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**Will the Emergence of the Euro
Affect World Commodity Prices?**

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ABSTRACT

The emergence of the euro as a key currency, perhaps eventually rivalling the US dollar in importance, may have important macroeconomic implications for industrial as well as developing economies in the years ahead. This paper focuses on two related questions. First, what effects, if any, will the euro have on the volatility of world primary commodity prices? Second, why should the impact of the euro on real commodity prices be of interest to economic analysts and policy makers?

Our econometric analysis provides evidence that episodes of internal stability of exchange rates among the 11 euro countries during 1957-98 were associated with periods of lower real commodity price volatility. These stabilizing effects are statistically significant for fertilizer, metals and petroleum, and cereals. A reasonable inference, therefore, is that the establishment of the euro on 1 January 1999 should be expected to contribute to reduced volatility of world commodity prices, other things equal, if the alternative is the higher euro-11 volatility in the 1970s. On the other hand, a move from the relatively stable euro-11 environment of the early 1990s to a single currency (with *zero* intra-euro-11 volatility) will have only a modest impact (reducing) commodity price volatility.

Keywords: Commodity prices, volatility, exchange rate regimes, euro

JEL classification:

F33 (International monetary arrangements and institutions)

O13 (Agriculture, natural resources, …., other primary products)

F41 (Open economy macroeconomics)

I   OVERVIEW

The emergence of the euro as a key currency, perhaps eventually rivalling the US dollar in importance, may have important macroeconomic implications for industrial as well as developing economies in the years ahead. This paper focuses on two related questions. First, what effects, if any, will the euro have on the volatility of world primary commodity prices? Second, why should the impact of the euro on real commodity prices be of interest to economic analysts and policy makers?

The second question is addressed first in order to motivate our subsequent empirical analysis. Section II provides background information on the importance of commodity trade for the EU countries, as well as other country groups. It then reviews earlier work by Cuddington and Liang (1997), who conclude that the choice of exchange regime in the world economy has had an important effect on world commodity price volatility. Section III discusses proxies for exchange rate volatility for the global financial system as a whole and the 'euro-11' countries, respectively.[1] Section IV provides some graphical evidence on commodity price volatility and exchange rate volatility within the euro-11 countries and elsewhere. Section V presents a more formal econometric examination of whether currency volatility among the euro-11 has affected the volatility of world commodity prices. Using data from January 1957 through December 1998, we find evidence that episodes of greater exchange rate stability within the euro-zone has been associated with lower volatility of world prices for some, but not all, of the primary commodity prices studied. The last section concludes by making inferences about the likely impact of the emergence of the euro on world commodity price volatility in the future. It also suggests directions for future research.

II   MOTIVATION AND BACKGROUND

Primary commodities still loom large in international trade. According to Marian Radetzki (1990: 5):

> In 1985, primary commodities as defined here accounted for some 40 per cent of overall exports of goods. The importance of commodities in total trade has been declining over time, and that of manufactures has been rising. The share of commodities was 48 per

cent in 1965 and 46 per cent in 1975. Despite the decline, primary commodities are still sufficiently important to have a substantial impact on the world macroeconomy. Changes in the prices of commodities are known to have strong repercussions, for instance, on global inflation rates, or on the exchange rates and levels of economic activity of the raw materials exporting countries.

Within this general trend, he identified an important change *within* the aggregate of primary commodities:

> The fuels group has emerged as a dominant component of overall commodity exports. Oil dominates the group. In the 1980s, nine-tenths of the value of fuel exports comprised oil. As a result of the petroleum price increases in the 1970s, the share of this group in primary commodity trade rose from 20 per cent in 1965 to 42 per cent in 1975 and 47 per cent in 1985 … With falling petroleum exports and energy prices from the mid-1980s, the share of fuels in total commodity exports and in global trade declined in subsequent years'. (page 5)

Using Radetzki's figures, Table 1 calculates the shares of fuel and non-fuel commodities, respectively, in total merchandise exports. The figures show that fuel commodity exports rose from roughly 10 per cent of total exports in 1965 to 19 per cent in the 1970s and 1980s. In contrast, the share of non-fuel exports has fallen steadily from 38 per cent of total goods exports in 1965 to 21 per cent in 1985. Radetzki's data only cover the period through

TABLE 1
THE DECLINING SHARE OF NON-FUEL COMMODITIES
IN TOTAL EXPORTS OF GOODS

	1965	1975	1985
1. Commodity trade: share of total exports of goods	48%	46%	40%
2. Fuel commodities: share of *commodity* trade	20%	42%	47%
3. Non-fuel commodities: share of *commodity* trade	80%	58%	53%
4. Fuel commodities: share of total exports of goods (row 1x2)	10%	19%	19%
5. Non-fuel commodities: share of total exports of goods (row 1x3)	38%	27%	21%

Source: Radetzki (1990: 5) for rows 1-3; rows 4-5 are our calculations.

1985. An April 1999 press release by the World Trade Organization (WTO), however, suggests that the decline in the share of (fuel and non-fuel) primary commodities has continued in the 1990s: 'Commodity prices fell sharply in 1998, pushing the share of primary products in world exports below 20 per cent in current price terms for the first time in the post-war period'. Steep declines in both oil and non-oil primary commodity prices were a major contributing factor: 'Oil prices fell by 30 per cent and non-oil commodity prices fell by 20 per cent in 1998, with very different implications for various countries and regions of the world'.

The European Union is a significant net exporter or net importer for a number of primary commodities, as the data by SITC categories in Table 2 show.[2] Intra-EU commodity flows are also important, especially for agricultural crops. The WTO *Annual Report* (1998: Table IV.4) shows that intra-Western European exports of agricultural products are valued at $177.6 billion, two and one half times larger than intra-regional flows within other regional areas. Agricultural exports from Latin America to Western Europe and from North America to Western Europe were also substantial: $22.5 billion and $17.8 billion, respectively, in 1997. Much of European agricultural trade has been governed by the Common Agricultural Policy (CAP) since the early 1960s. By 1984, the CAP accounted for over 70 per cent of EC spending (Dinan 1994: 329). This suggests that changes in commodity price behaviour could have potentially large fiscal transfer consequences for euro-11 countries.

Many developing countries still rely heavily on the primary commodity exports. Table 3 shows, by region, countries that rely on the earnings of one primary commodity for more than 50 per cent, 20-49 per cent, and 10-19 per cent, respectively, of their total export earnings. As the list indicates, most commodity dependent exporters are developing, rather than industrialized, countries.

The behaviour of primary commodity prices—including their long-term trends, cycles, and short-run volatility—has differed considerably from the behaviour of manufactured goods. Regarding long-term trends, the Prebisch-Singer hypothesis in the economic development literature conjectured that there would be an ongoing secular deterioration in the relative price of primary commodities in terms of manufactures. Empirical evidence on the PS hypothesis remains mixed, in spite of extensive testing in the literature.[3]

TABLE 2
COMMODITY EXPORTS BY SITC CLASSES AND REGIONS
(PER CENT SHARES OF TOTAL EXPORTS IN RESPECTIVE SITC CATEGORY)

SITC code	SITC classes	Developed economies			Developing economies	E. Europe & former USSR
		Year	EU	Total		
		Origin of exports of major commodity classes				
0-9	All classes	1990	43.0	71.3	23.4	5.3
	(Primary commodities; intermediate goods; manufactured goods)	1996	39.6	67.5	28.7	3.8
0&1	Food and live animals	1990	45.8	68.5	28.4	3.1
	Beverages and tobacco	1996	43.3	66.4	30.5	3.1
2&4	Crude materials, oil	1990	29.0	63.2	28.9	8.0
	and fats (fuel excluded)	1996	25.2	59.4	34.1	6.5
3	lubricants, mineral	1990	14.0	27.5	56.7	15.8
	Fuels and related material	1996	14.9	32.5	56.6	10.9
		Destination of exports of major commodity classes				
0-9	All classes	1990	42.5	70.4	23.4	4.3
	(Primary commodities; intermediate goods; manufactured goods)	1996	37.0	65.5	29.0	3.8
0&1	Food, live animals	1990	48.1	71.3	22.9	5.2
	Beverages and tobacco	1996	44.0	67.1	25.4	5.7
2&4	Crude materials, oil	1990	41.8	68.3	25.9	5.2
	and fats (fuel excluded)	1996	35.9	62.0	32.1	3.8
3	lubricants, mineral	1990	31.9	61.5	24.2	6.8
	Fuels and related material	1996	31.6	62.1	25.1	5.1

Notes: These data are extracted from 1996 *International Trade Statistics Yearbook*, volume 2, Special Table D: 'Structure of world exports by commodity classes and by regions'. The original table has disaggregated regional data for all SITC categories at the 1 digit level, in the periods 1990 and 1993-96.

There seems to be less controversy about the higher cyclicity and volatility of primary commodity prices. Regarding their cyclical behaviour, primary commodity prices are sometimes used as a leading indicator of inflation, and touted as a useful intermediate target for monetary policy. It is increasingly

TABLE 3
COUNTRIES DEPENDENT ON A SINGLE PRIMARY COMMODITY
FOR EXPORT EARNINGS
(Annual average of exports in dollars, 1992-97)

	50 % or more of export earnings	20-49 % of export earnings	10-19 % of export earnings
MIDDLE EAST			
Crude petroleum	Bahrain, Saudi Arabia, Iran, Islamic Rep. of Iraq, Kuwait, Libya, Oman, Qatar, Yemen Rep.	Syrian Arab Rep. United Arab Emirates	Egypt
Aluminium			Bahrain
AFRICA			
Crude petroleum	Angola, Gabon, Nigeria, Congo, Rep.	Cameroon Equatorial Guinea	Algeria
Natural gas		Algeria	
Iron ore		Mauritania	
Copper	Zambia		Congo, Dem. Rep.
Gold		Ghana, South Africa	Mali, Zimbabwe
Timber (African hardwood)		Equatorial Guinea	Central African Rep. Swaziland, Gabon, Ghana
Cotton		Benin, Chad, Mali, Sudan	Burkina Faso
Tobacco	Malawi	Zimbabwe	
Arabica coffee	Burundi, Ethiopia	Rwanda	
Robusta coffee	Uganda		Cameroon
Cocoa	Sao Tome & Principe	Côte d'Ivoire, Ghana	Cameroon
Tea			Kenya, Rwanda
Sugar		Mauritius	Swaziland
Crude petroleum	Venezuela	Ecuador, Trinidad &Tobago	Colombia, Mexico
Copper		Chile	Peru
Gold			Guyana
Cotton			Paraguay
Arabica coffee			Colombia, Guatemala Honduras, Nicaragua, El Salvador
Sugar		Guyana, St. Kitts & Nevis	Belize
Bananas		St. Vincent, Honduras	St. Lucia, Costa Rica, Ecuador
Fishmeal			Peru
Rice			Guyana
			Table continues

5

Table 3 continues

	50 % or more of export earnings	20-49 % of export earnings	10-19 % of export earnings
WESTERN HEMISPHERE			
Crude petroleum	Venezuela	Ecuador, Trinidad &Tobago	Colombia, Mexico
Copper		Chile	Peru
Gold			Guyana
Cotton			Paraguay
Arabica coffee			Colombia, Guatemala Honduras, Nicaragua, El Salvador
Sugar		Guyana, St. Kitts & Nevis	Belize
Bananas		St. Vincent, Honduras	St. Lucia, Costa Rica, Ecuador
Fishmeal			Peru
Rice			Guyana
EUROPE, ASIA AND THE PACIFIC			
Crude petroleum		Azerbaijan, Papua New Guinea, Brunei Darussalam, Norway, Russia	Indonesia, Kazakhstan Vietnam
Natural Gas	Turkmenistan		
Aluminium		Tajikistan	
Copper		Mongolia	Kazakhstan, Papua New Guinea
Gold		Papua New Guinea	Uzbekistan
Timber (Asian hardwood)		Lao P.D.R. Solomon Islands	Cambodia, Papua New Guinea, Indonesia, Myanmar
Timber (softwood)			Latvia, New Zealand
Copra & coconut oil	Kiribati		
Cotton		Pakistan, Uzbekistan	Azerbaijan, Tajikistan, Turkmenistan

Source: Cashin, Liang, and McDermott (1999).

recognized that international commodity markets, including oil but other commodities as well, play an important role in explaining post World War II business cycles in the US and other industrial economies, and in transmitting macro shocks internationally.[4]

Reinhart and Wickham (1994: 186) point out that 'there is a sustained and sharp increase in the variance of commodity prices [over their sample period 1957: I-1993: II]. The increase is evident in all the indices but is most pronounced in all the commodities and food groupings'. Not only has

commodity price volatility risen over time, commodity prices have consistently exhibited greater volatility than have manufactured goods. Figure 1 nicely illustrates the fact that commodity price volatility has exceeded that of manufactured goods. This has been true for decades of either high or low volatility.

FIGURE 1
THE INSTABILITY INDEX FOR MANUFACTURES AND COMMODITIES, 1990-92
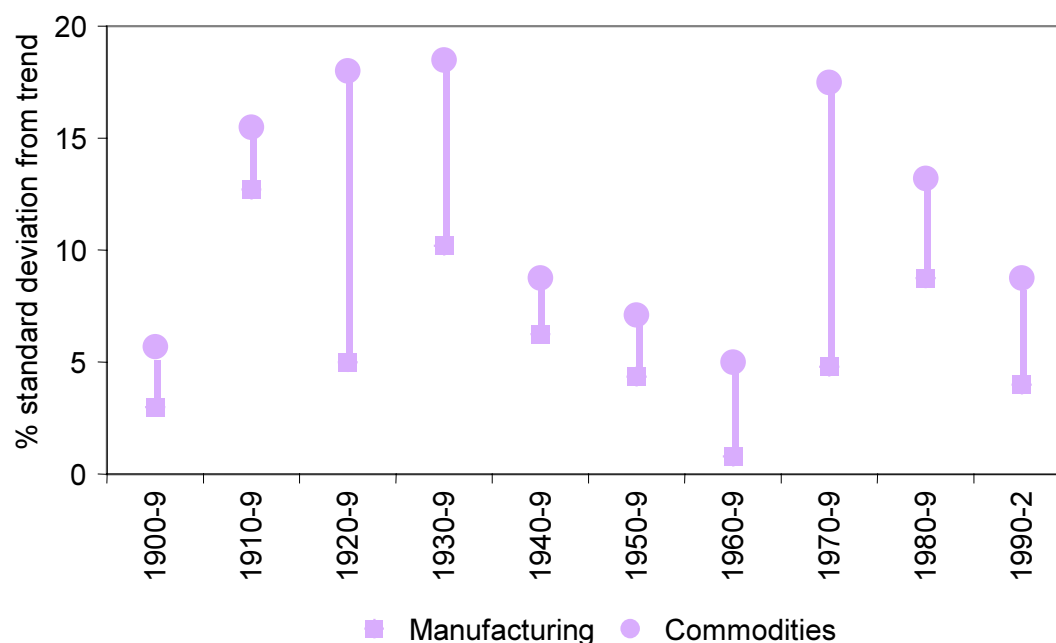
Source: World Bank (1994).

Some authors have noted that primary commodities seem to have volatility that is the same order of magnitude as the volatility of exchange rates or stock prices. Using data covering the period of 1972-82, Bui and Pippenger (1990) find that primary commodity prices, such as the prices of rubber and tin, that are determined in the well-organized auction markets were *more* volatile than exchange rates. This has led economic modellers to suggest that world primary commodity prices should be thought of as 'asset prices'. See, e.g., Jeff Frankel's (1986) model of commodity price 'overshooting' in response to money supply shocks, which adapts the classic Dornbusch exchange rate overshooting model to explain the volatility of commodity prices (but in a closed economy context).

Reduced-reform and structural models of the macroeconomic determinants of real commodity prices typically highlight the important roles of worldwide industrial production, the real US dollar exchange rate, and

changes in commodity export supply. See, e.g., Borensztein and Reinhart (1994), Dornbusch (1985), and Gilbert (1989). Borensztein and Reinhart (1994: Table 9) show a variance decomposition analysis for real commodity prices for the period 1971: I-1992: III and three subperiods: 1971: I-1984: IV, 1985: I-1988: IV, and 1989: I-1992: III. They conclude that the 'real exchange rate of the US dollar explains a fairly stable proportion of the variance of commodity prices throughout the subperiods' (1994: 256).[5]

A number of analysts have emphasized the strong link between movements in the (nominal and real) dollar exchange rate and the relative price of commodities in terms of manufactures. Hence, it is natural to ask: does only the dollar exchange rate matter? Should we consider the differential effects of the movements in several key currencies? Gilbert (1989) argues convincingly that the appropriate definition of the effective exchange rate is critical in getting meaningful results when studying the impact of exchange rate changes on world commodity prices.

International macroeconomists and policy makers have long been interested in the impact of alternative exchange arrangements on international trade and financial flows. It has been argued that under flexible exchange rate regimes, export and import prices become more volatile, thereby adding additional risk to export and import transactions. The impact of this additional price uncertainty on export and import volumes, however, has been difficult to isolate empirically.[6] One possible explanation is the increased sophistication of firms in hedging exchange rate-induced price risks.

Mussa (1986) and others have emphasized that flexible exchange rate regimes imply higher volatility of real as well as nominal exchange rates. Subsequent researchers have, however, been unable to detect the impact of choice of exchange regime on many, if any, other macroeconomic aggregates. See, e.g., Baxter and Stockman (1989), Flood and Rose (1995) and Rogers (1995).

Cuddington and Liang (1997) demonstrate, using three alternative datasets on commodity prices, that the volatility of real primary commodity prices has been higher under flexible exchange regimes. That analysis covered a number of different exchange regimes over the period from the 1880s to 1996. Exchange regimes were distinguished either using a dummy variable, DUMFLEX, that takes the value of zero during fixed exchange rate episodes and one during flexible exchange rate periods, or a measure of US

dollar/SDR volatility. Either delineation of exchange regimes does not, of course, imply that *all* countries shared the same exchange regime; rather they are rough characterizations of the international monetary system as a whole. For example, during the Bretton Woods fixed exchange rate period, most countries fixed their exchange rates in terms of US dollars, which in turn was pegged to the price of gold. Yet, Canada had a flexible exchange rate for much of this period. Similarly, the post-Bretton Woods period after 1973 is characterized as a floating rate period, even though many LDCs (initially, at least) retained fixed exchange rates. Within Europe, as well, a number of countries adopted exchange rate arrangements exhibiting 'limited flexibility'. Examples include: (i) the so-called 'snake in the tunnel' arrangement, which came into effect in 1972, following the 1970 Werner Report; (ii) the European Monetary System (EMS) with its ECU-based exchange rate mechanism (ERM), which emerged in 1979; and ultimately (iii) the euro system, which came into existence on 1 January 1999.[7]

III   INDICATORS OF EXCHANGE RATE VOLATILITY

Having provided some background and motivation we now turn to the more analytical of our two questions in the Introduction: what effects, if any, will the emergence of the euro have on the volatility of world primary commodity prices (denominated in US dollars)? In order to use historical data to shed light on the inherently forward-looking question being addressed in this paper, we ask: were periods of relative stability among the bilateral exchange rates of the euro-11 countries between 1957 and 1998 associated with relatively stable commodity prices?

To address this question we extend Cuddington-Liang (1997), who investigated the extent to which commodity price volatility can be explained by the choice of worldwide exchange regime (measured by a dummy variable indicating the worldwide exchange regime) or dollar exchange rate volatility. During the 1957-98 period, the US dollar was the dominant currency in the international monetary system. Hence the chosen quantitative proxy for the exchange rate uncertainty in the global market as a whole is the volatility of the US dollar/SDR rate, denoted as *var_u*s.[8] Volatility is defined here as the square of the period-to-period change in the logarithm of the US dollar/SDR exchange rate.

The IMF's *International Financial Statistics* (IFS) publication provides the following background on the SDR: 'Beginning in July 1974, the value of the

FIGURE 2
VOLATILITY OF THE US DOLLAR/SDR EXCHANGE RATE
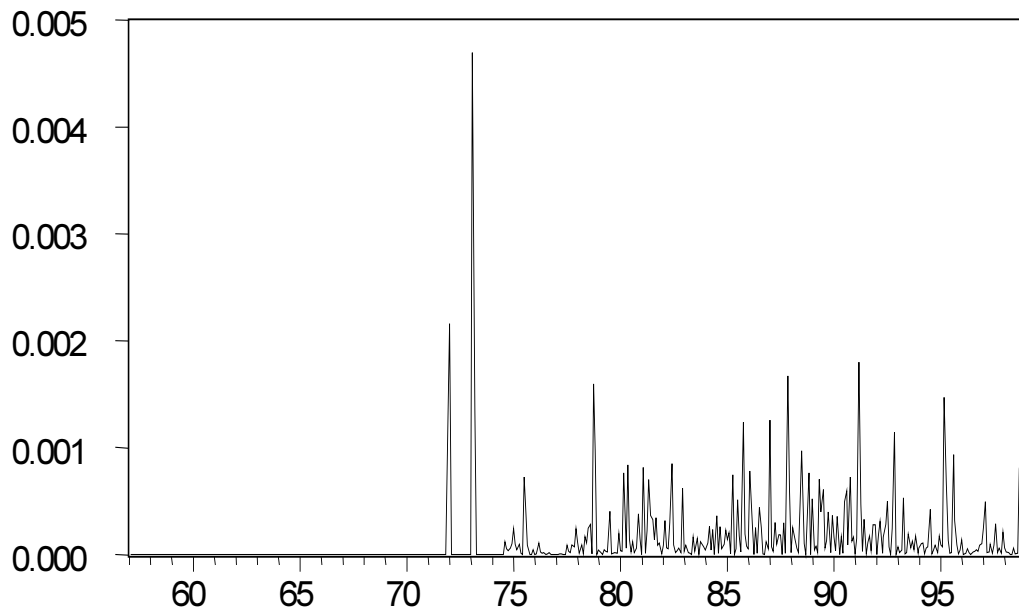
FIGURE 3
VOLATILITY OF THE EURO-11 CURRENCIES
(VIS-À-VIS THE DEUTSCHMARK)
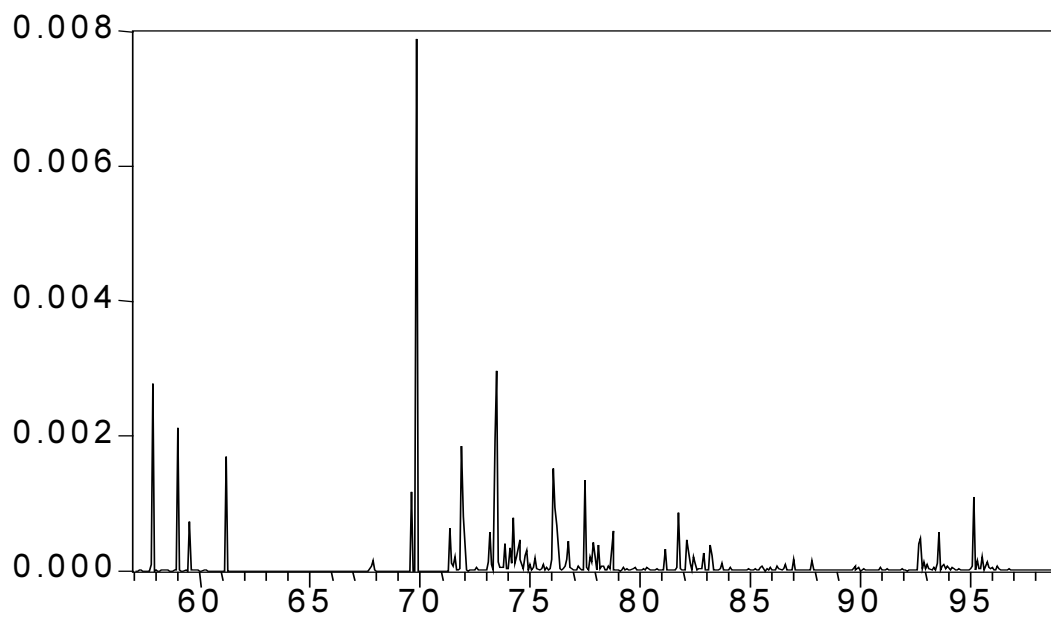
SDR is determined by the Fund on the basis of a basket of currencies with each currency assigned a weight in the determination of that value. In the derivation of the SDR value, the currencies in the basket are valued at their market exchange rates for the US dollar, and the US dollar equivalents of each of the currencies are summed to yield the rate of the SDR in terms of the US dollar. The method of calculating the U. dollar/SDR exchange rate remains the same, although the number and weights of currencies in the SDR basket have changed over time' (IFS Feb. 1995: x-xi). Since 1 January 1981, the value of the SDR has been based on a basket of five currencies.[9] The composition and currency amounts in the SDR valuation basket were last revised as of 1 January 1996. It currently contains 0.582 US dollars, 0.446 Deutsche marks, 27.2 Japanese yen, 0.813 French francs, and 0.105 pound sterling (IMF 1998: 93, Table 15).[10] Thus, the dollar/SDR rate is currently defined as:

(1) $S(\$/SDR) = .582 + .446\,S(\$/DM) + 27.2\,S(\$/yen) + .813\,S(\$/FF) + .105\,S(\$/pound)$.

Given this definition of the USD/SDR exchange rate, it is clear that *var_us*, which is the square of its period-to-period percentage, will be affected not only by the squares of the percentage changes in the five component currencies, but also their pairwise correlations.[11] As Figure 2 shows, the values of var_us were zero, expect for a couple of realignments in the early 1970s, reflecting the exchange rate stability of the Bretton Woods period. *Var_us* increased considerably thereafter.[12]

For the question in this paper, an index for the German effective exchange rate vis-à-vis the other 10 euro currencies was constructed by taking the weighted average of the 10 bilateral nominal exchange rate against the German mark.[13] Our measure of intra-euro currency volatility, *var_euro*, is defined as the percentage change of this constructed German effective exchange rate over previous period squared. The formula is analogous to that for the dollar/SDR rate above.

Var_euro is shown in Figure 3. A comparison of Figures 3 and 4 shows that the behaviours of volatility measures for the US/SDR exchange rate and the euro-11 exchange rates are quite different over the sample period. The correlation coefficient between the two volatility indices is only 0.06.

Presumably the emergence of the euro-zone will be associated with reduced volatility in the *var_euro* index. The hypothesis to be tested is whether

reduced nominal exchange rate volatility within the euro-11 countries leads to greater stability in real commodity prices, other things equal (including exchange rate volatility in the non-euro countries).

IV   VOLATILITY OF REAL COMMODITY PRICES AND MAJOR WORLD CURRENCIES

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 the manufacturing unit value (muv) index. The latter is a unit value index for manufactured good exports from a number of industrial countries.[14]

TABLE 4
SPECIFICATIONS AND SUMMARY STATISTICS FOR COMMODITY PRICES
(JANUARY 1957–DECEMBER 1998)

Definition of commodities	
Beverages	coffee, cocoa beans, tea
Food	wheat, maize, rice, soybeans soybean meal, soybean oil, palm oil, coconut oil, fish meal, groundnut oil, beef, lamb, sugar, bananas
Cereals	wheat, maize, rice
Fertilizer	phosphate rock, triple super-phosphate
Metals	copper, aluminium, iron ore, tin, zinc, lead
Petroleum	crude oil, gasoline, natural gas, heating oil
Agricultural raw materials	hardwood, softwood, cotton, wool, rubber, tobacco, hides

Summary statistics
for first difference of the MUV-deflated commodity prices indices

	Mean	Median	Maximum	Minimum	Std. dev.	Kurtosis
Beverages	-0.0019	-0.0033	0.3115	-0.1940	0.0521	7.5302
Food	-0.0016	-0.0002	0.1420	-0.1368	0.0312	6.1128
Cereals	-0.0020	-0.0031	0.2862	-0.1423	0.0383	10.4097
Fertilizer	-0.0007	-0.0019	1.1183	-0.5254	0.0638	201.5273
Metals	-0.0017	-0.0011	0.1259	-0.2263	0.0365	7.8668
Petroleum	1.31E-05	-0.0017	1.2164	-0.2721	0.0828	97.7781
Agricultural raw materials	-0.0003	0.0009	0.1151	-0.1219	0.0275	5.5766

Source: The IMF's International Financial Statistics.

FIGURE 4
MONTH-TO-MONTH PERCENTAGE CHANGES IN REAL COMMODITY PRICES
(JANUARY 1957-DECEMBER 1998)

Beverages

Fertilizer

Cereal

Food

Petroleum

Metals

Figure 4 continues

13
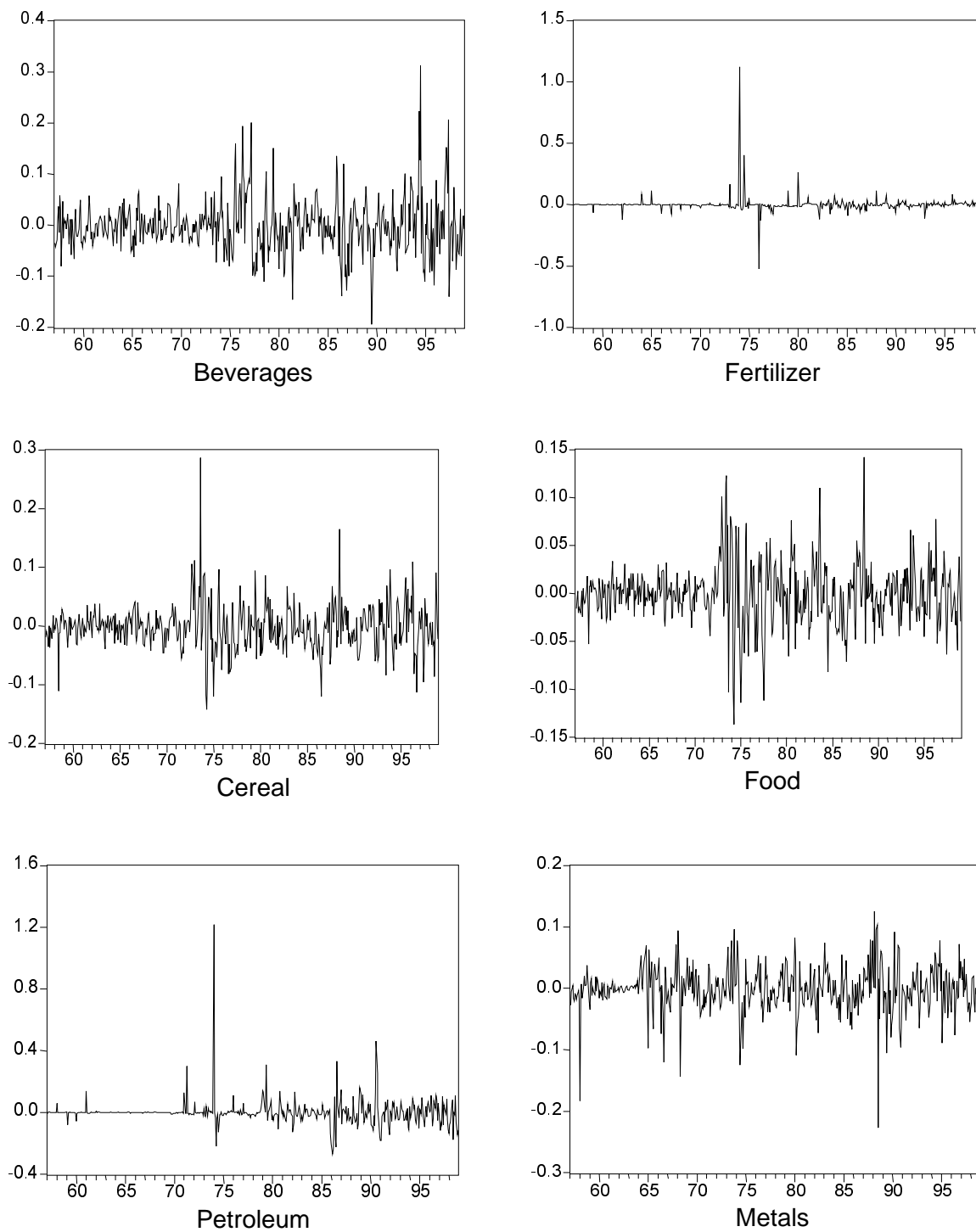
Figure 4 continues
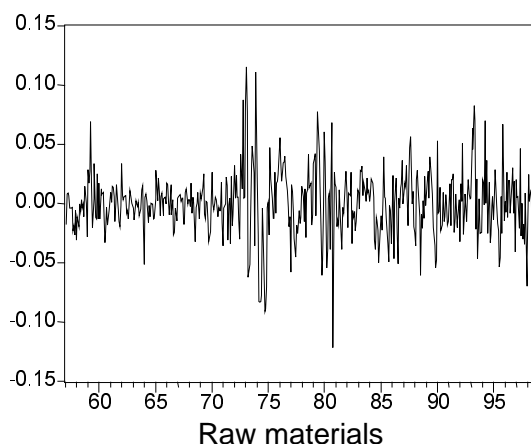
Raw materials

The IFS reports monthly commodity price indexes for agricultural raw materials, beverages, fertilizer, food, metals, and petroleum from January 1957 through December 1998. In addition, a sub-category of the food index, the cereals, is also considered since it is one major item covered by the European Common Agricultural Policy (CAP). Table 4 provides a detailed description of each commodity index and some summary statistics for their first-differences.

Figure 4 plots the percentage change over previous period (i.e. the first-difference of the natural logarithm) for each real commodity price series. Careful visual examination of Figures 2-4 suggests apparent co-movement between the volatility of beverages, food and agricultural raw materials with the volatility of the US dollar/SDR rate. The volatility of petroleum prices, on the other hand, may be correlated with euro-11 exchange rate volatility. A more systematic examination of these conjectures is presented in the next section.

V   ECONOMETRIC ANALYSIS

5.1   Trends in the commodity price series

Before measuring the temporal variation in the second moments of a time series, it is important to consider alternative specifications of its trend. An ongoing debate among scholars testing the PS hypothesis has been whether

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real commodity prices are better modelled as trend stationary (TS) or difference stationary (DS) processes.[15] These two univariate specifications for the trend in a time series, $\log y_t$ – which is the natural logarithm of real commodity price indices in our context—can be nested within the following general specification:

(2) $\log y_t = \alpha + \beta * time + e_t$

where e_t follows a possibly integrated ARMA process:

(3) $(1 - \rho L) * A(L) * e_t = B(L) * \varepsilon_t$.

A(L) and B(L) are autoregressive and moving average lag polynomials. The largest root in the AR polynomial, ρ, is factored out for convenience. When $|\rho| < 1$, the model (2)-(3) is just a log-linear time trend model with an ARMA error process accounting for the transitory fluctuations around a deterministic trend. For the case of an integrated error process where, by definition, $\rho = 1$, the model (2)-(3) collapses to a DS (or 'stochastic trend') model with a stochastic growth rate equal to:

(4) $d \log y_t = \beta + u_t$

and a stationary ARMA error process:

(5) $A(L) * u_t = B(L) * \varepsilon_t$.

In the latter case, $\log y_t$ is called a unit root process with (possibly non-zero) drift β. The error process in (5) is assumed to be stationary, so that fluctuations around the stochastic drift in the series are transitory.

Different detrending methods can differ markedly in the degree to which they filler out particular frequencies (Baxter 1988), implying that measures of conditional and unconditional volatility may be quite sensitive to the choice of detrending method. Our approach is to carry out Phillips-Perron[16] (PP) unit root tests on each series. If the null hypothesis of a unit root is not rejected, we fit a DS model and consider the volatility in the error process in that specification in our subsequent analysis. If the unit root hypothesis is rejected, we use a TS model instead.

Both Dickey-Fuller and Phillips-Perron unit root tests involve running an auxiliary regression of the form:[17]

(6) $d \log y_t = \omega + \phi \log y_{t-1} + \theta * time + \zeta_t$

and testing the hypothesis that $1-\rho \equiv \phi = 0$ using appropriate critical values obtained from Monte Carlo simulations. (The two tests differ in terms of how they cope with serially correlated errors in (6), but this need not concern us here.)

Table 5 reports the results of the Phillips-Perron tests on the real commodity price series. The power of unit root tests, as well as the appropriate critical values, depend on whether an intercept and time trend are included in the regression used to obtain the PP statistics. The tests in Table 5 were based on the general-to-specific methodology (GTS) of first including both a constant and a time trend in the estimation. If the null hypothesis of a unit root is not rejected in the most general specification, the significance of the trend is tested to see if it can be omitted, thereby increasing the power of the unit root test. If the null hypothesis is still not rejected, the significance of the intercept term in the PP regression is tested to see if it too can be dropped before the PP unit test is re-done on the more parsimonious specification. The third column of the Table indicates whether a time trend (T), an intercept (C), or neither term (N) is included in the PP regression equation chosen using the GTS method.[18]

TABLE 5
PHILLIPS-PERRON UNIT ROOT TESTS

	PP's τ statistic	Final specification of the PP test (GTS method)
Beverages	0.40	N
Fertilizer	0.08	N
Food	0.85	N
Cereals	0.79	N
Metals	-3.35 *	T & C
Petroleum	-0.21	N
Agricultural raw materials	-4.05 *	T & C

* indicates rejection of the unit root hypothesis at the 5% significance level.

As Table 5 indicates, the null hypothesis of unit root is rejected at the 5 per cent significance level for the series of agricultural raw materials and metals. The tests suggest that these two series do not have unit roots, but rather are (trend) stationary. For all other series, failure to reject the unit root hypothesis suggests that a DS model is a reasonable univariate characterization.

Although the focus of the present paper is on volatility, the estimated time trends for the commodity price indices are reported in Table 6 below for the interested reader. If the DS model is chosen, the estimated time trend in the commodity price is the coefficient β in (4). If the TS model is chosen, because the unit root hypothesis is rejected, the estimated trend in prices is β in (2).

5.2   Commodity price volatility and exchange rate fluctuation

In many ways, commodity prices behave like asset prices. It has long been observed that asset returns tend to be platykurtic (i.e. they exhibit 'fat tails'). Volatility clustering is also common. That is, large changes tend to be followed by large changes, and small changes tend to be followed by small changes. Visual inspection of Figure 4 reveals this volatility clustering quite clearly. In this study, a univariate generalized autoregressive conditional heteroskedastic (GARCH) model is used to account for time-varying conditional variance structure of the errors in the detrended series. The hypothesis of interest is the extent to which changes in volatility for the US dollar and the euro-11 currencies are associated with changes in the volatility of world commodity prices (deflated).

5.2.1 GARCH models

In what follows, we assume that real primary commodity prices can be characterized by a deterministic or stochastic time trend model with an ARMA error process as in (2)-(3) or (4)-(5), depending on the outcome of the unit root tests in Table 5.[19] To capture the persistence in the conditional variance of prices, the innovations ε_t in (3) or (5) are assumed to follow a GARCH(p,q) process, as in Bollerslev (1986):

$$(7) \quad \varepsilon_t \big| I_{t-1} \sim N(0, h_t)$$

$$(8) \quad h_t = \delta + \sum_{i=1}^{p} \alpha_i \varepsilon_{t-i}^2 + \sum_{i=1}^{q} \beta_i h_{t-i}$$

where I_{t-1} is the information set through time t-1. The ε_t's are serially uncorrelated (but not stochastically independent because they are related through their second moments). According to (8), the variance of ε_t at time t depends on past news about volatility (the lagged ε_{t-i}^2 terms) and past

TABLE 6
UNIVARIATE MODELS WITH GARCH(1,1) ERROR PROCESSES
JANUARY 1957-DECEMBER 1998

	$\delta \times 10^3$	α	β	Log likelihood
Beverages	0.05	0.16	0.83	866.73
	(2.47)	(5.17)	(29.93)	
	$dlogy_t = -0.0025 + e_t$ where $(1-0.33L)\,e_t = \varepsilon_t$			
	(-0.95)		(7.17)	
Fertilizer	0.04	0.0002	0.99	714.12
	(5.54)	(1.35)	(396.32)	
	$dlogy_t = -0.001 + e_t$			
	(-0.26)			
Food	0.01	0.11	0.88	1137.86
	(1.89)	(5.09)	(47.16)	
	$dlogy_t = -0.002 + e_t$ where $e_t = (1+0.32L)\varepsilon_t$			
	(-1.53)		(6.66)	
Cereals	-0.03	0.11	0.88	1005.40
	(-3.24)	(5.69)	(42.19)	
	$dlogy_t = -0.0027 + e_t$ where $(1-0.33L)\,e_t = \varepsilon_t$			
	(-1.33)		(6.55)	
Metals	0.15	0.22	0.67	1017.43
	(6.52)	(4.90)	(13.68)	
	$logy_t = -4.23 - 0.0013*time + e_t$; $(1-0.96L)e_t = (1+0.35L)\varepsilon_t$			
	(-30.35)(-3.52)		(66.07)	(6.70)
Petroleum	4.07	0.21	0.25	564.01
	(4.72)	(3.25)	(1.54)	
	$dlogy_t = 0.0026 + e_t$; $(1+0.38L)e_t = (1+0.64L)\varepsilon_t$			
	(0.36)		(-2.14)	(4.52)
Agricultural raw materials	0.02	0.17	0.82	1158.65
	(2.41)	(4.85)	(27.04)	
	$Logy_t = -4.55 + e_t$; $(1-0.97L + 0.06L^{13})\,e_t = (1+0.24L)\varepsilon_t$			
	(-350.97)	(-0.57) (4.96)		(5.02)

Note: t-statistics in parentheses.

forecast variance (the h_{t-i} terms).[20] A sufficient condition for well-defined variance and covariance (including non-negative variance) is that all coefficients in equation (8) lie inside the unit circle. For example, in a GARCH(1,1) process, it requires $\alpha > 0$, $\beta > 0$, and $\alpha + \beta < 1$.

5.2.2 Empirical results

This sub-section first examines whether GARCH provides a good description for the behaviour of the six monthly commodity price series, ignoring the possible nominal exchange rate volatility effects on the conditional variance. Table 6 (opposite) presents the estimated coefficients for the univariate commodity price models with GARCH error processes. Note that the GARCH effects are significant for all commodities/indices. The strength of this significance is one indication of the appropriateness of the GARCH specification for the commodity price data. For most of the monthly commodity price series, GARCH(1,1) provides a sufficiently good fit, as it does typically for financial market variables. Diagnostic checks (correlograms and ARCH LM tests) confirm that there is no further serial correlation, nor ARCH effects, in the residuals after fitting the ARMA error processes reported in the Table. The sum $\alpha + \beta$ measures the persistence of volatility shocks. For most commodity series, this sum is very close to unity, implying that volatility shocks die out rather slowly.

To examine our main question in this paper, we now consider the potential importance of variables capturing the degree of nominal exchange rate volatility in the conditional variance equations estimated in Table 6. The conditional variance of ε_t in (8) is therefore replaced with:

$$(9) \qquad h_t = \delta + \delta_{US} * var\_us_t + \delta_{Euro} * var\_euro_t + \alpha \varepsilon_{t-1}^2 + \beta h_{t-1}$$

where *var_us* and *var_euro* are the indices for the dollar/SDR and euro-11 exchange rate volatility, respectively. We carry out a model selection exercise to decide whether either or both of the exchange rate volatility variables should be included. This is done by estimating the version of (9), as well as restricted version where $\delta_{US}=0$, $\delta_{Euro}=0$, or $\delta_{US}=\delta_{Euro}=0$. The latter restrictions, of course, indicate that nominal dollar exchange rate and/or euro-11 exchange rate volatilities have insignificant effects on world commodity price volatility.

The likelihood ratio tests used to examine these restrictions in the GARCH specification (9) are reported in Table 7. The restrictions $\delta_{US}=\delta_{Euro}=0$ are rejected at less than 5 per cent level for every commodity price indices. Column 3 and 4 report the likelihood ratio statistics for the restrictions $\delta_{US}=0$ and $\delta_{Euro}=0$, respectively. The results indicate that both the dollar exchange rate and intra-euro exchange rate fluctuations have significant impacts on the volatility of cereals, petroleum, and metals prices. On the other hand, the price volatilities for beverages, food and agricultural raw

materials depend only on the volatility of the US dollar/SDR exchange rate. Interestingly, the volatility of fertilizer prices does not depend on the volatility of the dollar/SDR exchange rate once the volatility of the euro-11 currencies is taken into account.

TABLE 7
THE LIKELIHOOD RATIO STATISTICS
FOR THE RESTRICTED AND UNRESTRICTED MODELS

	Ho: $\delta_{Euro}=0$		Ho: $\delta_{US}=0$		Ho: $\delta US=\delta_{Euro}=0$	
Beverage	0.60		13.32	*	14.12	*
Fertilizer	289.99	*	-155.97		242.98	*
Food	0.25		14.29		15.41	*
Cereals	6.65	*	25.96	*	33.93	*
Metals	64.71	*	10.30	*	65.92	*
Petroleum	527.06	*	206.53	*	535.72	*
Ag. raw materials	1.02		7.44	**	7.47	**

Note   * (**) indicates rejection of the null at the 5% (10%) significance level. The statistic has a chi-square distribution with the degree of freedom equal to the number of restrictions.

If either δ_{US} or δ_{Euro} are found to be insignificantly different from zero, the associated exchange rate volatility variable is dropped when estimating the final specification for (9) reported below. Based on the likelihood ratio tests in Table 7 (and in order to conserve space), a GARCH model with (i) both *var_us* and *var_euro*, (ii) *var_us* alone, or (iii) *var_euro* alone was reported in Table 8 for each commodity group. The characterizations of the means of the various price processes [in (2) or (3), plus (6)] remain similar to those in Table 6. The t-statistics on δ_{US} and/or δ_{Euro}—the coefficients on the dollar/SDR and euro exchange rate volatility measures, respectively—are statistically different from zero at 5 per cent significance level for every price series except for the price series of agricultural raw materials, which is only at the 10 per cent significance level. Moreover, the size and sign of the estimated δ_{US} and/or δ_{Euro} parameters indicate how the nominal exchange rate fluctuations affect the commodity price volatility. Since δ_{US} and δ_{Euro} are significantly positive, there is evidence that the volatility of nominal exchange rates have been *positively* associated with the volatility of real commodity prices.

For four commodity groups—cereals, fertilizer, metals and petroleum, the intra-euro exchange rate volatility has had statistically significant, positive effects on price volatility. Interestingly, the conditional volatility of the

TABLE 8
ACCOUNTING FOR EXCHANGE RATE FLUCTUATIONS
IN THE GARCH(1,1) MODEL

	$\delta \times 10^3$	δ_{US}	δ_{Euro}	α	β
Beverages	0.04	0.46		0.14	0.83
	(2.18)	(2.90)**		(4.74)	(27.17)
	$dlogy_t = -0.0031+e_t$ where $(1-0.32L)\,e_t=\varepsilon_t$				
	(-1.30)	(6.83)			
Fertilizer	-1.66E-03		0.46	2.30E-0	0.98
	(-1.54)		(32.04)**	(6.96)	(633.26)
	$dlogy_t = -0.002+e_t$				
	(-0.11)				
Food	0.01	0.18		0.05	0.91
	(2.59)	(3.15)**		(2.99)	(39.30)
	$dlogy_t = -0.0019+e_t$ where $e_t=(1+0.34L)\varepsilon_t$				
	(-1.39)	(7.91)			
Cereals	0.02	0.36	0.13	0.02	0.91
	(2.22)	(3.73)**	(3.82)**	(1.78)	(37.10)
	$dlogy_t = -0.0024+e_t$ where $(1-0.34L)\,e_t=\varepsilon_t$				
	(-1.24)	(8.04)			
Metals	2.72E-03	0.29	0.64	0.30	0.70
	(1.51)	(2.22)**	(9.67)**	(7.07)	(28.17)
	$logy_t=-4.25 - 0.001*time+e_t$; $(1-0.96L)e_t=(1+0.35L)\varepsilon_t$				
	(-84.58) (-4.59)	(102.64)	(6.81)		
Petroleum	1.508E-04	2.68	1.04	0.02	0.89
	(0.79)	(25.86)**	(7.13)**	(4.18)	(158.20)
	$dlogy_t=0.001 +e_t$; $(1+0.42L)e_t=(1+0.67L)\varepsilon_t$				
	(-2.24)	(-3.13)	(5.94)		
Raw materials	0.02	0.14		0.12	0.83
	(2.78)	(1.92)*		(4.09)	(24.38)
	$Logy_t=-4.55+e_t$; $(1-0.97L + 0.06L^{13})\,e_t=(1+0.24L)\varepsilon_t$				
	(-354.69)	(62.23) (-4.84)	(4.89)		

Note   *   indicates significant at the 10% significance level.
         **   indicates significant at the 5% significance level.

relative prices of fertilizer has been closely associated only with the volatility of the intra-euro nominal exchange rates. On the other hand, both

the US dollar and euro-11 exchange rate volatility has influenced the conditional volatility of the relative prices of cereals, metals, and petroleum products. The relative magnitude of influence from the two sources of exchange rate fluctuations, however, differs among cereals, metals and petroleum products.

# VI ASSESSING THE ECONOMIC SIGNIFICANCE OF EURO VOLATILITY ON COMMODITY PRICE VOLATILITY

Although the previous section indicates that exchange rate volatility measures have, in some cases, had *statistically* significant effects on commodity price volatility, one would like to assess their *economic* significance. For example, the EU's Common Agricultural Policy has operated a series of target and intervention prices aimed at the market for cereals. Thus, one might ask: what has been the impact of euro-11 exchange rate on the volatility of cereal prices? Are the effects large enough to warrant consideration in policy discussions over the future impacts of the euro? This section proposes a methodology for addressing this question.

Consider the GARCH model for cereal prices in Table 8, where both *var_us* and *var_euro* are statistically significant. First, we save the estimated conditional variance series h_t [see (9)] for the estimated GARCH model. Next, we calculate the extent to which *var_euro$_t$* contributes to h_t in each period, assuming the historical series for *var_us* remains unchanged.[21] The *initial* impact of var_euro$_t$ on conditional volatility h_t in (9) equals impulse$_t$=δ_{Euro}*var_euro$_t$. These impulses are then perpetrated through the AR(1) process in the conditional variance equation to determine the cumulative effect on the conditional variance. That is, cum$_t$ = =δ_{Euro}*var_euro$_t$. + β*cum$_{t-1}$.

Figure 5 shows the conditional variance series for cereals, along with the impulse and cum series, which show the contribution of var_euro$_t$ to the conditional variance in each period. Note that the contribution of var_euro to the conditional variance in cereal prices is higher in the 1970s through the early 1980s than it is in the mid-1980s and early 1990s when the euro-11 exchange rate was less volatile. Figure 6 shows the impulse and cumulative effects of *var_euro$_t$*, respectively, as percentages of the conditional variance in the same period. During the 1970s, for example, the

euro-11 volatility explains roughly 20 per cent if we use the cum_t/h_t measure. In the 1960s, this ratio ranged from 20-40 per cent.

FIGURE 5
CONDITIONAL VOLATILITY IN CEREALS GARCH MODEL
AND CONTRIBUTION OF VAR-EER

- - - - - Conditional Volatility
——— var_eer cumulative effect
——— var_eer impulse
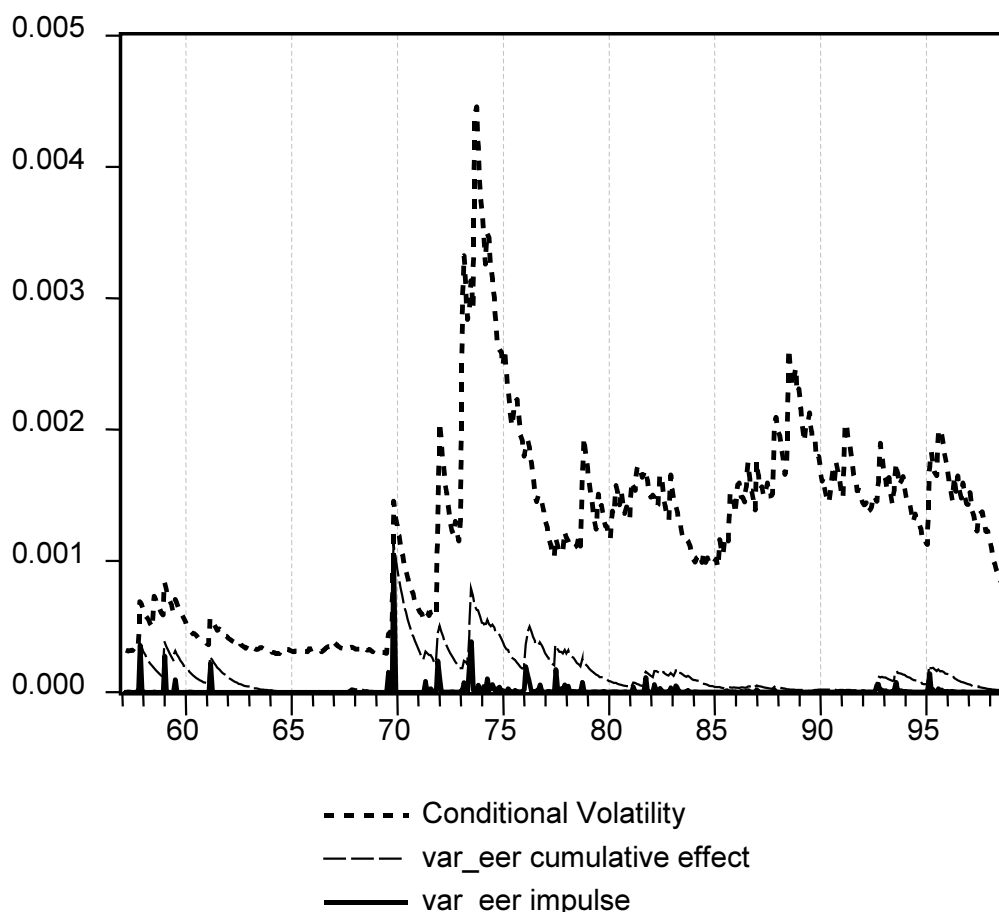
From this figure, one might conclude that reducing euro-11 volatility from the levels seen in the 1970s to zero might be expected to reduce the conditional volatility of cereals prices by roughly 20 per cent. If, on the other hand, one considers the modest volatility of the euro-11 under the EMS in the late 1980s and early 1990s, as the benchmark, there will be rather limited marginal gains in terms of reduced cereal price volatility when euro-11 is reduced to zero—the ultimate consequence of adopting a single currency in the euro-11 zone.

FIGURE 6
PER CENT CONTRIBUTIONS OF VAR-EER IMPULSE AND CUMULATIVE EFFECTS
TO CONDITIONAL VOLATILITY

— — — var_eer cumulative effect

———— var_eer impulse
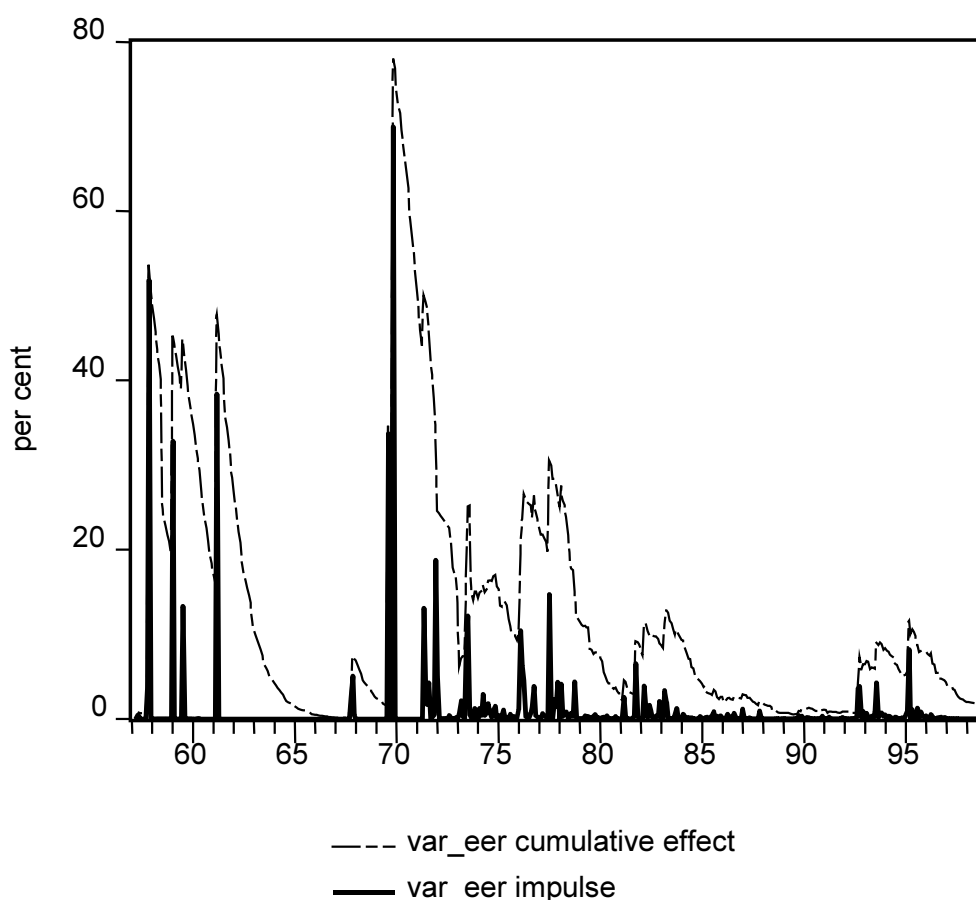
VII REVERSE CAUSALITY?

Some readers of the initial draft of this paper questioned whether there might be reverse causality in the volatility relations described in the previous section. In particular, is it possible that episodes of higher commodity price volatility led to greater exchange rate volatility? For example, it seems plausible that higher petroleum price volatility in the early 1970s induced a shift in exchange rate regimes toward greater flexibility.

To explore the reverse causality issue, we estimated univariate time series models for the logs of the dollar/SDR exchange rate and the euro-11 exchange rate allowing for GARCH error processes. As the exchange rate series are clearly integrated processes where, by definition, $\rho=1$, this

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amounts to estimating a DS model where y_t is now the nominal exchange rate (either the dollar/SDR rate or the euro-11 rate). The innovations in the error process in (5) are assumed to follow a GARCH process as in (7)-(8).

The seven commodity price volatility measures in Table 9 below were then added to the GARCH equation (8) to get:

(10) $h_t = \delta_1 + \delta_{2*} var\_petrol + \delta_{3*} var\_bev + ... + \delta_{7*} var\_metal + \alpha \varepsilon_{t-1}^2 + \beta h_{t-1}$.

Table 9 compares the values of the log likelihood functions of various specifications involving subsets of the seven commodity volatility measures. Specification 1 contains all 7 commodity volatility variables on

TABLE 9A
SIGNS OF COEFFICIENTS AND LOG-LIKELIHOOD FOR VARIOUS
SPECIFICATIONS USING DOLLAR (DLUS_SDR) EXCHANGE RATE AS THE
INDEPENDENT VARIABLE IN THE CONDITIONAL VARIANCE EQUATION

Specification	Var-petrol	Var-bev	Var-rawmat	Var-food	Var-cereal	Var-fert	Var-metal	Log-likelihood
1	-	-	+	-	-	-	-	1540
2	-	-	+	-	-	-		1539
3	-	+	+	+	+			1622
4	-	-	+	+				**1739**
5	-	-	+					1547
6	-	-						1540
7	-							1539
8								1708

TABLE 9B
SIGNS OF COEFFICIENTS AND LOG-LIKELIHOOD FOR VARIOUS
SPECIFICATIONS USING THE EURO-11 EXCHANGE RATE (DLNEER_DM)
AS THE INDEPENDENT VARIABLE IN THE CONDITIONAL VARIANCE EQUATION

Specification	Var-petrol	Var-bev	Var-rawmat	Var-food	Var-cereal	Var-fert	Var-metal	Log-likelihood
1	-	-	+	+	-	+	+	1668
2	-	-	+	+	-	+		1667
3	-	-	+	+	-			1672
4	-	-	+	+				1688
5	-	-	+					1681
6	-	-						1673
7	-							1702
8								1702

the right-hand side of the conditional variance equation. Since there are so many combinations for the right-hand side specifications, the table only shows the results of dropping one variable at a time from specification 1. Specification 8 is the simple GARCH model with no commodity volatility variables in the conditional variance equation for the nominal exchange rate series. The specifications including the commodity volatility measures are generally implausible. While individual coefficients are sometimes significant and other times insignificant, they often have a negative sign. This would imply that higher commodity volatility is associated with lower, not higher, exchange rate volatility. This is the opposite of the effect that sceptics conjectured at the beginning of this section. Moreover, the inclusion of commodity volatility measures in either the dollar or euro exchange rate equations invariably *reduced* the value of the log-likelihood function.[22] This suggests that the simple GARCH specification for the two nominal exchange rate series, ignoring commodity price volatility measures, provides the best univariate specification.

VIII CONCLUDING REMARKS

Our econometric analysis provides evidence that episodes of internal stability of exchange rates among the 11 euro countries during 1957-98 were associated with periods of lower real commodity price volatility. Judging from the relative magnitudes of the δ_{Euro} coefficients in the GARCH specifications, these stabilizing effects are particularly significant for fertilizer, metals and petroleum, and to a lesser degree, cereals. A reasonable inference, therefore, is that the establishment of the euro on 1 January 1999 should contribute to reduced volatility of world commodity prices, other things equal. This contribution will presumably be very small if we use the EMS period of euro-11 stability as our reference point. If , on the other hand, we use as a benchmark the 1970s when euro-11 exchange rate volatility was relative high, the effects of reducing euro-11 exchange rate volatility to zero (by adopting a single currency) appear correspondingly larger.

What are the mechanisms through which this stabilizing influence operates? We offer some speculative answers. The existence of EMU implies smaller monetary and fiscal shocks among the euro-11 countries, thereby reducing the fluctuations on world aggregate demand and indirectly

the demand for primary commodities, which are important intermediate inputs in global production.

This paper has focused on the impact that the emergence of the euro will have on the volatility of world commodity prices. The reverse question may also be worth pursing in future research: What impact are periodic commodity price shocks likely to have on the sustainability of the euro system? Are euro countries going to be more or less susceptible to commodity price shocks as members of the euro zone than they would have been under more flexible exchange rate arrangements? If countries give up the flexibility of responding to external commodity price shocks by adjusting their exchange rates, will the unemployment costs of adjusting to these shocks rise for countries in the euro zone? Some indication of the likely magnitude of such shocks for intra-euro zone adjustment could presumably be gleaned from the commodity trade (oil and non-oil) composition of exports and imports of euro zone and non-euro zone countries. It would be interesting to assess the implications of periodic commodity price shocks for fiscal transfers within the EU. There may also be direct implications for the intra-EU flows of commodities and financial support under the CAP.

The findings of this study are important for the on-going debate on the desirability of alternative exchange systems. One of the main reasons why economists care about the regime-dependent behaviour of real macroeconomic variables is due to the so-called *social inefficiency* associated with the floating exchange rates (Hallwood and MacDonald 1994). If there are nominal rigidities in goods prices, nominal exchange rate changes lead to changes in real exchange rates. Moreover, if goods with different underlying market structures adjust to exchange rates changes differently, the relative prices of two traded goods will also change. Such changes may result in inefficient allocation of goods and services.

NOTES

[1] The eleven countries in the European Monetary System that have adopted the euro are Austria, Belgium, Finland, France, Germany, Ireland, Italy, Luxembourg, Netherlands, Portugal, and Spain.

2 At the highly aggregated one-digit level, the SITC categories 0-4 include primary commodities, while categories 5-9 are intermediate and final manufactured goods.

3 See, e.g., Reinhart and Wickham (1994) and the references therein.

4 See, e.g., Beckerman and Jenkinson (1986), Hamilton (1983) and Reinhart (1991).

5 Borensztein and Reinhart (1994: 259) use an IMF index of the real exchange rate of the United States relative to other industrial countries. It is based on value added deflators in manufacturing.

6 We know of no literature that considered the impact of exchange rate uncertainty on exports and imports of *commodity* trade in isolation.

7 See De Grauwe (1994) and Szasz (1999) for institutional details and historical context.

8 The US dollar/SDR exchange rate, as well as the nominal exchange rates of the other 10 euro countries against the Deutsche mark, were obtained from the *International Financial Statistics* (line rf, the period average exchange rate).

9 'Beginning in 1 January 1991, the SDR valuation basket consists of the currencies of the five members having the largest exports of goods and services during the period 1985-89…' (IFS Feb. 1995: xi)

10 Given exchange rates as of 1 January 1996, the percentage (value) weights were 39, 21, 18, 11, and 11 per cent, respectively.

11 Note that reductions in the period-to-period per cent changes in the bilateral cross rate between Germany and France, say, may or may not produce a reduction in var_us.

12 Examining Figure 2, we noted that the sterling crisis in the late 1960s does not show up. This is because dollar/SDR rate was fixed during this period. The redefinition of the SDR in terms of a basket of currencies was introduced in July 1974 after the Bretton Woods fixed exchange rate commitment collapsed.

13 The weights are constructed by taking the average of the ratio of each individual country's exports plus imports over the total sum of trade in the euro zone over the period of 1969-98 when the trade data are available. The effective exchange rate calculated using GDP weights is quite similar to the index used here.

14 The MUV series is obtained from the IMF commodity price database. Some of these exports are flows among industrial countries; some are exports to developing countries. Hence, the MUV is not a perfect measure of the unit value of manufacturing exports to (only) developing countries. By the same token, some industrial countries export primary commodities. It is risky, therefore, to interpret the relative price of commodities (in terms of the MUV) as a measure of any country's terms of trade. Rather the relative price of primary commodities in terms of manufactures should be viewed as a key relative price in the world macroeconomy.

15 Using their re-constructed price index to estimate the time trend model, Grilli and Yang (1988) found evidence supporting the PS hypothesis. Cuddington and Urzúa (1989) introduced the unit root issue into the PS debate. Their statistical analysis

concluded that there was 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 1920/21 and again rejected the unit root hypothesis. Cuddington (1992) considered 26 individual commodities and found 13 of them to be trend stationary, whereas the other 13 were difference stationary.

[16] Since u_t is shown to be heteroskedastic later in the study, Phillips-Perron unit root test is more appropriate than the augmented Dickey-Fuller test.

[17] To obtain this equation, multiply (2) by (1-ρL) to get: $\log y_t = [(1-\rho)\alpha - \rho\beta] + \rho \log y_{t-1} + (1-\rho)\beta * time + \zeta_t$.

[18] See Enders (1995: Chapter 4) for detailed discussion of the GTS method.

[19] Because it is well known that unit root tests have low power, we tried different specifications of the trend equation when estimating the GARCH equations. Our final results do not appear to be particularly sensitive to the detrending methods used.

[20] The inclusion of lagged conditional variances distinguishes the GARCH model from Engle's (1982) ARCH model, and might be motivated as capturing some sort of adaptive learning mechanism. The GARCH specification is more parsimonious and entails fewer coefficient restrictions than the ARCH model.

[21] Given that the correlation between *var_us* and *var_euro* in the data is only 0.056, this seems unobjectionable.

[22] Note that it is theoretically possible for the value of the log-likelihood to fall when additional terms are added to the conditional variance equation. This occurs in situations when the coefficient on some of the exogenous determinants of the conditional variance are negative.

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