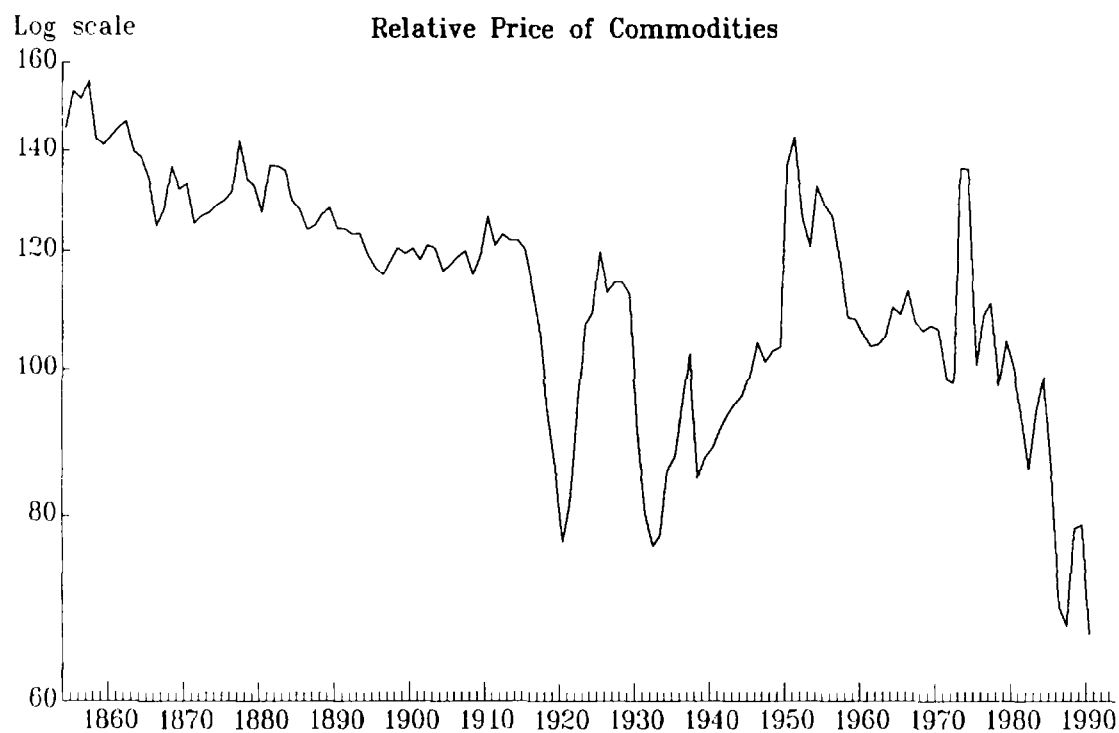
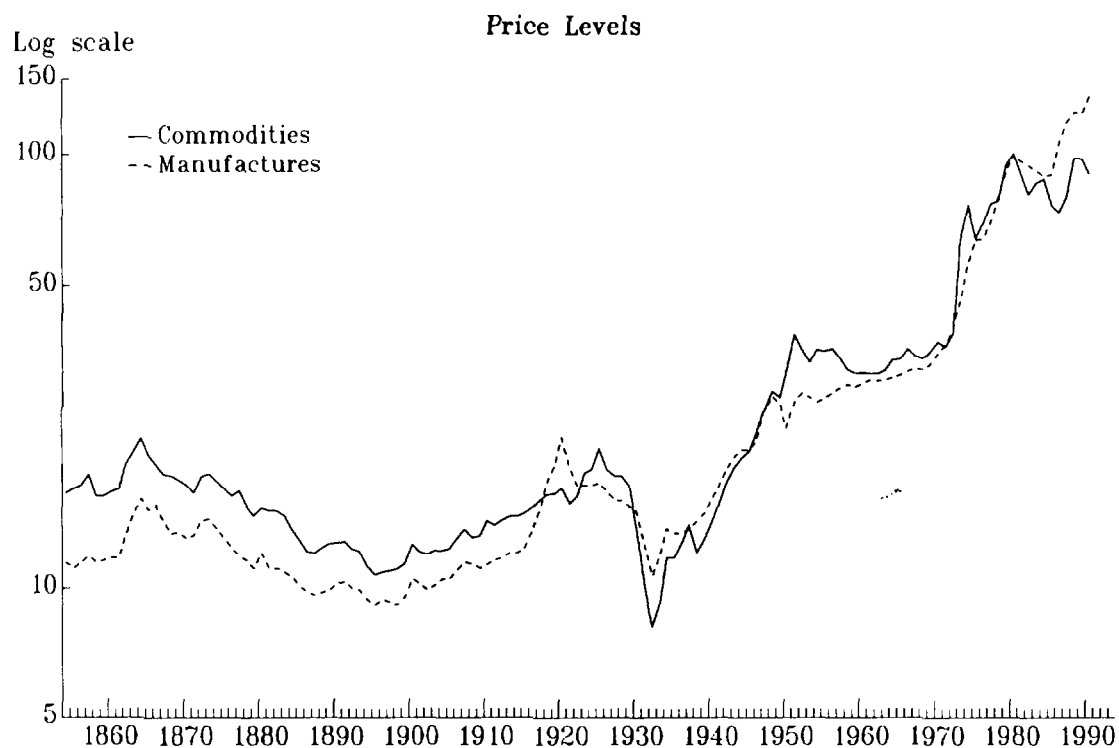
$$\dot{p}_r = \phi(p_r, y, b, i), \quad (1)$$

- + - +

where p_c and p_m are logarithms of indexes of primary commodity and manufactures prices, respectively; $p_r \equiv p_c - p_m$ (the relative price of primary commodities); y is the logarithm of real output in industrial countries; b is the real return to holding commodities for final use (net of storage costs) which could be positive or negative and which may change

^{1/} For other, more detailed, strategies for modeling commodity prices, see Brayton et al. (1990), Hook and Walton (1989), Sapsford (1987), and Winters and Sapsford (1990).

Figure 1. Commodity and Manufactures Prices, 1854-1990
(1980 = 100)



Relative price = commodity price index divided by manufactures price index multiplied by 100.

deterministically or stochastically over time; and i is the nominal interest rate.

An implication of this model is that commodity and manufactures prices should be cointegrated. More specifically, the relative price of commodities should be a stable long-run function of a few variables; it may be trended, however, owing to trends in the determining variables. The major empirical difficulty that has arisen in assessing cointegration has been that the model implies that the short-run relationships may be very different from those that would prevail in the long run. In the short run, the dominant relationship is that when the commodity price level is high, the inflation rate of the manufactures price should be high, because of the tendency of the former to jump in response to a monetary disturbance, followed by a more gradual response by the manufactures price. A relationship of this Gibson-paradox type was uncovered in Boughton, Branson, and Muttardy (1989), based on monthly data covering the period 1957-88. In the longer run, however, the common trends in the data would become more prominent. Consequently, any tendency toward cointegration might appear only in the long run.

The short- vs. long-run distinction may be illustrated by reference to some artificially constructed data that are generated according to a flexprice model. The constructed data are shown in Figure 2. ^{1/} The heavy line in the top panel is the logarithm of the commodity price level, which has a constant upward trend but which periodically jumps in response to shocks. Implicitly, these shocks are monetary in nature and do not alter the equilibrium relative price. The lighter curve is the price of manufactures, which responds gradually in response to the same shocks. Note that whenever the commodity price level is high relative to its trend, the CPI inflation rate is positive, and conversely; that relationship is illustrated in the middle panel.

The bottom panel of Figure 2 plots the logarithm of the relative price of commodities. By construction, this relative price has no trend; it fluctuates in a regular pattern around a stable mean. Because of the continuing pattern of shocks, the relative price does not settle down to an equilibrium level; in that sense, it is not stationary.

^{1/} The data are constructed as follows [where for a number such as $x = 9/2$, the notation is that $\text{integer}(x) = 4$ and $\text{remainder}(x) = 1$]:

$$i_t = \text{integer}(t/10) + 1, \quad t = 1, \dots, 1200;$$

$$j_1 = 1, \text{ and } j_t = \text{remainder}(i_{t-1}/2), \quad t = 2, \dots, 1200;$$

$$k_t = 5 \text{ if } j_t = 0, \text{ and otherwise } k_t = 1;$$

$$p_{ct} = k_t + 0.1t;$$

and $p_{mt} = p_{ct} - 4((-1)^{j_t})((t-1)/10 + 1 - i_{t-1}); p_{m1} = p_{c1}.$

The cointegration properties of these data, as summarized in Table 1, clearly depend on the length of the sample being examined. The table lists several test statistics both for a 60-period sample and one of 1,200 periods. 1/ First, regardless of length, the commodity price is obviously $I(1)$, but the gradually-adjusting manufactures price is more problematic. With the short sample, p_m is estimated to be $I(2)$ (i.e., the inflation rate is nonstationary), whereas it is $I(1)$ in the long run. Second, the levels of the two series are cointegrated, but the null hypothesis is rejected only when the longer sample is used. 2/ Note that with this simple data generation process, testing for cointegration of levels is exactly equivalent to testing for whether the relative price of commodities is $I(0)$. Third, with the short sample, because the two price indexes appear to be integrated at different levels, it is appropriate to test for cointegration of the commodity price level with the manufactures inflation rate, as was done in Boughton, Branson, and Muttardy (1989). And indeed, one finds (at the bottom of the table) that the two series are so linked.

The relationship between this example and the real-world data is as follows. First, even if there were no trend in the relative price (i.e., if movements in the data were predominantly generated by monetary or offsetting real shocks), there would be no assurance that the relative price would appear to be stationary in the short run; with a longer run of low-frequency data, however, absence of trend and stationarity would be more closely related. Second, if the theoretical model is correct in characterizing the two types of prices as flexible and sticky (and both as being $I(1)$), and if history has been dominated by monetary rather than real shocks, then the manufactures inflation rate should be cointegrated with the commodity price level in the short run or when using high-frequency data, but not in the long run or with lower-frequency data. It should be noted, however, that this condition is not necessary; the empirical finding of a short-run Gibson paradox does not necessarily confirm the model.

III. Trend Analysis

There have been several previous empirical studies of the long-run trends in commodity prices; notable among these have been Prebisch (1950), Singer (1950), Lewis (1952), Spraos (1980), Sapsford (1985a, 1985b), and Grilli and Yang (1988); there also have been a number of quite recent papers using the Grilli-Yang data. Most but not all of these studies have

1/ The data are deterministic, so the test statistics are illustrative. A more substantive analysis would require a monte carlo study incorporating less regular disturbances.

2/ The Durbin-Watson statistic is misleading with the long sample. Because it is based on one-period changes, it shows just as much nonstationarity in the long sample as in the short one. If the frequency of the data were lengthened as well, then the Durbin-Watson statistic would rise rapidly toward 2.

Figure 2. Stylized Behavior of Commodities and Manufactures Prices

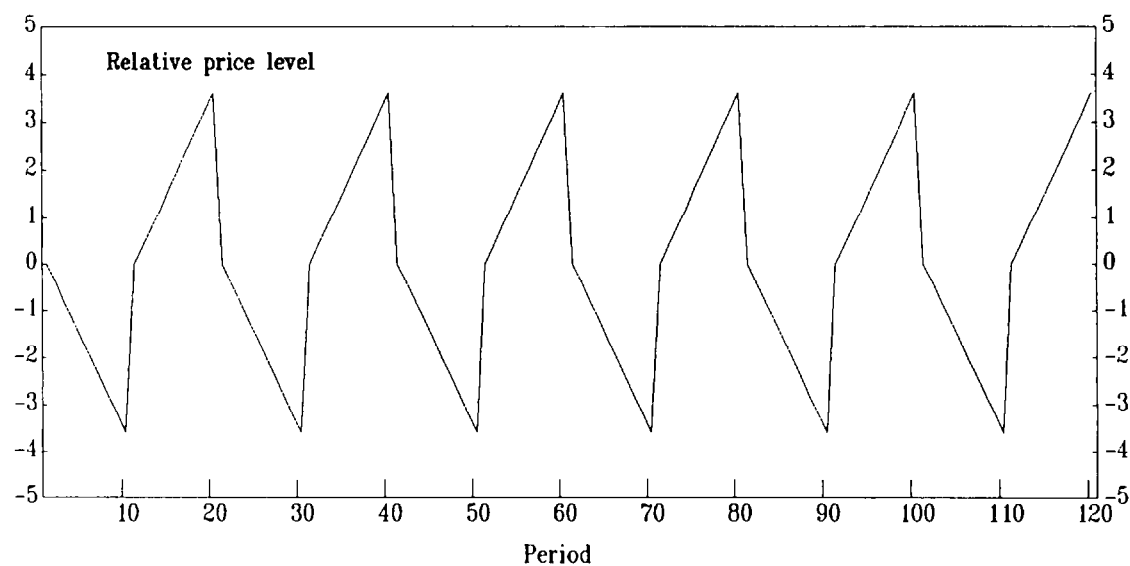
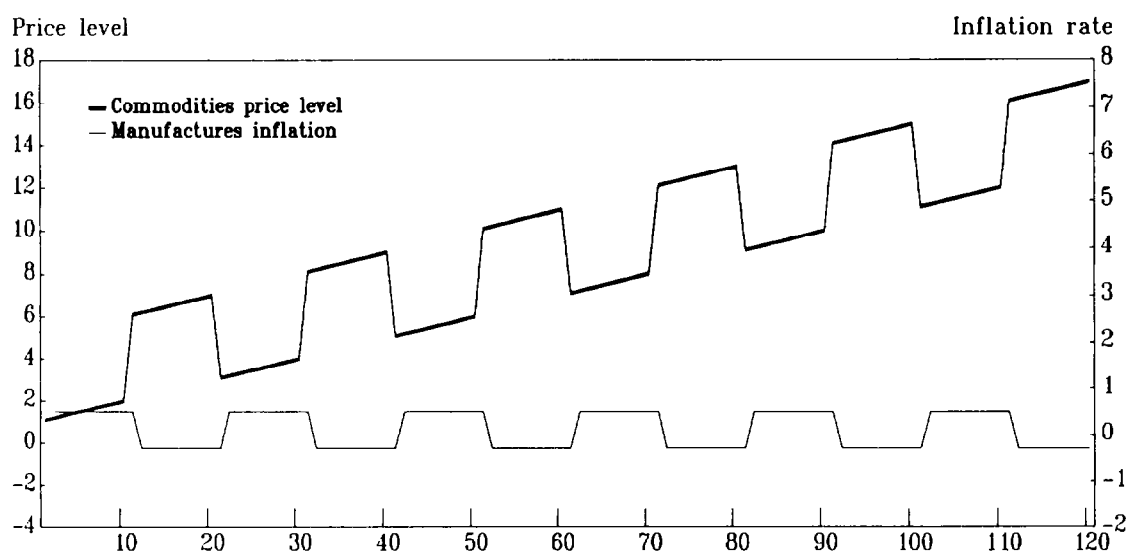
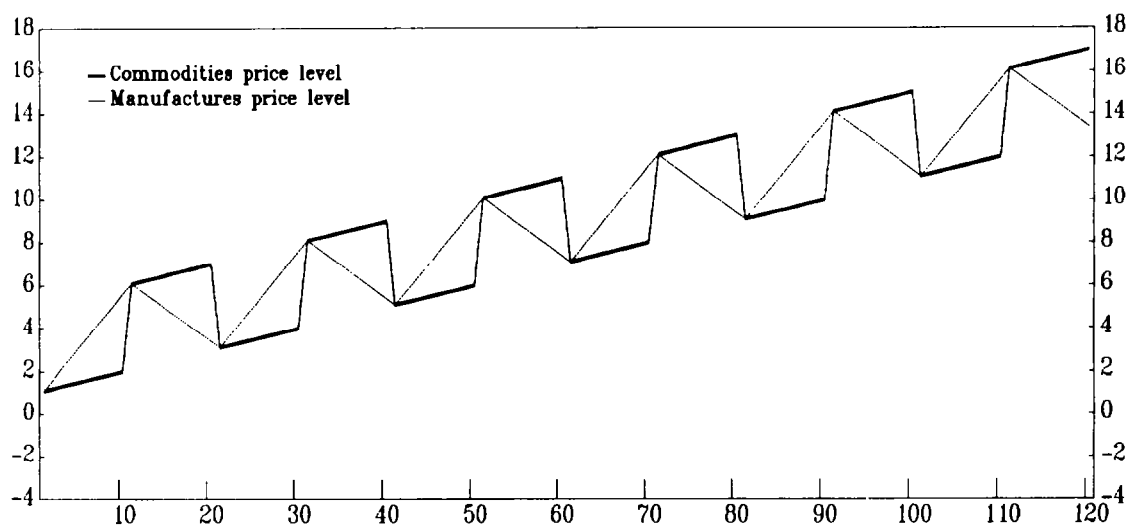


Table 1. Cointegration Tests on Constructed Data 1/

	60 periods		1,200 periods	
First-order integration	<u>P_c</u>	<u>P_m</u>	<u>P_c</u>	<u>P_m</u>
Durbin-Watson <u>2/</u>	2.01**	0.34	2.00**	0.55**
t-statistic <u>3/</u>	-7.42**	-2.35	-34.03**	-13.28**
Cointegration of levels:				
Durbin-Watson <u>4/</u>	0.27		0.31	
t-statistic <u>5/</u>	-1.70		-9.95**	
Cointegration of p _c				
with Δp _m : <u>6/</u>				
Durbin-Watson	0.86**			
t-statistic	-3.63*			

1/ p_C and p_m are, respectively, the constructed data on commodity and manufactures prices. * and ** indicate rejection of the null hypothesis with 95 and 99 percent confidence, respectively.

2/ Sargan-Bhargava test on the null hypothesis that the index is I(2) or higher. If the Durbin-Watson statistic for the regression $\Delta p = \text{constant} + \epsilon > DW_u$, the null hypothesis is rejected. For 100 observations (the largest figure given in Sargan and Bhargava (1983)), $DW_u \approx 0.40$ at the 95 percent confidence level; at 99 percent, 0.54.

3/ Dickey-Fuller test of the t-statistic on β in the regression $\Delta \Delta p = \beta \Delta p_{-1} + \epsilon$. For 100 observations and 95 percent confidence, the critical value for rejecting the null $\beta \geq 0$ is 3.37; at 99 percent, 4.07. See Engle and Granger (1987).

4/ Durbin-Watson statistic for the regression $p_C = \alpha + \beta p_m + \mu$.

5/ T-statistic on β in the regression $\Delta \mu = \beta \mu_{-1} + \epsilon$, where μ is the time series of residuals from the equation just above.

6/ Tests same as above, with p_m replaced by Δp_m . These tests are relevant only where the hypothesis that p_m is I(2) has not been rejected.

concluded that there were substantial negative trends in the relative price of primary commodities, but they have been hampered by limitations in the data, the model, and/or the statistical methodology.

The work by Raul Prebisch and Hans Singer was based on terms of trade data for Great Britain from the 1870s through the 1940s, with implications drawn for the terms of trade between commodities and manufactures on the assumptions that Britain mainly imported the former and exported the latter, and that Britain's experience could be taken as representative of global conditions. Both concluded that there had been a secular tendency for Britain's terms of trade to improve and therefore for the terms of trade of commodities exporters to deteriorate.

Arthur Lewis, working independently around the same time as Prebisch and Singer, was interested primarily in developing consistent global data on trade in commodities and manufactures. He derived time series on prices covering 1870-1913, 1921-38, and 1950, based primarily on data from the League of Nations, and then drew some general conclusions about the relationships between quantities of production and trade, and between trade volumes and relative prices. He found, for example, that the price of food relative to that of manufactures was explained by a regression on indexes of manufacturing output and food production, with positive and negative coefficients, respectively; no trend was included in that regression. The relative price of raw materials, however, was related positively to manufacturing output and negatively to a log-linear trend. 1/

John Spraos (1980) set out to test the validity of the Prebisch-Singer hypothesis using improved data (derived partly from Lewis' work) over a longer time period. Spraos found negative trends in the commodity terms of trade for the pre-World War II period, but not for the full period 1900-1970. However, David Sapsford (1985a, 1985b) noted that the absence of a long-run trend arose because of a sharp upward break in 1950. Specifically, Sapsford (1985b) estimated a significant negative trend of -1.3 percent per annum in the real commodities price over the period 1900-38 (minus the war years 1914-20), with an 82.3 percent (!) upward jump in the 1940s and a -1.8 percent trend thereafter. Spraos, however, replied to Sapsford that reliance on a break after World War II leaves the underlying question of the long-term trend unresolved, since the source of that break is unclear and any number of such breaks could appear.

Grilli and Yang (1988) developed an annual data base covering 1900-86, using price data for 24 internationally traded nonfuel primary commodities. They then regressed the logarithm of the ratio of that index to unit values of manufactures exports on a linear time trend. They found an apparently

1/ Lewis' conclusion is suspect, however, because there was a very strong positive trend component in his manufacturing output data, which was the other argument in the equation determining the level of the price of raw materials. It is thus difficult to disentangle the trend from the influence of output.

stable negative trend of about 0.6 percent per annum. Von Hagen (1989), however, criticized that conclusion, on the grounds that Grilli and Yang did not allow for heteroscedasticity and that the trend model that they estimated is dominated by an error-correction model. Von Hagen found--using the Grilli-Yang data--that commodity and manufactures prices were cointegrated and that there was no significant trend in the relative price. Cuddington and Urzúa (1989) also concluded that there was no negative trend in the Grilli-Yang data, once allowance was made for a structural break in 1920. In contrast, Helg (1990) found a negative trend in the post-1920 period, and Ardeni and Wright (1990) found a negative trend over the full period.

The point that must be emphasized regarding the possibility of a trend is that it cannot be analyzed independently of a model of price behavior. Time series analysis provides an initial fix on the issues, but--as evidenced by the conflicting conclusions in the recent literature--it cannot resolve them. Partly this difficulty arises because of the limitations of the data that have been available. Figure 3, which shows both the official IMF data (a slightly modified and updated version of the data used by Sapsford) and the Grilli-Yang data, illustrates this difficulty. Owing to the two large gaps in the IMF data, the relatively short span of the Grilli-Yang data, and the sudden large movements in each series, the statistical analysis of trends is inevitably subject to wide margins of error. One could either confirm or reject a negative trend in these data, depending on the methodology. But even when the data are interpolated and extrapolated back to 1854, as has been done for this study, time series analysis yields only limited insights.

To illustrate further, in the top panel of Figure 4, a log-linear trend in the relative price of primary commodities has been fitted by OLS for the full period, 1854-1990. The slope is -0.31 percent per annum, and the fitted cumulative decline is 35 percent. There are obvious problems of serial correlation and heteroscedasticity in this fitted trend, and it is not clear whether the observed decline has resulted from a deterministic trend or a random walk (i.e., a unit root). ^{1/}

In the bottom panel of Figure 4, allowance has been made for an apparently significant break in the trend and intercept at 1950; see Sapsford (1985a, 1985b). The expanded fit (with an AR(1) residual) suggests that the downward trend was 0.35 percent through 1949 and 1.00 percent thereafter, with a cumulative decline of 48 percent. With that equation, the White test (using the trend and the squared trend as regressors) is unable to reject unconditional homoscedasticity at the 5 percent level. There is, however, very severe excess kurtosis, and the equation is certainly not an adequate explanation of changes in the relative price. There may

^{1/} The regression is specified as $p_r = \alpha + \beta t/100 + \mu$. With OLS estimation, $\beta = -0.311$ with t-statistic 10.7. When the equation is refit with an AR(2) residual, $\beta = -0.334$ with t-statistic 3.70; $\epsilon = \mu - 1.007\mu_1 - 0.188\mu_2$, with t-statistics 11.48 and 2.13, respectively.

be a deterministic trend, but the estimated trend could also be serving as a crude proxy for other determinants.

Another striking feature of the data is the increase over time in the variance of the commodity price index. The top panel of Figure 5 displays the 15-year moving average of the standard deviations of the annual first differences in the logarithms of the two price indexes. Up to World War I, both indexes show relatively little variation. The manufactures price index then becomes sharply more variable, but it settles down again; by the 1970s and 1980s it is only slightly more variable than in the early years of the sample (see also Table 2). The commodities index, in contrast, becomes more variable in the 1920s and 1930s, gradually settles down through the 1960s, and then becomes quite volatile again. Consequently, there is a secular rise in the relative volatility of commodity prices, 1/ until the standard deviation of changes in commodity prices becomes roughly twice as high as that of manufactures prices in the 1970s and 1980s (Table 2 and bottom panel of Figure 5). 2/

Table 2. Standard Deviations of Changes in Commodity and Manufactures Prices

	(In percent)			
	1855-1990	1855-1913	1914-1970	1971-1990
Commodities	9.21	4.97	9.76	15.13
Manufactures	6.62	4.56	7.37	7.25
Ratio	1.39	1.09	1.32	2.09

1/ The trend line was estimated over the period 1870-1990 by the equation

$$R_t = 0.709 + 0.861 t/100, \\ (0.088) (0.124) \quad R^2 = 0.27, DW = 1.66$$

where R_t is the ratio of the two 15-year moving standard deviations.

2/ It is possible that the change in the variance arises from changes in the way the data are constructed over time; that problem was identified by Romer (1986) in studying U.S. unemployment data from 1900 through 1980. It may be noted, however, that the data on U.K. traded goods prices developed by Schlote (1952)--which are measured consistently from 1854 through 1933--show a clear jump in relative volatility after the onset of World War I.

Figure 3. Alternative Data Series on Relative Commodity Prices
1870 - 1986

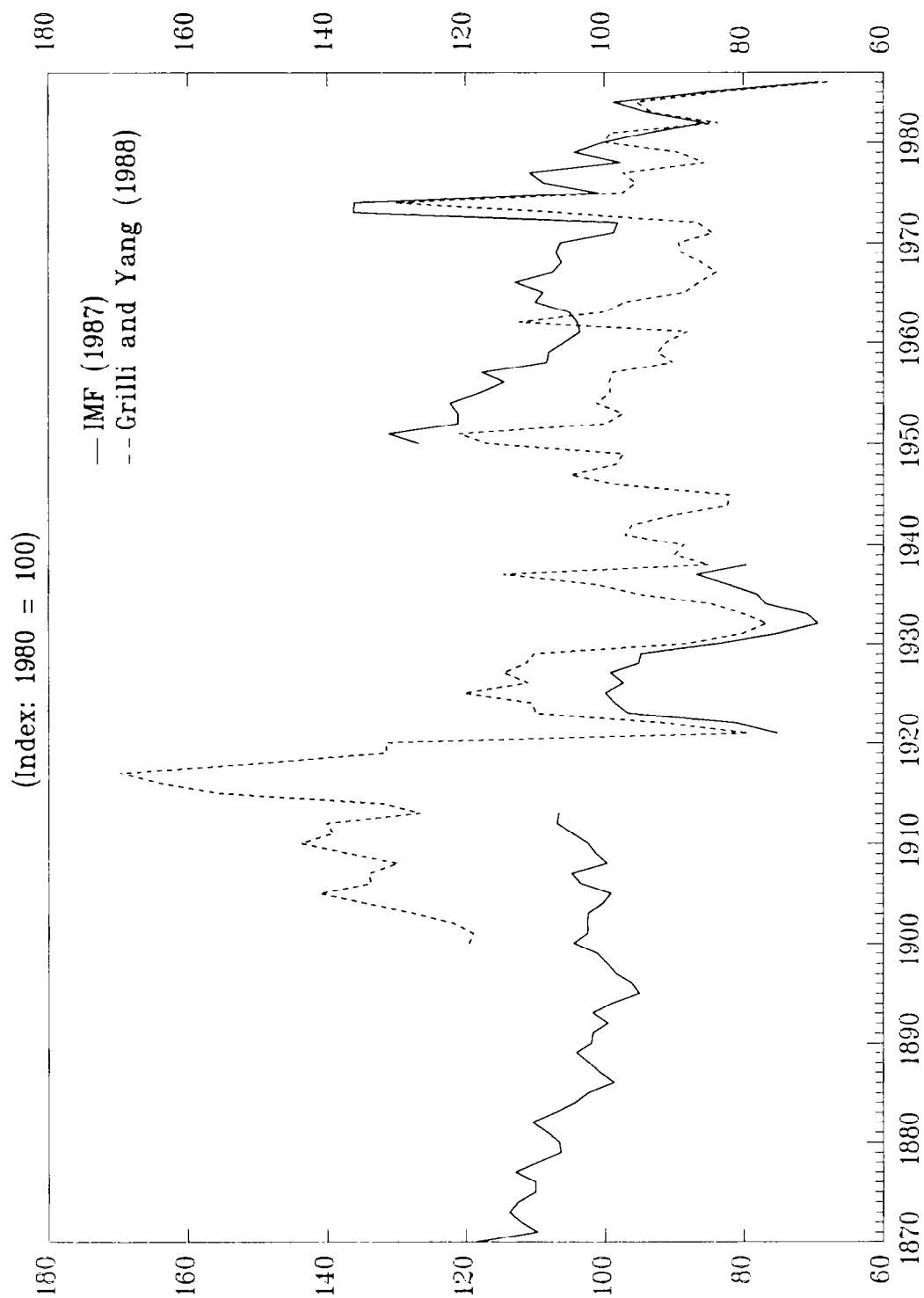


Figure 4. Estimated Trends in the Relative
Price of Commodities, 1854-1990
(1980 = 0)

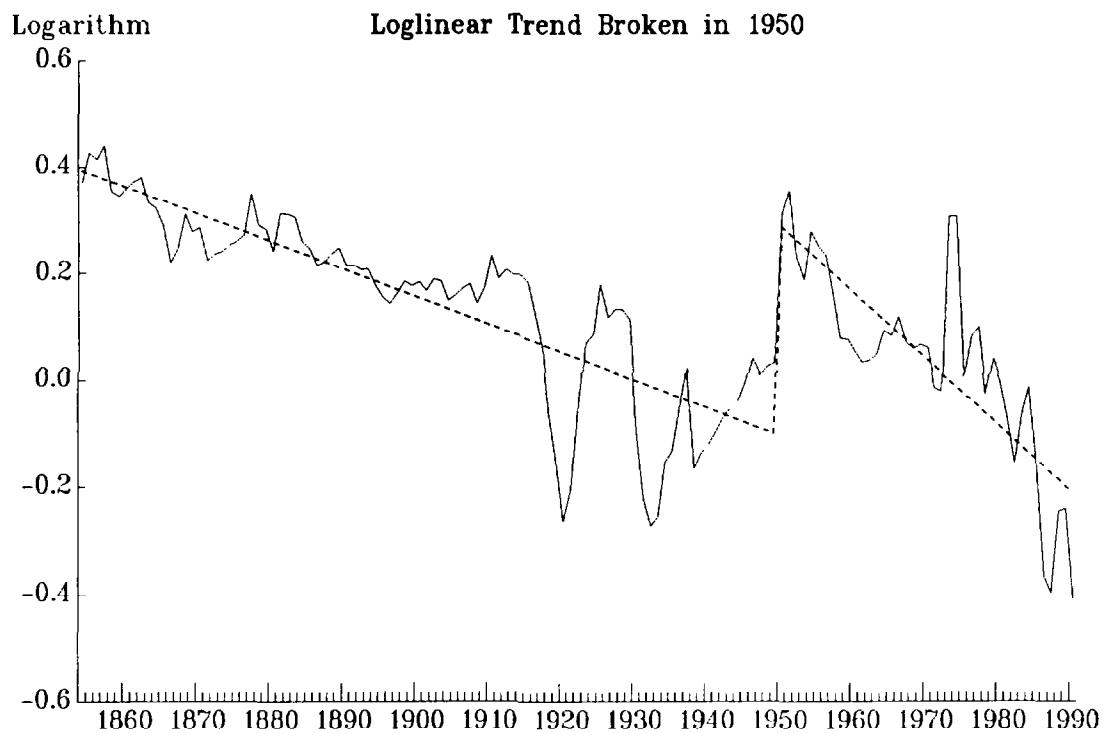
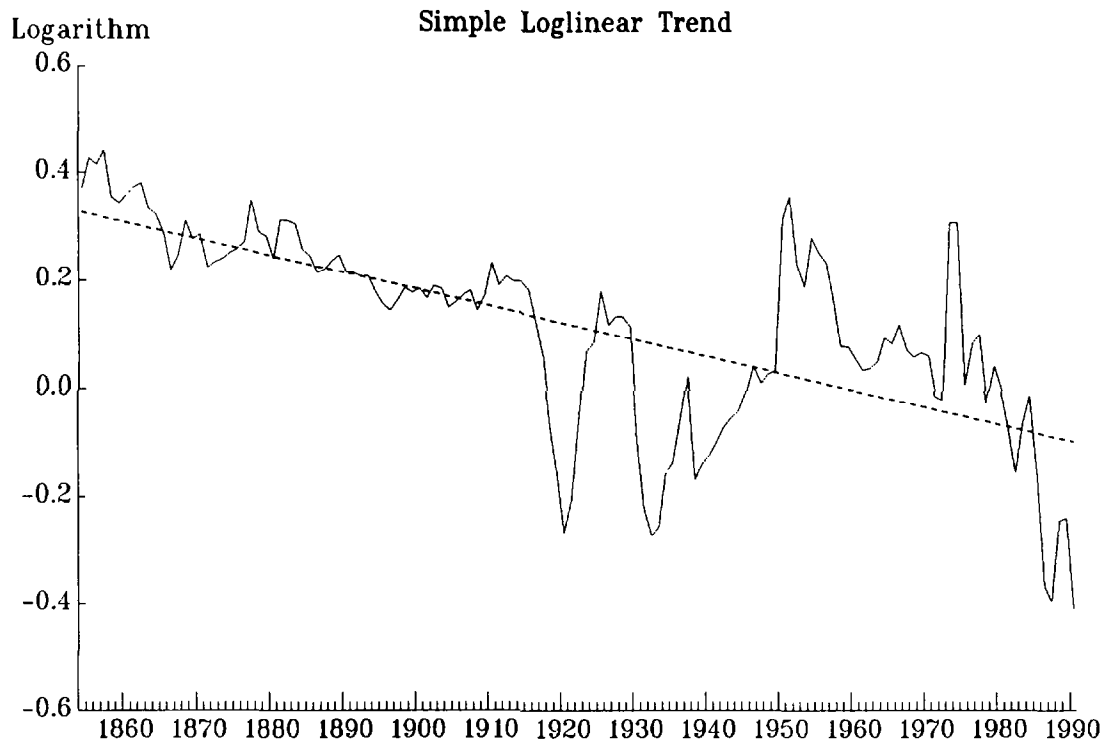
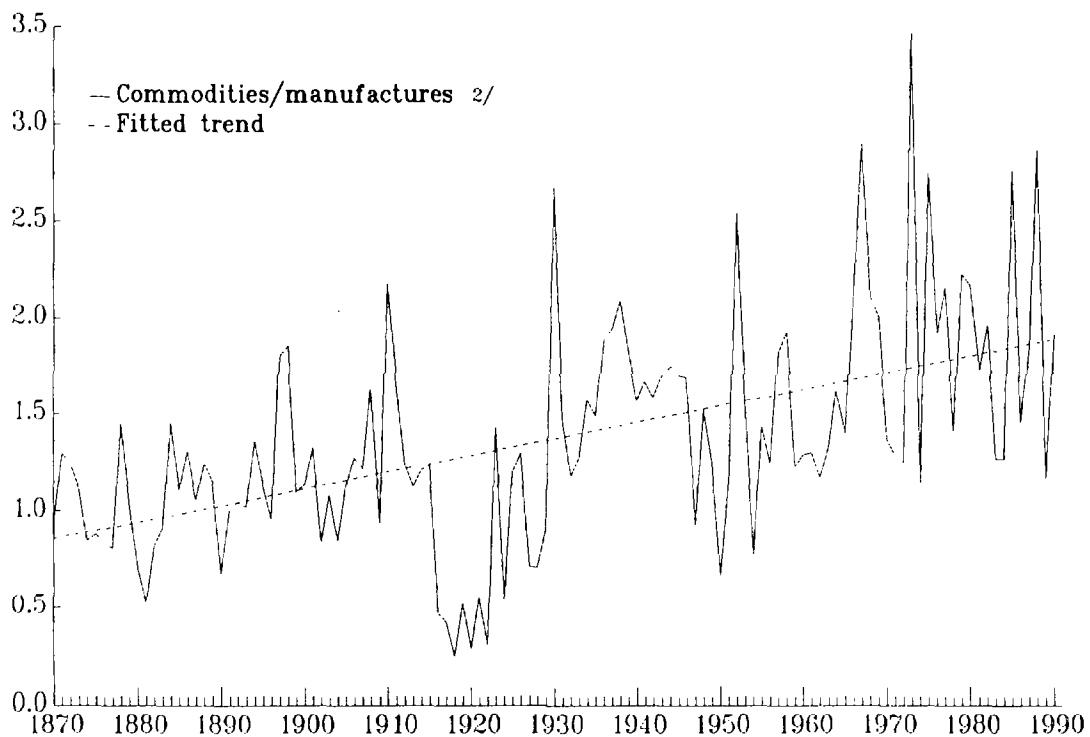
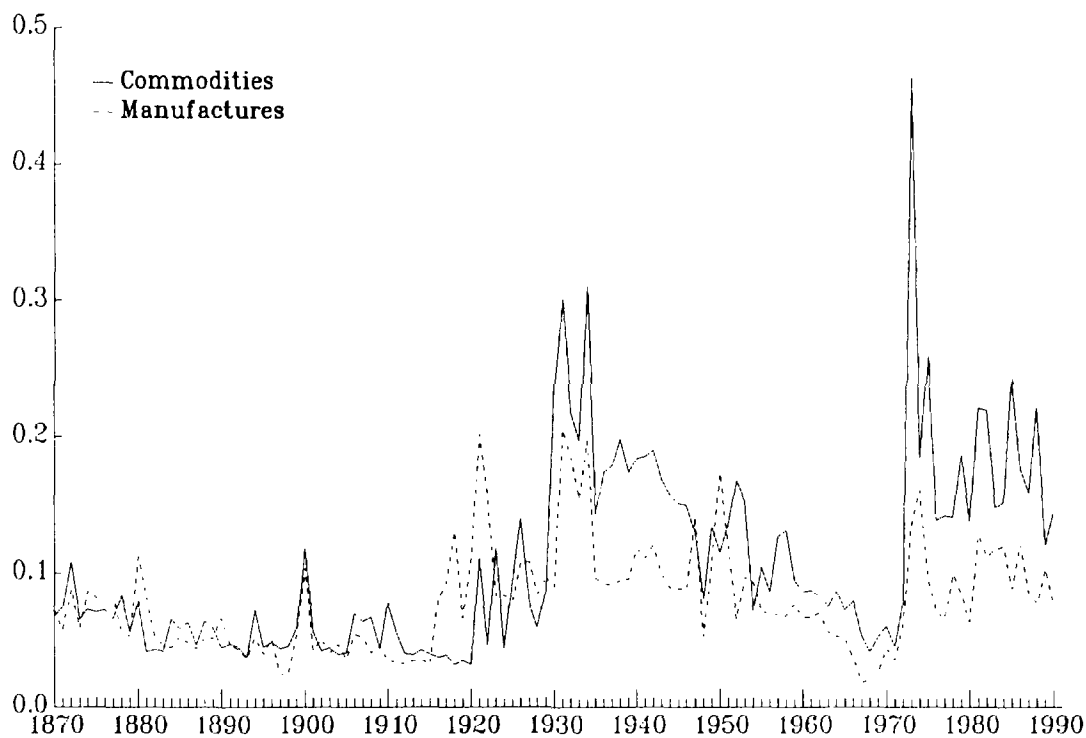


Figure 5. Standard Deviations of Log Changes in Price Indexes ^{1/}
1870 - 1990



1/ Fifteen-year uncentered moving standard deviations of annual first differences in the logarithms of the indexes.
2/ Ratio of the two series shown in the top panel.

IV. Estimating an Error-Correction Model

As a prelude to specifying a more detailed empirical model, it is necessary to examine more closely the time-series properties of the data. 1/ In view of the possibility of deterministic trends in all of the data, appropriate tests can be derived from Perron (1988), as long as one ignores possible breaks in the relationships. For obviously nonstationary series such as the price indexes for commodities and manufactures (p_c and p_m), the issue is whether the indexes are $I(1)$ or $I(2)$. 2/ The null hypothesis is that the once-differenced data conform to a random walk with drift: $\Delta p_t = \gamma + \Delta p_{t-1} + \mu_t$, where μ_t need not be white noise. If that hypothesis is rejected, then the data are deemed to be $I(1)$. The basic regression to be estimated is

$$\Delta^2 p_t = \tilde{\gamma} + \beta t + (\tilde{\alpha}-1)\Delta p_{t-1} + \tilde{\mu}_t, \quad (2)$$

where μ_t will be modeled as being $AR(2)$.

For present purposes, the interesting test statistics are an F-test for $(\beta, \tilde{\alpha}) = (0, 1)$ and t-statistics for $\beta = 0$, $\tilde{\alpha} = 0$, and (if $\beta = 0$ is not rejected) $\tilde{\alpha} = 1$ conditional on $\beta = 0$. The test statistics, listed in Table 3, strongly reject the existence of unit roots in the first differences of either p_c or p_m , and they fail to reject the absence of a deterministic trend in those first differences. Both series will therefore be assumed to be $I(1)$.

The next step is to examine the properties of the relative price of commodities, p_r . Tests on the level of p_r fail to reject the existence of a unit root, while the trend is significant at the 90 percent confidence level. Tests on first differences reject the unit root and fail to reject the absence of trend. Thus the relative price appears at this stage to have a negative deterministic trend and to be $I(1)$. This conclusion is muddled, however, by the possible existence of breaks in the trend. For example, when allowance is made for a break in the intercept and the trend at 1950, the t-statistic on α rises to -4.79, which is significant approximately at

1/ Most of the empirical tests have been carried out in PC-GIVE; see Hendry (1989). The Phillips-Perron statistics were computed by Mark Taylor, using RATS.

2/ Though not shown here, the test statistics discussed below all indicate that neither index is $I(0)$.

Table 3. Stationarity Properties of the Data, 1854 - 1990 1/

	Commodity price index	Manu- factures price index	Relative commodities price		Real output	Nominal Interest Rate
			levels	differences		
F-test for $(\beta, \alpha) = (0, 1)$ <u>2/</u>	36.35***	30.14***	5.05	51.95***	35.99***	30.33***
t-test for $\alpha = 1$ <u>3/</u>	-8.53***	-7.76***	-3.13	-10.20***	-8.48***	-7.79***
t-test for $\beta = 0$ <u>4/</u>	1.37	2.12	-2.48*	-0.38	1.96	1.16
t-test for $\alpha = 1$, conditional on $\beta = 0$ <u>5/</u>	-8.40***	-7.38***	-1.85	-10.25***	-8.17***	-7.70***

1/ See equation (2); the test statistics are given in Perron (1988), Table 1. All tests are in first differences except as noted. For tests in levels, the null hypothesis is that the index is $I(1)$ against the alternative of being $I(0)$; for tests in first differences, the null hypothesis is that the index is $I(2)$, against the alternative of being $I(1)$. Where tests in levels are not shown, the null hypothesis has not been rejected at that level by any test. *, **, and *** indicate rejection of the null hypothesis at the 90, 95, or 99 percent confidence level, respectively. The output data commence in 1869; money and the interest rate in 1867.

2/ The Phillips-Perron statistic $Z(\Phi_3)$; the 90, 95, and 99 percent confidence levels are 5.44, 6.43, and 8.61, respectively (interpolated from Table VI in Dickey and Fuller (1981)).

3/ The Phillips-Perron statistic $Z(t_\alpha)$; the 90, 95, and 99 percent confidence levels are 3.15, 3.45, and 4.02, respectively (interpolated from Table 8.5.2 in Fuller (1976)).

4/ The Phillips-Perron statistic $Z(t_\beta)$; the 90, 95, and 99 percent confidence levels are 2.38, 2.79, and 3.51, respectively (interpolated from Table III in Dickey and Fuller (1981)).

5/ The Phillips-Perron statistic $Z(t_{\alpha*})$; the 90, 95, and 99 percent confidence levels are 2.58, 2.89, and 3.49, respectively (interpolated from the distribution on r_μ in Table 8.5.2 in Fuller (1976)).

the 99 percent level. 1/ Because the choice of breaks is arbitrary, it thus remains ambiguous as to whether p_r is $I(0)$ or $I(1)$.

Even if the level of the relative price has a deterministic trend and is nonstationary, p_c and p_r may be cointegrated: There may be a nonunitary relationship between the two, and there may be other factors that have induced secular movements in the relative price. Bivariate tests for the existence of a cointegrating relationship, however, do not eliminate the ambiguity. The problem again is the existence of possible breaks in the relationship. When the simplest cointegration tests are performed (ignoring the trend and the breaks), the two indexes appear not to be cointegrated. When the trend and a dummy for 1950 onward are included, they are found to be cointegrated, although the long-run relationship is less than unitary. 2/ That is, one would conclude from this test that the relative price of commodities tends to fall over time both because of whatever is causing the negative trend and because of a less than proportional relationship between the two price indexes. This conclusion, however, is unsatisfactory in the absence of further investigation of the underlying relationships.

1/ The full regression, estimated with a first-order autoregressive residual, is

$$\Delta p_r = - 0.700 p_{r-1} + 0.271 - 0.352 \text{ trend} + 0.295 \text{ dum50} \\ (0.146) \quad (0.064) \quad (0.087) \quad (0.062) \\ - 0.650 \text{ dt50}, \quad \rho = 0.584 \\ (0.216) \quad (0.136)$$

where trend = the year (1855, ..., 1990) /100; dum50 = 1 from 1950 onward; dt50 = dum50*trend; and ρ is the autocorrelation coefficient (estimated by Gauss-Newton iteration). For a description and rationale for this test, and for the critical values, see Rappaport and Reichlin (1989).

2/ The basic cointegrating equation, using a second-order autoregressive distribution lag, is

$$p_c = 0.54 + 0.85 p_m + \mu_t \\ (0.21) \quad (0.07)$$

In this form, μ_t has a Durbin-Watson statistic of 0.23 (insignificantly different from zero at the 95 percent level by the Sargan-Bhargava test), and the regression $\Delta \mu_t = - \beta \mu_{t-1}$ yields $t_\beta = -2.87$, which is insignificantly different from zero by the Dickey-Fuller test. The solution of the expanded model is

$$p_c = 0.77 + 0.84 p_m - 0.44 \text{ trend} + 0.40 \text{ dum50} + \mu_t \\ (0.13) \quad (0.06) \quad (0.09) \quad (0.07)$$

The Durbin-Watson statistic now is 0.51, and $t_\beta = -4.27$, both of which imply rejection of the unit root hypothesis. The residuals of both equations, however, display severe kurtosis and first- and second-order autocorrelation.

The final step prior to estimating a model such as equation (1) is to test for the existence of cointegrating vectors that might link all of the included variables. That is, if the forcing variables--real output and nominal interest rates--are also $I(1)$, the levels of those variables may also be related to the levels of commodity prices, and their omission may be a cause of some of the strange results discussed above. As shown in the last two columns of Table 3, each of the forcing variables is estimated to be $I(1)$. 1/

Initial investigation of the full model involved estimating an autoregressive distributed lag on all of the variables that appear in equation (1). The steady state solution of that equation revealed that there are significant long-run (cointegrating) linkages. 2/ As a further check on the existence of a cointegrating vector, a VAR was computed and subjected to the tests developed in Johansen (1989) and Johansen and Juselius (1990). Those tests confirmed the existence of a single cointegrating vector. 3/

Further inspection of the residuals from the general ADL revealed that there were not one but several shifts in the relationships, most of which appear to reflect major disturbances to the supply or demand for primary commodities (top panel of Figure 6). As expected in view of the sharp secular increase in the relative variance of commodity prices, the variance of this proxy for market conditions also rises over time, culminating in the very large movements of the late 1980s. When these effects are estimated as

1/ See Appendix I for a description of the data. As with the two price indexes, tests were also conducted to see if each series was $I(0)$; all of those tests failed to reject the existence of a unit root in the levels of the data.

2/ For example, a 4th-order ADL yielded this long-run result:

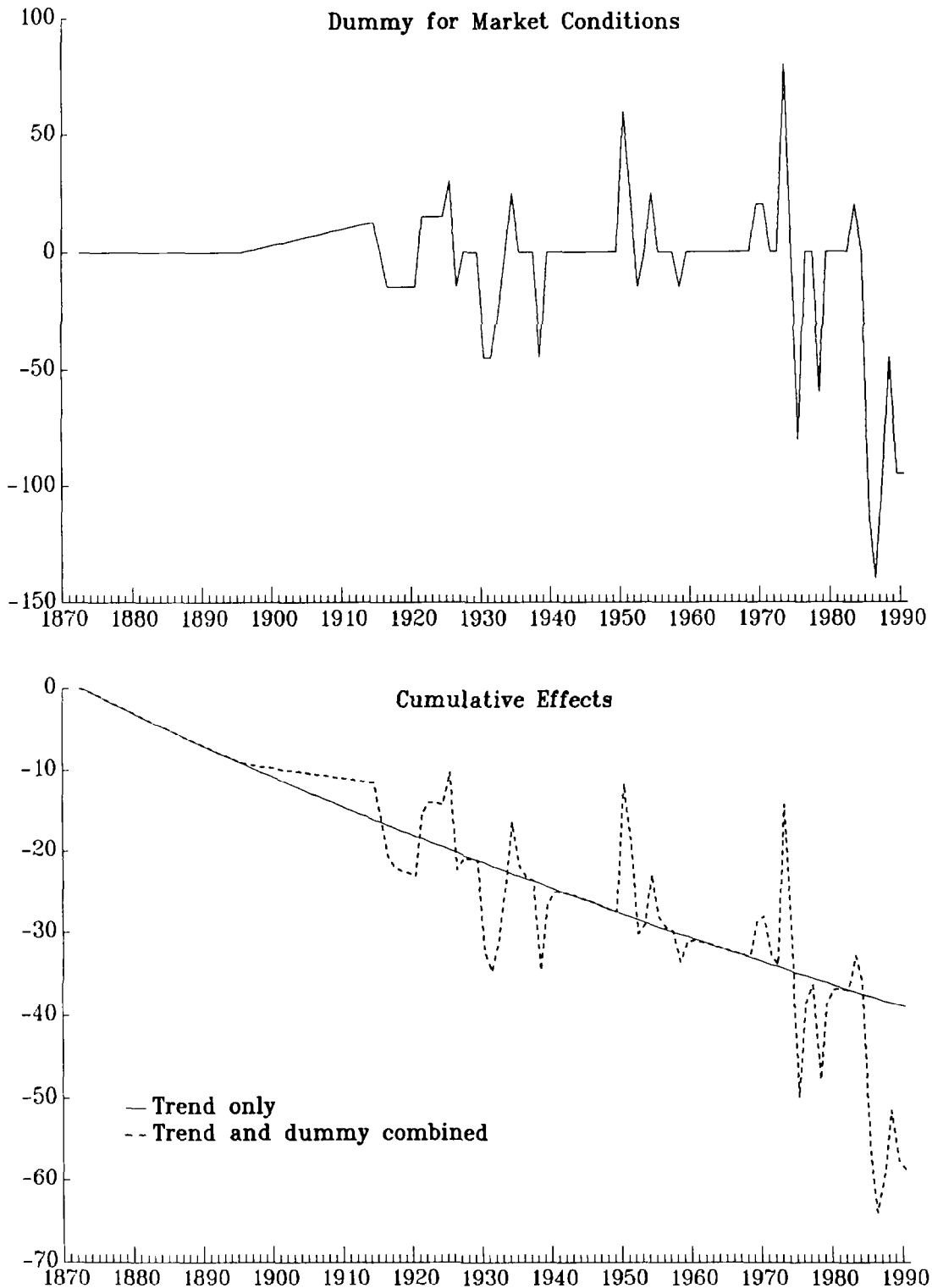
$$p_c = 0.31 + 0.87 p_m - 5.00 i + 0.32 y - 0.015 \text{ trend},$$

$$(0.26) \quad (0.24) \quad (2.90) \quad (0.16) \quad (0.006)$$

with standard errors in parentheses; the sample period was 1854-1990. The residuals from this long-run equation have a Durbin-Watson statistic of 0.39, which is high enough to reject the unit root hypothesis at the 95 percent level. The ADL displays significant skewness and excess kurtosis, but no significant autocorrelation or heteroscedasticity. Note that the coefficient on p_m is insignificantly different from unity. The negative sign on the coefficient on the interest rate is inconsistent with the model.

3/ With no trend in the system and five lags on each variable, the trace statistic (see Table A1 in Johansen and Juselius (1990)) was 56.67, compared with a 99 percent confidence level of 53.79 for rejecting the null hypothesis of no cointegrating vectors; the maximal eigenvalue statistic was 31.72, compared with a 95 percent level of 27.17 and a 99 percent level of 31.94. With a trend included, the corresponding statistics were 66.93 and 41.02, both of which exceed the 99 percent threshold. In each case, the hypothesis of no more than a single cointegrating vector was not rejected at the 95 percent level.

Figure 6. Estimated Trend and Dummy Effects, 1872-1990



part of the error-correction model, the major shocks to market conditions are found to be as follows: 1/

- a gradual upward shift in p_r of about 5 percent during the period of strong growth between the discovery of gold in the Yukon, South Africa, and Australia in the 1890s, and the onset of World War I;
- a downward shift of about 6 percent while agricultural prices were controlled during World War I, followed by an upward swing of 12 percent in the postwar boom of the early 1920s;
- a downward shift of more than 16 percent at the onset of the depression of the 1930s; followed by an upward swing of 8 percent by 1934, at least partly in response to the implementation of price supports and production controls on agricultural products in the United States and to the initial success of international commodities agreements, such as those on rubber, tea, and tin;
- a downward shift of 14 percent in 1938 reflecting both bumper crops and declining aggregate demand in industrial countries;
- an upward shift of 22 percent in 1950 as the Korean war inflation was felt primarily in the largely uncontrolled (and even supported) agriculture sectors rather than in the more pervasively controlled manufacturing sectors, followed by a downward shift of just under 9 percent in 1954;
- an upward shift of more than 30 percent in 1973, when the Soviet grain purchases of the preceding year touched off an unprecedented increase in demand for grains and other agricultural commodities, followed by negative swings of 23 percent in 1975 and 18 percent in 1978, periods when generally good crops meant that commodities did not match the rapid inflation in manufactures prices; and
- a massive decline starting in 1985 that appears to reflect a wide variety of influences on the supply of and demand for commodities, as well as general macroeconomic developments (see below); the unexplained portion of the decline reached 42 percent in 1986 and was still more than 30 percent in 1990.

In addition to these major swings, these estimates reveal a steady log-linear negative trend in p_r of 0.35 percent per annum. The cumulative effect of this trend over the full estimation period 1872-1990 was close to 40 percent, and the combined effect of the trend and the dummy variable for

1/ For historical discussions of developments affecting commodity markets, see Commodity Research Bureau (1939 et seq.).

the 1985-90 period was around 60 percent (bottom panel of Figure 5). 1/ In other words, primary commodity prices in 1990 are estimated to be about 40 percent of the level they would have been in the absence of these trend and dummy effects, given the history of developments in manufactures prices and other macroeconomic variables in industrial countries.

The final estimate of the error-correction model is as follows: 2/

$$\begin{aligned} \Delta p_r = & - 0.086 - 0.203 \text{ ECM} - 0.194 \Delta_3 \Delta p_m \\ & (0.003) \quad (0.006) \quad (0.025) \\ & - 7.837 \Delta_2 i + 4.581 \Delta_2 \Delta i_{-2} + 0.132 \Delta_2 \Delta y \\ & (0.436) \quad (0.350) \quad (0.020) \end{aligned} \quad (3)$$

$$R^2 = 0.94, \quad \sigma = 2.08\%, \quad \text{sample} = 1872 - 1990$$

where $\text{EC} = p_{r-1} - 0.389 y_{-1} - 0.017 \text{ detrend}$,

and detrend is the negative of the trend, plus the market-conditions dummy (top panel of Figure 5). All of the variables in this equation are $I(0)$; thus the relative price of commodities is stationary, conditional on output and the deterministic trend.

In some but not all respects, equation (3) conforms to the model as characterized in equation (1). The equilibrium relationship with output is as expected, but the interest rate has only a dynamic effect: a rise in the rate temporarily reduces the relative price of commodities, but it does not permanently raise it as hypothesized. About 20 percent of any disturbance from the equilibrium relationship is estimated to be closed each year. The detrend term could be interpreted as a proxy for b (the net real return to holding commodities), but it could also represent shifts or trends in any of

1/ The cumulative effects take account of the lagged effects working through the error-correction term. Ignoring all other variables, the model is estimated in the form $\Delta p = -(1-\alpha)p_{-1} + \beta_1 D + \beta_2 \text{trend}$, where D is the market-conditions dummy. If $D = 1$ in periods $1, \dots, N$ and 0 thereafter, then its cumulative impact at time T ($T \geq N$) is $\beta_1 \sum_{i=1}^N \alpha^{T-i}$. The cumulative impact of the trend is $\beta_2 \sum_{i=0}^{T-1} \alpha^i$. Percentage changes are calculated in relation to the base period; that is, they are exponents of the log changes.

2/ Standard errors, in parentheses, are corrected for heteroscedasticity. The residuals from equation (3) display some significant negative first-order serial correlation, but they are otherwise normal and homoscedastic. Estimation by recursive least squares reveals no significant instabilities as measured by N -step Chow tests. The serial correlation is induced by the construction of the combined dummy variable; if that term is somewhat simplified, the serial correlation is eliminated but the residuals become heteroscedastic and the recursive Chow tests indicate instabilities.

the structural relationships governing relative price movements, and to some extent it could have resulted from the splicing together of data that are not fully consistent over time.

V. Conclusions

This paper has confirmed the Prebisch-Singer hypothesis that the price of primary commodities tends to decline secularly relative to the prices of manufactured goods. The empirical tests have extended earlier work in several directions: (a) by interpolating and extrapolating data so as to construct annual aggregate time series on commodity and manufactures prices from 1854 through 1990; (b) by estimating the trend coefficient in the context of an error-correction model that explains relative price movements in relation to developments in real output and interest rates in industrial countries; (c) by taking account of both the trend and a number of major shifts in market conditions while investigating the time series properties of the data; and (d) by relating the long-run behavior of the data to shorter-run properties that had been revealed in earlier work.

Although there is a deterministic secular trend, whereby the relative price of primary commodities has declined by around 1/3 of 1 percent per annum, the relative price is otherwise stationary. There is a strong and significant tendency for it to revert to its trend-adjusted mean level. Output growth in industrial countries tends to lower the price of manufactures relative to those of primary commodities, while increases in monetary growth rates or in nominal interest rates tend to raise the price of manufactures relative to commodities. Since 1985, however, the relative price of commodities has plummeted to an unprecedented level, and that decline cannot be explained either by the longer-run trend or by the other variables in the model.

One excluded factor that may help to explain the sharp decline of the past few years is the exchange rate. The U.S. dollar price of manufactures rose by some 35 percent from 1984 to 1990, even though inflation generally was low and stable in industrial countries. The discrepancy largely reflects the inflationary effect of the dollar's depreciation on the dollar price of goods manufactured and marketed in countries outside the United States. World prices of primary commodities, however, may have been less affected, perhaps because these commodities are more widely traded in more highly competitive markets. (For general discussions of the effect of exchange rate changes on commodity prices, see Ridler and Yandle (1972) and Boughton, Branson, and Muttardy (1989).) A second possible factor is increased competition among suppliers in commodity markets as exporting countries may have attempted to generate additional revenues in the wake of the debt crisis. Third, there have been a number of factors affecting specific commodities--health concerns that have reduced the demand for tropical oils and coffee; environmental concerns that have reduced the demand for phosphates; the weakening of international marketing arrangements, including those for coffee, cocoa, and tin; a series of bumper crops in cereals; sharp increases in metals production in response to

earlier strength in demand, and so forth--that have combined to depress the overall price index for primary commodities. Sorting out these various influences, and assessing the likelihood of their persistence into the 1990s, would require much more detailed study.

In the absence of more information about the causes of the downward trend in the relative price of commodities, it would be premature to draw normative conclusions from it. It is possible, for example, that there has been a secular underestimation of quality improvements in manufactured goods such that a properly adjusted index of manufactured goods would have risen by a third of a percent less per annum, completely offsetting the estimated trend. Perhaps more disturbing is the long-term increase in the variability of commodity prices. The abrupt price swings that occurred in the 1970s and the sustained relative decline in the second half of the 1980s have had adverse consequences on both sides of the market. Even if the depressed conditions of recent years are reversed in the 1990s, both importing and exporting countries could continue to face uncertain terms of trade until market forces are more well understood.

Finally, it may be noted that these results lend some support for the idea of using movements in primary commodity prices as an indicator of global inflationary developments. It has been shown here that, with proper allowance for the influence of trends and other factors, commodity prices are cointegrated with manufactures prices. The Gibson paradox displayed by short-run data, the level of commodity prices seemingly cointegrated with the general inflation rate, has been shown to vanish in the longer run.

Derivation of the Data

The data have been derived from a number of sources. The starting point for the two price indexes was a pair of data series developed and maintained by the Research Department at the IMF: (a) prices of 34 non-fuel primary commodities that are important in world trade (available monthly from January 1957), and (b) unit values of manufactured goods exported by industrial countries (available monthly from January 1948). Both series are expressed in U.S. dollars and indexed to 1980=100; the data used here are annual averages of monthly data. For a detailed description of the commodity price data, see IMF (1986), Appendix I. The prices of manufactured goods cover SITC categories 5 through 8 and are averaged from OECD country-specific data using 1985 exports of manufactured goods as weights.

Both price series have been extended back to 1900 by splicing the IMF data to unit value indexes from United Nations (1969). The U.N. data for the post- World War II period are derived from comprehensive world trade data, but the coverage becomes progressively narrower as the data reach farther back in time; it is estimated (p. 366 of United Nations (1968)) that the primary product coverage is about 50-55 percent for the period 1900-13. (Further details on the U.N. indexes are provided in United Nations (1968).) Primary commodities ("other goods") comprise SITC sections 0-4 and 9. These data thus include petroleum trade in primary commodities, which reduces the correspondence in coverage with the IMF data. The importance of this difference, however, is much smaller for the period in question than it would be for the more recent period. Schlote (1952, p. 58) notes, for example, that "mineral oils" accounted for just 3.1 percent of British imports of raw materials in the period 1909-13, and 10.8 percent for 1927-29.

The U.N. data omit the years 1914-20, 1939-47, and 1949. Earlier studies using these data have left these gaps in place, but for this study the timing of changes during the missing years has been estimated on the basis of partial data from national sources. To interpolate the World War I period, I have used British trade data from Schlote (1952, Table 26). Schlote's data cover three categories: (a) foodstuffs and livestock, (b) raw materials and semi-manufactured goods, and (c) manufactured goods. Schlote presents a price index covering total trade (including re-exports) in manufactures; to get an index for primary commodities, I have aggregated the indexes for imports of (a) and (b) (since exports of primary commodities were a very small component of British overseas trade in that period). Lewis (1952, p. 136) estimates that for the period 1913-29, the ratio of raw materials to food in world trade was 9:6; I therefore have assigned a weight of 60 percent to raw materials for this aggregation. These indexes have been converted from pounds sterling to U.S. dollars using annual average exchange rates, from Board of Governors of the Federal Reserve System (1943). Next, I have divided the total change in the U.N. indexes for world trade for 1913-21 by the total change in the British indexes, multiplied the year-to-year percentage changes in the British indexes by that factor, and used the resulting percentage changes to interpolate the U.N. data.

The World War II period and its aftermath (1939-49) have been interpolated using annual data from the United States as well as the United Kingdom. The U.S. data, from the Department of Commerce (1947, 1951), include price indexes for imports and exports of (a) crude materials, (b) crude foodstuffs, (c) manufactured foodstuffs, (d) semimanufactures, and (e) finished manufactures; and percentage distributions in trade, which I used to derive weights. The first two categories (a and b) have been aggregated to get primary commodities, and the remaining three to get manufactured goods.

The United Kingdom data for the 1939-49 period come from the Central Statistical Office (1952), with a few missing observations filled in from the Records and Statistics supplement to the Economist for September 6, 1947. The categories in this case are (a) food, drink, and tobacco, (b) raw materials and articles mainly unmanufactured, and (c) articles wholly or mainly manufactured. Import prices for (a) and (b) have been averaged to derive a price series for primary commodities, and total import and export prices for (c) have been averaged to derive a series for manufactures. In each case, average relative value figures for the period 1938-50 (from the same source) have been used as weights, and the indexes have been converted to U.S. dollars at annual average exchange rates.

The next step in deriving data for this period is to average the price movements for the United States and the United Kingdom. In order to obtain comparable data, and to avoid distortions arising from the disruption of trade during the war, I have used relative trade data from the League of Nations (1945, pp. 157, 158, and 166) covering the period 1930-38. Those data show that the United States accounted for 49 percent of the two countries' total trade in primary commodities, and 44 percent of total trade in manufactures. The two countries together accounted for around 25 percent of total world trade in both commodities and manufactures. Percentage changes in the weighted average price series have been used to interpolate the world data, as in the World War I period discussed above.

Data for 1854-99 have been derived from Schlote (1952), generally following the choices selected by Lewis (1952). These data cover British traded goods and are expressed in U.S. gold dollars at the 1913 parity. Raw materials prices ("raw materials and semi-manufactured goods") are an average of British import and export prices (excluding re-exports), weighted by trade shares. Food prices ("foodstuffs and livestock") are just for imports; Schlote notes that British food exports in that period consisted primarily of processed goods. The two are weighted by trade shares to obtain a price index for primary commodities. Prices of manufactured goods are for total trade (again excluding re-exports); Lewis used an unweighted average so as not to give too much emphasis to exports, but the weighted average seems more appropriate for the present study. All of these data have been derived through 1913; the period of overlap with the United Nations data (1900-13) has been used to adjust percentage changes so that the two series could be smoothly spliced.

The remaining data--real output and the nominal interest rate--relate only to the United States. Data are available for the United Kingdom as well, but aggregation across the two countries over such a long time period would raise difficult measurement issues. Real output is defined as real net national product. Data from 1939 on are from the Council of Economic Advisors (1991). That series has been extended back to 1869 using annual percentage changes in data from Table 4.8 in Friedman and Schwartz (1982). The interest rate is the yield on Aaa corporate bonds, with data for 1919-41 from Board of Governors (1943) and data for 1942-90 from Council of Economic Advisors (1991). That series has been extended back to 1867 using data from column 8 in Friedman and Schwartz' Table 4.8. There was a fairly stable difference in the two series during the period 1919-39, which averaged 36 basis points; that amount was added to each observation in the Friedman-Schwartz data. The coefficients reported in this paper are semi-elasticities; that is, they are scaled for the interest rate measured as a decimal rather than in percent.

Derivation of the Reduced Form

The relevant portion of the flexprice model developed in Boughton and Branson (1991) comprises the following three equations, characterizing a two-sector economy producing and consuming both a primary commodity and a manufactured good, with the commodity price determined in a flexible market with forward-looking expectations:

arbitrage in the commodity market:

$$(A1) \quad i = \dot{p}_c + b$$

demand for the manufactured good:

$$(A2) \quad d = \delta(p_c - p_m) - \sigma(i - \dot{p}_m)$$

and gradual adjustment of excess demand for the manufactured good:

$$(A3) \quad \dot{p}_m = \pi(d - y_m)$$

where the notation is as in the text, with the addition of y_m (output of the manufactured good); $y = \alpha y_m + (1-\alpha)y_c$, where y_c is final output of primary commodities. If commodities are used only as inputs, $\alpha = 1$. Equation (1) in the text is derived by combining equations (A1) through (A3) and solving for the rate of change in the relative price ($p_r \equiv p_c - p_m$):

$$(A4) \quad \dot{p}_r = -b - (\eta\delta)p_r + \eta y_m + (\eta\sigma)i$$

where $\eta = \pi/(1-\pi\sigma) > 0$ by assumption (if $\eta \leq 0$, a positive shock to the excess demand for the manufactured good would not raise the price of the good).

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