COMMODITY PRICE VOLATILITY ACROSS EXCHANGE RATE REGIMES

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ABSTRACT

This paper documents a new "stylized fact" regarding the relative price of primary commodities in terms of manufactured goods. Using alternative data sets covering the period from 1880 to 1996, this key relative price among two categories of ***tradable*** goods is shown to exhibit greater volatility under flexible-exchange rate regimes than fixed-exchange rate regimes. Implications of this finding for open-economy macro modeling are briefly discussed.

Key words: commodity prices; price volatility; exchange rate regimes; time series models

JEL classification: F33; E3; C22

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I. Introduction

Empirical studies of business cycles in open economies have sometimes concluded that nominal and real exchange rates are virtually the *only* macroeconomic variables that have experienced dramatic changes in volatility across exchange rate regimes.[1] This paper documents a new "stylized fact" regarding commodity prices using alternative data sets covering the period from 1880 to 1996*: The volatility of real primary commodity prices, defined as nominal commodity prices deflated by the manufacturing unit value index, is higher under flexible-exchange rate regimes than fixed-exchange rate regimes*.[2]

Why should we care about this finding? More than 40 percent of world trade consists of primary commodities. Both long-term trends and short-term fluctuations in primary commodity prices have important consequences for the world economy. Concerns over possibly adverse long-term trends in commodity prices have occupied development economists at least since the late 1940s.[3] While conclusions regarding

[1] See Mussa (1986), Baxter and Stockman (1989), Flood and Rose (1995), and Rogers (1995), which focus on the post World War II period. Baxter and Stockman (1989) find that the volatility of real exports and imports increased in the post-1973 period for most of the countries in their sample, but these increases did not depend systematically on the exchange rate regimes chosen by the countries after 1973. Grilli and Kaminsky (1991) consider a longer historical perspective. Using monthly data from 1886 to 1986, which spans several changes in exchange regime, they claim that even real exchange rate behavior is regime-independent. Liang (1998a), however, finds their results are quite sensitive to the choice of econometric technique and which subperiods are included in the analysis.

[2] Appendix V, available from the authors, shows that the volatility of *nominal* commodity prices also increases under flexible exchange regimes.

[3] See Prebisch (1950) and Singer (1950), and more recently Grilli and Yang (1988), Cuddington and Urzúa (1989), Cuddington (1992), Powell (1991), Reinhart and Wickham (1994), among many others.

long-term trends are mixed, there is general agreement that the *volatility* of primary commodity export prices is costly for both LDCs and industrial countries. For example, it generates uncertainty in LDC export revenues, which can have negative effects on investment and growth. Moreover, macroeconomic mismanagement in the face of commodity export booms and busts has been a major problem. Understanding the nature of commodity price volatility is a prerequisite for sensible risk management strategies for multinational corporations as well as policymakers.

From a theoretical standpoint, our findings have implications for the plausibility of various macroeconomic models used in studying alternative exchange rate regimes. As Stockman (1983) highlights, two types of models have been used in the literature on real exchange rate volatility across regime. Sticky-price models (e.g., Dornbusch (1976), Frenkel (1981), and Mussa (1982)) exhibit nonneutrality with respect to the nominal exchange rate regime. Adopting floating exchange rates implies higher nominal and real exchange rate variability. In contrast, a class of equilibrium models (see, e.g., Helpman (1981), Lucas (1982) and Stockman (1980)) satisfies "the nominal exchange regime neutrality proposition," i.e. the time series properties of all real variables are invariant to the choice of exchange regime. Hence, the alleged difference in real exchange rate behavior across exchange regimes has generally been viewed as a consequence of price stickiness. Stockman shows, however, that some equilibrium models, for example those with nontradable as well as tradable goods – not just sticky price models – may also exhibit nonneutralities with respect to the exchange rate regime.

There is little research, however, examining the volatility of the relative price of two *traded* goods across exchange rate regimes. The nominal price of a traded good may increase during floating exchange periods simply because of the mechanical effect of denominating the price in a common currency. If the exchange rate pass-through effect is similar for two different traded goods, however, the *relative* price of any two of these goods should not be expected to exhibit shifts in volatility across nominal exchange rate regime. On the other hand, if the mechanisms through which prices respond to exchange rate shocks are different for goods with different market structures, the relative price of two traded goods may also exhibit regime-dependent volatility.

To examine this contention empirically, this study focuses on the relative prices of primary commodities in terms of manufactured goods. Although both goods are internationally traded, the underlying market structures have long been perceived to be different. While the prices of most non-oil primary commodities are determined in well-organized auction markets, the markets for many manufactured goods are thought to be monopolistically competitive. If regime-dependent behavior is found for relative commodity prices, this can be interpreted as further evidence against theoretical models that exhibit nominal exchange rate regime neutrality, and in support of 'sticky prices' attributable to noncompetitive market structures.

The paper is organized as follows. Section II motivates the study by presenting graphical evidence on commodity price behavior over historical periods characterized by different exchange rate regimes. Section III presents more formal testing of the null hypothesis that there are no changes in volatility across exchange regimes. The bulk of the evidence rejects this null hypothesis. The last section concludes.

II. A Preliminary Look at the Data

1. Description of Data Sets

This study considers three alternative data sets: (1) the annual data set of Grilli and Yang (GY) from 1900 to 1992, and (2) Boughton's data set with annual observations from 1854 to 1990, and (3) the International Monetary Fund's *International Financial Statistics* (IFS) monthly data for the post WWII period. Following the norm in the commodity price literature, this study examines a *real* commodity price index, defined as the ratio of the chosen nominal commodity index deflated by a manufacturing unit value (MUV) index.

GY (1988) constructed a dollar index of 24 internationally traded non-fuel commodity prices from United Nations sources, beginning in 1900 through 1992.[4] They also provide data on a MUV index and a U.S. producer price index for use as alternative deflators.[5] James Boughton (1992) presents a nominal commodity price index and the MUV index based on national data and U.N. indexes. His series span the period 1854-1990 and include 34 non-fuel commodity prices:[6] The IMF's *International Financial Statistics* (IFS) reports monthly commodity price indexes for agricultural raw materials, beverages, fertilizer, food, metals and sugar from January 1957 through May 1996.

[4] The GY index is defined as a base-weighted arithmetic average of 24 non-fuel commodities, using values of world export for 1977-79 as weights. For the periods of 1914-20 and 1939-47, interpolation were used to fill the gaps in the UN data sources. The series have been updated through 1992 by Grilli and Yang since their original study.

[5] Their results regarding the long-run trend of relative commodity prices are similar regardless of the choice of deflator.

[6] Data for the World War II years, 1939-45, are obtained by interpolation using annual data from the United States and the United Kingdom.

Volatility is ideally analyzed as a high-frequency characteristic in the data. Hence, monthly series are, in principle, superior to annual series. The monthly data permit a detailed analysis of short-run volatility patterns using GARCH models, which are typically not present in annual-frequency data. This facilitates a time series analysis of volatility patterns, including the distinction between conditional and unconditional volatility. Although much of the work comparing the volatility of various macroeconomic variables across exchange rate regimes uses monthly or quarterly post-WWII data, this data span has the disadvantage that it contains only one major change in exchange rate regime. The exact date of the regime shift is open to debate.[7] Moreover, the post Bretton Woods period has been associated with a number of major oil shocks, which in turn have impacted other commodity prices. It was also a period characterized by considerable government intervention in international capital as well as primary commodity markets. Thus, it leaves open to question any conclusion that the change in exchange regime rather than these other factors have been the underlying cause of the increase in commodity price volatility.

The advantage of the long time-span annual series is that they contain a number of exchange rate regime shifts, not all of them coincident with major oil price shocks, wars,[8] the prevalence of capital controls, or other potentially causal events. However,

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[7] Not all currencies started floating against the dollar in 1971. The subsequent test results, however, are not very sensitive to the exact dating of regime shifts. Changing the regime shift date to 1972 or 1973, or excluding the period of 1971-73, we still reject the null hypothesis of no volatility shift across exchange regimes.

[8] A referee conjectured that changes in commodity price volatility might be the result of wars, rather than shifts in exchange rate regime. Appendix IV, available on request, tests and rejects this hypothesis.

the limited number of observations during some of the brief exchange regime episodes may leave skeptical readers unconvinced.[9] Nevertheless, if both long-run and short-run data sets produce similar results, our findings on the non-neutrality of nominal exchange regime on real commodity price volatility will be more firmly established.

2. Historical Description of Exchange Rate Regimes

Using Eichengreen (1994) as a guide, three fixed and three flexible exchange rate regimes were identified over the period for which we have commodity price data. (See also Bordo and Schwartz (1996), who date the regimes slightly differently, for a discussion.) There are three fixed exchange rate episodes. The first is the classical gold standard from 1880 to 1913, a period characterized by fixed nominal exchange rates and essentially no capital controls. The second was the interwar years from 1927 to 1931. Finally, there was the Bretton Woods system from 1946-71. During this period, the dollar was pegged to gold while other countries pegged to the dollar. While the IMF promoted current account convertibility, many countries maintained capital controls.

There are three flexible exchange rate episodes during the period studied here. The first episode, which was as close to a free float period as history provides, is from the beginning of the First World War until the mid-1920s. The second floating exchange episode is 1932-38, a period when government intervention in the foreign exchange market was pervasive. On August 1971, the devaluation of the dollar brought the Bretton

[9] For example, there are only five observations in the fixed exchange period of 1927-31 and seven observations in the flexible exchange period of 1932-38.

Woods period to a close. Thus, the post-1972 period is the final floating rate episode considered.

3. Some Visual Evidence on Commodity Price Volatility across Exchange Regimes

Figure 1 plots year-over-year percentage changes in the long-span annual series, with the shaded years corresponding to the fixed exchange rate episodes. Simple visual examinations of the data sets point to striking differences in the real commodity price behavior across exchange rate regimes. It appears that prices were relatively tranquil during the fixed exchange episodes and became much more volatile during the flexible exchange periods.

*** INSERT Fig. 1 and Fig. 2 here ****

Figure 2 plots the percentage changes in the monthly IFS real commodity price indices. For agricultural raw materials, beverages, fertilizer, and food, there is a pronounced increase in volatility under flexible exchange rate regimes. It is less clear whether there have been shifts in volatility across exchange regime for the relative prices of metals and sugar. A simple statistical summary of real commodity price volatility across the six exchange rate regimes is shown in Table 1; it also summarizes the dating of the changes in exchange regime. We now proceed with formal testing of our volatility hypothesis.

III. <u>Econometric Evidence on Volatility Shifts Across Exchange Rate Regimes</u>

Most empirical studies of regime-varying volatility used data on exchange rates for the post-WWII period (see fn.1). An exception is the Grilli and Kaminsky's (1991) study In this paper, we first consider our two annual data sets to determine whether the regime-dependent behavior of relative commodity prices is present over our long spans of data containing several changes in exchange regime. We next turn to a more detailed analysis on the monthly data for the post-war period, using a regime-dependent GARCH specification.

As a preliminary step, Phillips-Perron tests for the presence of unit roots are carried out for each price series. For the GY and Boughton series, the unit root hypothesis is rejected. The null of unit root is rejected at the 5% significance level for the monthly series of agricultural raw materials, metals and sugar; for the other monthly series, the unit root hypothesis is not rejected. For brevity, these results are not reported here. [10]

1. The Long-Run Annual Data

The long-span time series data permit the investigation of commodity price volatility across six different exchange regimes.[11] Before studying the volatility across different exchange regimes, we test the simple hypothesis that the volatility is constant across the fixed exchange rate periods for the long-run data sets. A similar test is carried out for the flexible rate periods. Suppose that $\Delta log(y_t)$ (y_t being the real commodity price) is normally distributed, we can then use analysis of variance (or ANOVA) methods to test the null hypothesis of equal variance across all fixed (flexible) exchange rate episodes. The test results in the upper panel of Table 3 indicate that the null can not be rejected for either fixed or flexible exchange regimes.

[10] See Appendix I (available on request) for details.

[11] In subsequent tests, the WWII observations (1939-1945) are excluded.

Given our failure to reject the null hypotheses that the commodity price variance is equal across all fixed exchange rate periods and across all flexible exchange rate periods, we proceed to test whether the variance is equal across all regimes. The test shows a clear rejection of the null of equal variance across fixed and flexible exchange regimes. However, because of the violation of the normality assumption in the whole $\Delta log(y_t)$ series, the exact confidence levels of the test should not be taken too literally.[12]

A more systematic approach to this issue is to get a careful measurement of the temporal variation in the second moments of the price series by first choosing an appropriate detrending, differencing, or filtering method. The extensive literature on the Prebisch-Singer (PS) hypothesis (of a secular deterioration in the relative price of primary commodities in terms of manufactures) as now considered both trend stationary (TS) and difference stationary (DS) models for the mean of the relative price process.[13]

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[12] Unlike some earlier applications of the ANOVA tests in the literature, we tested the underlying normality assumption. The Jaque-Bera statistics suggested that the null hypothesis of normality can not be rejected for the $\Delta log(y_t)$ series across three fixed (flexible) exchange regimes, except during the fixed exchange rate episode in the Boughton data set. There are, however, clear signs of non-normality for the whole $\Delta log(y_t)$ series in either the GY or the Boughton data sets. When the Bawman-Shenton and Doornik-Hansen tests for normality are used, however, the null of normality is rejected at 5% level in every case except the flexible exchange rate episode in the Boughton data set.

[13] Using their re-constructed price index to estimate a TS model, Grilli and Yang (1988) found evidence supporting the PS hypothesis. Cuddington and Urzúa (1989) found no secular downward trend in the GY index regardless of whether the DS or TS model is used. However, their results crucially rely on the specification of a change in the mean occurred in 1920. Perron (1990) used "additive outlier" method to deal with the structural break in 1920/1921 and again rejected the unit root hypothesis. Cuddington (1992)

The low power of unit root tests in distinguishing the TS and DS alternatives led us to consider both detrending methods, as well as volatility of the actual series around an Hodrick-Prescott (HP) filtered series.[14] As we show below, our "volatility shift" results are robust to the choice of detrending method.

The empirical strategy is then to use dummy variables in the error variance specification for the residuals in the detrended, first-differenced, and HP filtered relative price series:

$$Var\ (\varepsilon_t) = \mu_1 + \mu_2 * FLEX \tag{1}$$

FLEX is a dummy variable that takes the value 0 for fixed exchange periods and 1 for flexible exchange periods. The magnitude of μ_2 measures how much the volatility associated with the flexible exchange regimes differs from that of fixed exchange regimes.

The hypothesis that the variances of ε_t across different exchange regimes are equal amounts to testing H_o: $\mu_2=0$ in equation (1). The test results (employing the Newey-West heteroskedasticity consistent covariance matrix) are summarized in the lower panel of Table 2. Regardless which detrending or filtering method is used, μ_2 is significantly different from zero at the 5% level. Since μ_2 is significantly positive in

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considered 26 individual commodities and found 13 of them to be trend stationary, whereas the other 13 were difference stationary.

[14] A number of studies (Cuddington and Urzúa (1989), Perron (1990)) have pointed out the possible existence of a structural break in 1920/1921. Hence, we also consider including the structural break in the trend specification. The qualitative conclusions of this subsection are unchanged when the structural break is allowed for both the TS and DS models. See appendix II, available from the authors.

every specification considered, there is strong evidence that flexible exchange rate regimes have been associated with a much higher real commodity price volatility than fixed exchange regimes.

2. The Monthly Data Sets on Commodity Price Sub-Indices

The standard deviation of period-over-period percentage changes has been widely used in the macroeconomics literature to measure volatility.[15] There are, however, more rigorous econometric tools now available to allow a simultaneous modeling of the mean and time varying variance in high frequency data. In this study, we use a univariate generalized autoregressive conditional heteroskedastic model (GARCH) to account for volatility clustering.[16] The hypothesis of interest is the extent to which change in nominal exchange regime leads to parameter shifts in the GARCH process for commodity prices. If GARCH parameters vary with exchange regime shifts, this implies changes in the degree to which shocks to commodity price volatility persist over time.

To examine this contention empirically, we first examine whether GARCH provides a good description for the behavior of the six monthly commodity price series, ignoring the possible regime shift effects on the conditional variance. (See appendix III). Into these specifications, exchange rate regime shifts are then introduced.

[15] See e.g. Grilli and Kaminsky (1991), and Flood and Rose (1995).

[16] A referee points out that an alternative to the GARCH model as a way of modeling changing variance is the stochastic volatility models. See Kim et al. (1998) for a comparison of the stochastic volatility and GARCH models.

After appropriate detrending based on the unit root test results, a dummy variable, FLEX, indicating the presence of a flexible exchange regime, is added to the conditional variance equation.[17] Specifically, the residual ε_t is modeled as an GARCH(1,1) process:

$$\varepsilon_t | I_{t-1} \sim N(0, h_t) \tag{2}$$

$$h_t = \delta_1 + \delta_2 * FLEX + \alpha\varepsilon_{t-1}{}^2 + \beta h_{t-1} \tag{3}$$

or the more general:

$$h_t = \delta_1 + \delta_2 * FLEX + \alpha\varepsilon_{t-1}{}^2 + \gamma\varepsilon_{t-1}{}^2 d_{t-1} + \beta h_{t-1} \tag{3'}$$

where FLEX=1 for observations in subsample starting from August 1971 and 0 otherwise. Equation (3') is a Threshold ARCH (TARCH) specification where $d_t = 1$ if $\varepsilon_t < 0$ and 0 otherwise (Glosten, Jagnanathan, and Runkle (1989)). Hence, positive ε_t has an impact of α on the conditional variance while bad news has an impact of $\alpha+\gamma$.

The estimation results are presented in Table 3. For most financial time series, GARCH(1,1) provides a sufficiently good fit. This is also true for the monthly commodity price series. Diagnostic checks (correlograms and ARCH LM tests) confirm that there is no further serial correlation, nor ARCH effects, in the residuals. For real metal price series, TARCH(1,1) provides a better fit than the GARCH(1,1) model. The

[17] Appendix VI repeats the monthly analysis using the month-to-month variability in the U.S. dollar-SDR exchange rate, rather than FLEX, as the quantitative proxy for the exchange rate uncertainty in the global system as a whole. Except for fertilizer, the qualitative results are similar to those using FLEX as the proxy for the international monetary system.

TARCH term is negative, implying that upward price spikes increase volatility.[18] Except in the case of fertilizer,[19] all ARCH and GARCH terms are significantly greater than zero, suggesting the GARCH specification is appropriate for the commodity price data.

Likelihood ratio tests or t-tests can be used to test the restriction $\delta_2=0$ in the GARCH or TARCH specifications [(3) or (3')]. Table 3 reports both. The t-tests indicate that δ_2 is statistically different from zero at 5% significance level for the price series of beverages and metals, and at 10% significance level for agricultural raw material, food and sugar. The likelihood ratio statistics, on the other hand, show that the restriction $\delta_2=0$, is rejected at less than 5% level in all commodity price sub-indexes.[20] With the possible exception of fertilizer, these tests clearly indicate that flexible exchange regimes are associated with a much higher conditional variance and hence a higher unconditional variance.

The sum of $\alpha + \beta + \gamma$ measures the persistence of volatility shocks. Interestingly, this sum clearly declines in all cases when the regime shift variable FLEX is included. (Moreover, it no longer exceeds unity, which was a problem in the case of real food prices when exchange regime shifts were ignored; compare Appendix III results). Diebold (1986, p.55) suggests if there are regime shifts in the underlying data generating

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[18] The latter finding is consistent with Deaton and Laroque's (1992) "stock-out" model, which suggests that the conditional variance is increasing in prices. Price surges due to collapsing supply or booming demand may lead to an inventory "stock-out". In this situation, inventories are not available to perform their typical buffer stock role in stabilizing prices.

[19] For real fertilizer price index, the ARCH term is negative but not significant. A negative ARCH term cannot ensure that the conditional and unconditional variances are positive for all realizations of ε_t.

[20] The statistic has a chi-square distribution with one degree of freedom under the null hypothesis.

process, failing to account for them may lead to the conclusion that shocks to volatility are more persistent than they actually are. The series may appear to follow *integrated* ARCH error processes due to this specification error. The evidence reported here is an empirical example of this pitfall.

IV. Concluding Remarks

This paper investigates differences in real primary commodity price volatility across fixed and flexible exchange rate regimes. The findings suggest that the exchange rate arrangement may imply an important source of systematic risk to world commodity markets. Using three different data sets over different sample periods, as well as alternative detrending methods, there is strong evidence supporting the conjecture that the flexible exchange periods have been associated with higher real commodity price volatility than the fixed exchange periods.

These conclusions regarding commodity price volatility should be of interest to international macroeconomists, as well as those studying global commodity markets, for at least two reasons. First, the puzzle of why commodity price volatility varies with exchange rate arrangement may well be related to a puzzle in the literature regarding fundamental determinants of the exchange rate. Flood and Rose (1995) find that there are no significant changes in the time series processes of output, money, or prices across exchange regimes. Since the volatility of nominal exchange rates exhibits apparent shifts across exchange regimes, they conclude that the above macro aggregates can not be the true fundamentals for exchange rates. It would be interesting to see whether this problem of identifying underlying fundamentals is also present in international commodity markets.

Second, from a theoretical perspective it remains to be determined whether our finding that the relative price of two tradable goods increases under flexible exchange rates can be explained in the context of equilibrium models -- even those described by Stockman as exhibiting 'nominal exchange rate nonneutrality.' The explanation behind the non-neutrality phenomenon is still unresolved.

The empirical results here, however, suggest that the mechanism linking the exchange rate to primary commodity prices is different than the mechanism linking exchange rates and manufactured goods prices. This in turn seems to reflect some form of price stickiness, perhaps arising from different market structures for the two types of goods. Using an industrial organization model, Liang (1998b) demonstrates how market structures and economic agents' perception about future exchange rate movements may influence the volatility of relative commodity prices. Recent work by Goldberg and Knetter (1997) also concludes that empirical price-exchange rate relationship may reflect underlying differences in market structure across industries.

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Figure 1

Real Commodity Price Indices and Exchange Rate Regime

Long-Span Annual Data;

Percentage changes over pervious period

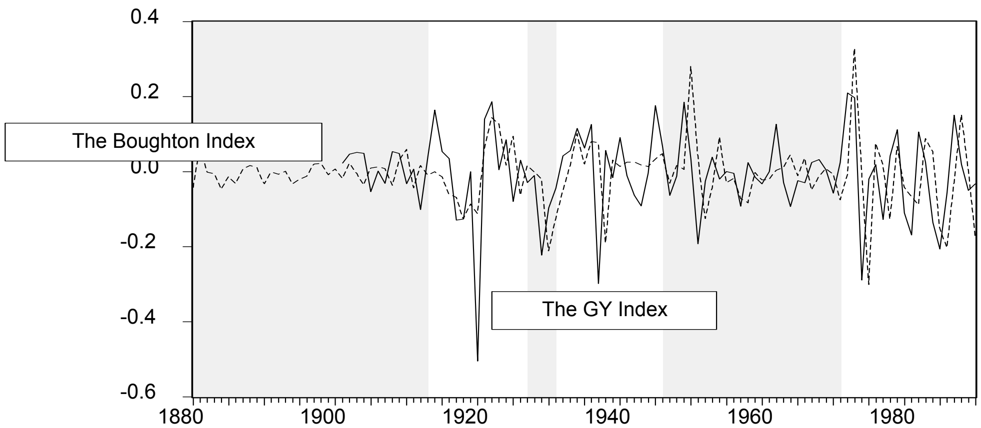

Figure 2
Real Commodity Price Indices and Exchange Rate Regime
January 1957-May 1996;
percentage changes over previous period

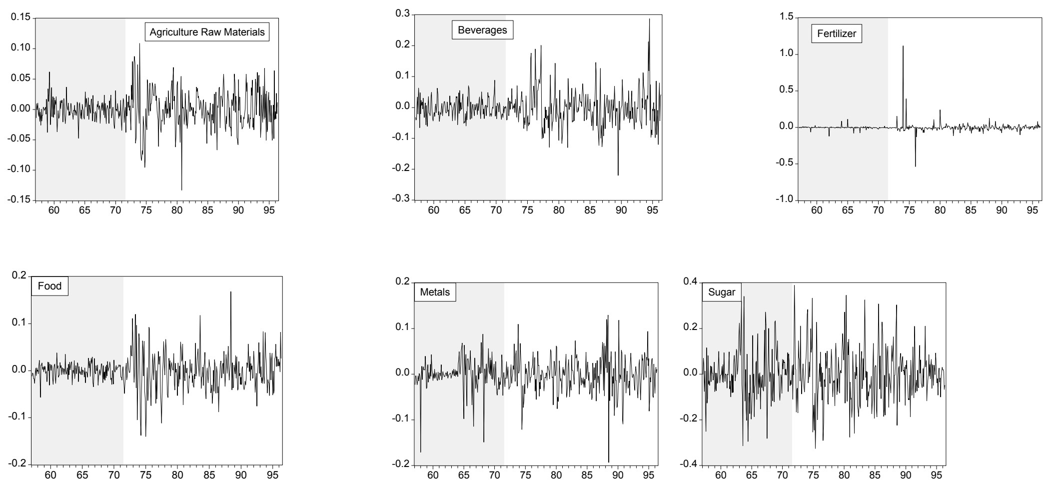

Table 1
Statistical Summary of Real Commodity Price Volatility Across Exchange Regimes

(percentage changes over previous period)

	Fixed	Free Float	Fixed	Controlled Float	Fixed	Controlled Float
	1880-1913	*1914-1926*	*1927-1931*	*1932-1938*	*1946-1971*	*1972-1992*
GY Index: 1900-92						
	(13 periods)	*(13 periods)*	*(5 periods)*	*(7 periods)*	*(26 periods)*	*(21 periods)*
mean	0.45	-0.93	-5.46	0.75	0.11	-1.80
std dev	4.83	17.25	9.24	13.56	7.92	13.09
Boughton Index: 1880-92						
	(33 periods)	*(13 periods)*	*(5 periods)*	*(7 periods)*	*(26 periods)*	*(19 periods)*
mean	0.10	-0.42	-1.23	-0.93	-0.42	0.36
std dev	4.23	8.56	11.22	12.75	9.79	19.41

IFS Index: 1957 January--1996 May (Standard Deviations)		
	Fixed: Jan. 1957-July. 1971	**Flexible: Aug. 1971-May 1996**
Agricultural Raw Material	1.50	3.20
Beverages	2.82	5.95
Fertilizer	1.79	8.13
Food	1.42	3.78
Metals	3.24	4.00
Sugar	10.43	11.50

Table 2
Volatility Analysis of the Long-run Data

ANOVA Tests (*critical values in parenthesis*)		
Ho: Equal variance across fixed (flexible) periods		
	Fixed Exchange Rate Episodes (44 and 64 periods respectively for GY and Boughton data sets)	Flexible Exchange Rate Episodes (41 and 39 periods respectively for GY and Boughton data sets))
GY Data Set	0.94 (3.23)	1.04 (3.23)
The Boughton Index	1.67(3.15)	1.80 (3.23)
Ho: Equal variance between fixed and flexible periods		
GY Data Set		2.62* (2.11)
The Boughton Index		4.02* (2.02)
Regression Tests: Estimated Values of $\mu_2 \times 10^2$ in Equation (1) (*t-statistics in parenthesis*)		
The TS Model		
The GY Index		1.28* (2.07)
The Boughton Index		0.70* (2.29)
The DS Model		
The GY Index		1.48* (2.15)
The Boughton Index		0.78* (2.65)
The HP Filter		
The GY Index		1.09* (2.28)
The Boughton Index		0.90* (3.20)

Asterisks (*) indicate that the null hypothesis can be rejected at the 5% significance level.

Table 3
Accounting for Exchange Regime Shifts in the GARCH(1,1) Model

	$\delta_1 \times 10^3$	$\delta_2 \times 10^3$	α	γ	β	***likelihood ratio test comparing models w/out FLEX***
Agricultural Raw Material	0.03 *(1.53)*	0.07 *(1.67*)*	0.18 *(2.80)*		0.71 *(5.84)*	10.10**
Beverages	0.07 *(1.70)*	0.24 *(2.11**)*	0.11 *(2.46)*		0.80 *(10.20)*	18.58**
Fertilizer	0.10 *(0.98)*	1.68 *(1.30)*	-4.37×10^3 *(-1.08)*		0.76 *(2.87)*	53.70**
Food	0.02 *(1.74)*	0.10 *(1.77*)*	0.09 *(2.04)*		0.80 *(9.45)*	26.34**
Metals	7.26×10^{-3} *(2.48)*	0.07 *(3.17**)*	0.30 *(4.20)*	-0.33 *(-4.25)*	0.82 *(23.34)*	26.30**
Sugar	0.60 *(2.08)*	1.88 *(1.64*)*	0.27 *(3.82)*		0.66 *(7.69)*	5.88**

*Rejecting the null at the 10% significance level.

** Rejecting the null at the 5% significance level.