Metals Market Efficiency in Relation to Foreign Exchange and Financial Markets

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TABLE OF CONTENTS

SUMMARY

The paper reports on the results of an extension of the testing of the efficiency of metal trades at the London Metal Exchange (LME). Earlier work by the author pointed to the importance of financial market variables such as interest rates, exchange rates and inflation to primary commodity markets--because they affect the terms on which futures or forward traders will be prepared to hold title to the commodity--and modelled these relationships. This theoretical development allows an extension of the concept of futures or forward market efficiency to include the potential for arbitrage between commodity prices and interest rates and exchange rates. A second point of interest is a revisiting of the question of the size of the response of metal prices to exchange rate changes. Earlier empirical analysis by the author had produced results which indicated, contrary to theory, that the elasticity of primary commodity prices to exchange rate changes was greater than one.

The period of study covered LME forward prices over the period 1978-85 and included seven metals. Previous tests of market efficiency by the author, using weak-form efficiency tests, had concluded that there was no evidence of departures from efficiency in LME copper trading, but that evidence for lead, tin and zinc was mixed. The tin market appeared to exhibit a significant risk premium while in lead, and perhaps also in zinc, current prices could be predicted from previous prices. This paper extends the analysis to cover aluminum, nickel and silver and to include semi-strong form tests of efficiency of the LME in relation to foreign exchange and treasury bill markets.

In light of the tests carried out, none of the LME metals markets appear to conform to the efficient market paradigm. There is clear evidence of weak-form inefficiency through lagged dependence in the case of aluminum and possibly also copper, lead and zinc; there is somewhat weaker evidence of bias in aluminum, nickel, tin and zinc; there is evidence that the lagged forward premium (backwardation) has predictive value in aluminum and lead; and some evidence that exchange rate changes (and possibly also changes in inflation) are not transmitted efficiently to metals prices in the case of aluminum and possibly also lead, silver, tin and zinc.

Aluminum, nickel and zinc are seen as having negative bias over this sample period, implying a preponderance of short hedging (typically associated with producer sales), while tin provides evidence of positive bias. This latter result may well be the consequence of very heavy forward support activities by the International Tin Council during the period.

The exchange rate response parameters all take the predicted sign. The long-run elasticities are distributed fairly widely around the theoretical value of unity. Aluminum, nickel and zinc show much lower elasticities, which is consistent with producer pricing on a dollar basis (moreover, these three markets are dominated by a few large producers); while lead, silver and tin appear oversensitive to changes in the value of the dollar (i.e., elasticity

much larger than one). Copper alone has an estimated value close to the theoretical value of one.

The interest, inflation and activity innovation effects are relatively poorly defined, except in the case of copper. It is notable that the activity variable, industrial production, is insignificant in the tin equation; this again may demonstrate the effects of international support of tin prices.

Tests using differently constructed exchange rate indices yielded very different results. Theory suggests that in a multicommodity model the weights in the index will be complicated functions of all own and cross demand and supply elasticities and all market share parameters. Besides the GNP-weighted index used in the major part of the study other constructions of the exchange rate index included (i) the IMF's MERM index in which the weights reflect the weights of each currency in US overseas trade; and (ii) commodity-specific indices using weights proportional to, respectively, consumption of the commodity in each country, production of the commodity and production plus consumption. The commodity-specific indices gave the best fit with weighting by consumption shares giving superior results to weighting by either production shares or by an average of the two. The choice of the index had a significant effect on the size of the elasticity of the response of commodity prices to exchange rate changes. Use of the MERM index gave rise to high elasticities, reflecting the fact that the use of trade weights gives a particularly high impact to currencies (in particular to the Canadian dollar) which vary little in relation to the US dollar. Since the MERM index shows relatively less variation than the other indices, the variation that exists must take a higher weight in the estimated regression.

The size of the elasticity parameters should be regarded with some caution as they have been estimated in relationships which omit most of the fundamental determinants of price. This approach is reasonable as the study was mainly a test of market efficiency. In further work it may be useful to undertake the analysis using daily data which should strengthen the results on the question of the efficiency of these markets.

I. INTRODUCTION*

1. Many important primary commodities are traded on forward or futures exchanges. These commodities may be regarded as either physical or financial assets. Producers and consumers are concerned with physical properties of the commodity (location, purity, flavor, etc.), while futures traders are concerned with likely changes in the commodity price and, in particular, with movements in the basis (i.e. in the relative price of nearby and more distant futures). One should therefore expect that commodity prices will be affected not only by those factors which affect the supply and demand of the physical commodity ("fundamentals"), but also by financial market variables which affect the terms on which futures traders will be prepared to hold title to the commodity. But despite this elementary observation, scant attention has been paid to financial market variables in the commodity modeling literature.

2. There are three sets of financial market variables that will, in principle, be important in considering commodity price dynamics. These are exchange rates, interest rates and inflation rates. The relationship between exchange rates and commodity prices was discussed in an equilibrium model by Ridler and Yandle (1972). That analysis was extended to include inflation in Gilbert (1973). However, neither of those models apply directly to markets in which stocks are carried and where interest rate changes will also be important. This extension was made in Gilbert (1985). A rise in interest rates

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leads to a fall in asset prices as the given dividend stream is required to generate a higher yield. The same should be true in commodity markets where stock holders will be looking for a higher convenience yield. Gilbert (1985) also shows that if futures traders are risk neutral and if futures markets are unbiased commodity prices will respond only to the unanticipated components (the "innovations") of the exchange rate and inflation movements.

3. This allows an extension of the concept of futures market efficiency. A futures market is regarded as efficient if it is not possible to devise a trading rule based on a specified information set that will have positive expected profitability. In weak form efficiency tests, 1/ the information set is confined to the past price history of the asset price. Here, evidence of bias or of lagged dependence provides prima facie evidence of inefficiency, although there is no certainty that the implied trading rule would have covered transactions costs, or, in thin markets, that trades could have been made in sufficient volumes at the quoted prices. The results derived in Gilbert (1985) allow extension to semistrong form efficiency tests in which the information set is extended to include exchange rates, interest rates and inflation rates. Here one is asking whether it is possible to devise trading rules that involve simultaneous positions in commodity and forex markets, commodity and t-bill markets or commodity and physical markets which would have positive expected profitability.

4. The London metals markets provide a convenient and important testing ground for this theory. The London Metal Exchange (the LME) trades (or has

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1/ See Fama (1970).

recently traded) 1/ contracts in seven metals. These are silver (Ag), aluminum (Al), copper (Cu), nickel (Ni), lead (Pb), tin (Sn) and zinc (Zn). Silver, aluminum and copper are also traded on Comex in New York, and tin on the Kuala Lumpur Futures Exchange (Malaysia). 2/ The LME prices of these metals effectively constitute the free market prices, at least outside the United States, and in copper (and previously in tin) most trade outside the United States and the centrally planned economies takes place at or close to the LME price.

5. Gilbert (1986a) used weak form tests to examine the efficiency of LME trading in copper, lead, tin and zinc. He concluded that there was no evidence of departures from efficiency in the copper market where LME trading has been most active, but that the evidence for efficiency of the other three markets was more mixed. The tin market appeared to exhibit a significant risk premium, while in lead, and perhaps also in zinc, the previous price history allowed prediction of current movements. 3/ In this paper, we extend that analysis to cover all seven LME metals and consider semistrong form tests of the efficiency of the LME in relation to forex and t-bill markets.

6. These tests also allow us to throw more light on the issue of the exchange rate response of commodity prices. The models derived in Ridler and Yandle and Gilbert (1973, 1985) imply that the equilibrium exchange rate elasticity of dollar commodity prices to a change in the value of the dollar

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1/ Trading in the tin contract was suspended indefinitely in October 1985 after the International Tin Council default.

2/ Previously in Penang.

3/ This work was prompted by earlier discussions in Goss (1981, 1983) which were vitiated by serious econometric problems.

will fall within the (0,1) interval. In practice, however, dollar commodity prices appear to have been overresponsive to exchange rate changes during the first half of the eighties. The model derived by Van Duyne (1979) allows the possibility of overshooting, but does not affect the equilibrium elasticity. Gilbert (1987) provides some evidence for the view that developing country debt service obligations have forced supply shifts which would have the effect of increasing this elasticity. The results reported in this paper allow direct estimation of the exchange rate elasticity for the LME metals from the market price reactions to exchange rate changes.

7. The plan of the paper is as follows: in Section 2 we review the theory relating exchange rate changes to commodity prices; in Section 3 we discuss the data used for the efficiency tests reported in Section 5; Section 4 is devoted to econometric issues; in Section 6 we consider the implications of alternative weighting procedures in the construction of exchange rate indices; and Section 7 contains brief conclusions.

II. THE IMPACT OF EXCHANGE RATE CHANGES ON COMMODITY PRICES

8. We first derive the Ridler and Yandle (1972) result, as generalized in Gilbert (1973). Write the price of commodity as p. Suppose country i's dollar exchange rate is x (domestic units per dollar). There are n countries with the United States taking index 1, so that $x_1 = 1$. Denote consumption of the commodity in country i by $C_i = C^i\,[x_i p/q_i]$ where q_i is the appropriate price deflator in country i. Similarly, let production of the commodity be $Q_i = Q^i[x_i p/q_i]$. Write the country i demand and supply elasticities as e_i and ε_i, respectively, with e_i defined as positive. Let country i's share in world consumption of the commodity be $w_i = C_i/\Sigma_k C_k$ and its share in world production be $\omega_i = Q_i/\Sigma_k Q_k$. Define e and ε as the weighted average demand and supply elasticities respectively (i.e., $\Sigma_i w_i e_i$ and $\Sigma_i \omega_i \varepsilon_i$). Market clearing, over a sufficiently long period for stock changes to be negligible, requires

$$(1) \qquad \sum_{i=1}^{n} C^i[x_i p/q_i] = \sum_{i=1}^{n} Q^i[x_i p/q_i]$$

The Ridler-Yandle result is obtained by total differentiation of (1) followed by approximation of derivatives by first differences. This gives

$$(2) \qquad \Delta \ln p = \sum_{i=1}^{n} v_i[\Delta \ln x_i - \Delta \ln q_i]$$

[Ridler and Yandle (1972), Gilbert (1973)]. Ridler and Yandle showed that the weights $[v_i]$ are given by

(3) $$v_i = \frac{w_i e_i + \omega_i \varepsilon_i}{e + \varepsilon}$$

Equations (2) and (3) may be simplified by defining (commodity specific) exchange rate and relative price indices X and Q as

(4) $$\ln X = \sum_{i=2}^{n} v_i \ln x_i$$

and

(5) $$\ln Q = \sum_{i=2}^{n} v_i \ln q_i$$

Then (2) becomes

(6) $$\Delta \ln p = \Delta \ln X - \Delta \ln Q$$

9. Note that the summation in the exchange rate and price indices X and Π is over countries 2 to n, and thus excludes the United States. This implies that the weights in these indices sum to $1-v_1$. Consider a $100\theta\%$ general appreciation of the dollar so that $\Delta \ln x_i = \theta$ for $i=2,\ldots,n$. Suppose that there are no associated changes in the levels of the price deflators. Then (6) implies

(7) $$\Delta \ln p = [1-v_1]^{\theta}$$

and we obtain the bounding relationship $0 \leqq \Delta \ln p \leqq 1$. The commodity price falls by $100(1-v_1)\%$ of the dollar appreciation. Typical values, implied by (7), for the elasticity of commodity prices with respect to changes in the value of the dollar, would be of the order of 0.6-0.8.

10. Alternatively, suppose indices X^* and Q^* are defined with weights $v_i^* = v_i/[1-v_1]$, which sum to unity over countries 2...n. Then (6) becomes

$$(8) \qquad \Delta \ln p = -(1-v_1)[\Delta \ln X^* - \Delta \ln Q^*]$$

which implies the same result. However, the commodity price has unit elasticity with respect to the Ridler-Yandle indices (4) and (5) defined with weights summing (over countries 2,...,n) to 1-v , but elasticity $1-v_1$ with respect to the indices defined in (8) with weights summing to unity. This implies a need for some caution in comparing the results of different studies. In this paper we define the exchange rate indices as in (4) with the implication that the commodity price should exhibit a unit elasticity with respect to changes in these indices.

11. We need to consider two extensions to the simple Ridler-Yandle model developed above. The first is to a multicommodity world. The Ridler-Yandle model considers the effect of a dollar devaluation on the price of (say) copper with the price of all other commodities held constant. But it would clearly be inconsistent to apply this result to copper while holding the price of aluminum constant and, at the same time, to apply it to aluminum while holding the price of copper constant. Gilbert (1987) shows that the Ridler-Yandle result extends to a multicommodity world, and in particular that, given gross substitutability across all traded goods, commodity price elasticities with respect to exchange rate changes remain within the unit interval. However, the weights $[v_i]$ implied by the exchange rate indices now depend upon all own and cross supply and demand elasticities for all countries. This

suggests that there may be little point in attempting to construct commodity specific exchange rate indices. We examine this suggestion in section 6.

12. The second extension is to a model in which stocks of the commodity are held by rational risk neutral agents. This model is discussed in Gilbert (1985) where it is shown that the innovation (i.e. the unanticipated change) in the commodity price is related to the unanticipated change in the Ridler-Yandle exchange rate index X, the price index, and to a similarly weighted index of interest rates. Assumption of covered interest parity allows the interest rate index to be replaced by a single interest rate, and it is natural in a model in which the US dollar is numeraire to choose the US rate. The commodity price change then depends on current and future expected changes in these indices. Simple results are obtained by making appropriate constancy assumptions. A result analogous to that developed by Ridler and Yandle follows if changes in exchange rates, interest rates and price levels are unanticipated. This gives

$$(9) \qquad \nabla \ln p = -(\nabla \ln X - \nabla \ln Q) - \alpha \nabla r$$

(The coefficient will depend on the expected stockholding period and may, therefore, not be constant.) It is, however, more plausible to suppose that the rates of inflation rather than price levels are stationary. If agents regard all changes in inflation rates π_i as unanticipated then (9) is replaced by

$$(10) \qquad \nabla \ln p = -\nabla \ln X = \alpha(\nabla r - \nabla \bar{\pi})$$

where

$$(11) \qquad \Pi = \sum_{i=1}^{n} v_i \pi_i = \sum_{i=1}^{n} v_i \Delta \ln q_i$$

More generally, one might model the forcing variables lnX, r and Π as (nth order) autoregressive processes

$$(12) \qquad \lambda^z(L) z_t = \mu^z + \nu_t^z \qquad (z = \ln X,\ r,\ \Pi)$$

where L is the lag operator $Lz_t = z_{t-1}$ and $\lambda_0^z = 1\ \forall z$. In this situation, an innovation in any variable z will result in an expectation of further changes in z over the future. The restrictions in (11) are then relaxed and one obtains

$$(13) \qquad \nabla \ln p = -\alpha_1 \nabla \ln X - \alpha_2 \nabla r + \alpha_3 \nabla \Pi$$

However, if the period over which stocks are expected to be held continuously is sufficiently long, the α coefficients in (13) may be simply expressed in terms of the λ coefficients in (12):

$$(14) \qquad \nabla \ln p = -\Lambda^{\ln X} \nabla \ln X - \alpha(\Lambda^r \nabla r - \Lambda^{\Pi} \nabla \Pi)$$

where

$$\Lambda^z = (1 - \sum_{i=1}^{n} \lambda_i^z)^{-1} \qquad (z = \ln X,\ r,\ \Pi)$$

The expressions $\Lambda^z \nabla z$ denote the innovations scaled by the appropriate long run multiple. To avoid unpleasant notation we shall henceforth absorb this scaling factor into the innovation operator ∇ which allows us to revert to (10). Note that the unit elasticity property is restored now that the exchange rate innovation has been appropriately scaled.

13. Exchange rates, interest rates and prices together explain only a small proportion of commodity price movements. In a full commodity market model one would need to model the shifts in the supply and demand functions. In this paper, our concern is more limited and so attention is confined to the financial market variables. However, inclusion of a demand shift (A, activity) variable (again in scaled innovation form) increases the power of the efficiency tests. We therefore augment (10) as

$$\nabla \ln p = - \nabla \ln X - \alpha_1(\nabla r - \nabla \Pi) + \alpha_2 \nabla \ln A \tag{15}$$

Equation (15) forms our null hypothesis. Departures from (15) imply the possibility of profitable trading rules.

III. DATA AND VARIABLE CONSTRUCTION

14. The LME trades metal for "prompt" (i.e., spot) and 90-day (3-month) delivery. 1/ In this study we use a sample of 32 nonoverlapping end month LME prices covering the period 1978-85. The selected months are January, April, July and October--this avoids possible year-end problems. The metal price innovation is measured as

$$(16) \qquad \nabla \ln p_t = \ln p_t - \ln p^f_{t|t-1}$$

where $p^f_{t|t-1}$ is the forward price at date t-1 for delivery at date t. 2/

15. The aluminum and nickel contracts are relatively recent--taking into account the need for a lagged forward price, data on aluminum is only available from July 1979 and on nickel from April 1980. The tin contract was suspended on October 24 1985--we take the price on that date as the end October price. A longer sample would have been available if we had been prepared to use pre-1978 data. However, forward exchange rate data, required in the construction of the exchange rate innovations, is not easily available before 1978; and, in any case, there is no strong reason to expect any departures from market efficiency to be constant over long periods of time.

1/ Strictly, the 90-day contracts are forward and not futures contracts [see Gilbert (1986a)]--they are made with an LME ring-dealing member, who acts as principal, and not with an exchange clearing house they are canceled by making an offsetting contract with the same delivery date with the same broker; and they are not marked to market during the contract period.

2/ LME prices are quoted in sterling. The spot price p is converted into dollars at the spot sterling exchange rate, and the forward price at the three-month rate.

16. An alternative method of increasing the sample size would have been to use overlapping observations--either end-month observations [see Gilbert (1986a)] or daily observations. The use of overlapping observations in market efficiency studies gives rise to inferential problems in the econometrics which have been authoritatively analyzed by Hansen and Hodrick (1980). The major difficulty with the Hansen and Hodrick procedure, noted in Gilbert (1986a), is that it requires homoscedasticity of the innovation variance. With financial market data, this assumption is neither theoretically reasonable nor empirically realistic. 1/ A major advantage deriving from the use of nonoverlapping data is the possibility of explicitly modeling the innovation variance process. 2/

17. The exchange rate data consists of end month observations on OECD exchange rates against the US dollar. 3/ One would wish to define the exchange rate innovation for country j analogously to the price innovation (16):

$$(17) \qquad \nabla \ln x_{jt} = \ln x_{jt} - \ln x^{f}_{j,t|t-1}$$

1/ See Mandelbrot (1963, 1966, 1967) and Press (1967) for theoretical derivation of the distribution of asset price changes; and Engle and Bollerslev (1986) for a survey of financial market applications of the ARCH heteroscedasticity model.

2/ In principle it should be possible to extend the Hansen and Hodrick (1980) procedure to allow an ARCH error variance process. This is, however, beyond the scope of the current paper.

3/ Australia, Austria, Belgium-Luxembourg, Canada, Denmark, Finland, France, Federal Republic of Germany, Ireland, Italy, Japan, The Netherlands, Norway, Spain, Switzerland, United Kingdom. Source: spot rates--IMF, International Financial Statistics; forward rates (see text)--Financial Times. I am grateful to Sompheap Sem for assembling this data.

where $x^f_{j,t|t-1}$ is the 90-day forward exchange rate for currency j at date t-1 for delivery at date t. However, a complete set of forward exchange rate data is only easily available over this sample for a subset S of major currencies (those of Canada, France, Federal Republic of Germany, Japan, Switzerland and the United Kingdom). For those currencies, the exchange rate innovations were defined as in (17). For the remaining currencies forward rates were estimated as weighted averages of the forward discounts of the currencies in the subset S where the weights were obtained by regressing the actual exchange rate changes on the exchange rate changes in subset S. Thus

$$(18) \qquad \ln(x^f_{j,t|t-1}/x_{j,t-1}) = \sum_{k\varepsilon S} \hat{\gamma}_k (x^f_{k,t|t-1}/x_{k,t-1}) \qquad (j\notin S)$$

where the weights $\hat{\gamma}_k$ are obtained from OLS estimation of the regression

$$(19) \qquad \Delta \ln x_{jt} = \sum_{k\varepsilon S} \gamma_k \Delta \ln x_{kt} + u_{jt}$$

The exchange rate index X was then calculated using (4) as

$$(20) \qquad \nabla \ln X_t = \sum_{j=2}^{n} \nu_j \nabla \ln x_{jt}$$

where the weights $[v_j]$ are equal to the share of country j in total dollar GDP in 1978.

18. It is also possible in principle to calculate interest rate innovations from market data. Let r_t be the one period (three-month) interest rate at date t, $r^f_{t+1|t}$ be the implicit forward rate, and ${}_2r_t$ be the two-period (six-month) rate. Then, in the absence of any risk premium,

$$(21) \qquad r^{f}_{t+1|t} = \frac{(1 + {}_{2}r_{t})^{2}}{(1 + r_{t})} - 1$$

In practice, we found that these implicit forward rates for US t-bills were not informative in the sense that, knowing the current spot rate, knowledge of the implicit forward rate did not improve predictions of the next period's spot rate. This is in line with the results reported by Shiller et al. (1983) who conclude (p. 215) that "the simple theory that the slope of the term structure can be used to forecast future changes in the interest rate seems worthless."

19. We therefore calculated interest rate innovations from an estimated autoregression [as (12)]. 1/ This is the same procedure that was adopted for the inflation and industrial production variables which were computed using the same weights $[v_j]$ as in the exchange rate index (20),

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1/ The autoregressions were estimated on the aggregated inflation and industrial production indices. In Gilbert (1987) we computed industrial production innovations as recursive residuals. This procedure is preferable to the use of OLS residuals since one is not obliged to suppose that agents have access to equation estimates based on future sample information. However, degrees of freedom constraints prevented us from applying that procedure in this study.

but running over the complete set of countries including the United States. 1/

In each case, a first-order autoregression appeared sufficient. 2/

1/ The industrial production index (measured as an average over the quarter previous to the month in question--i.e., the January 1978 observation relates is the average for the fourth quarter of 1977) covered the same countries as entered the exchange rate index excluding Luxembourg and Switzerland. The inflation index was calculated in terms of producer prices (current dated) over the same set of countries. The price indices were:

Canada: producer prices--electrical machinery;
United States: producer prices--all goods;
Japan: producer prices--manufactured goods (total);
Australia: wholesale prices--machinery and equipment;
Austria: wholesale prices--transport equipment;
Belgium: wholesale prices--manufactured goods;
Denmark: wholesale prices--machinery and transport equipment;
Finland: producer prices--investment goods;
France: wholesale prices-semimanufactured goods;
Federal Republic of Germany: producer prices--manufacturing;
Ireland: wholesale prices--manufactures;
Italy: wholesale prices--finished investment goods;
Netherlands: producer prices--nonelectrical machinery;
Norway: wholesale prices--investment goods;
Spain: wholesale prices--metals and metal products;
Sweden: producer prices--(home market)--manufactured goods;
Switzerland: wholesale prices--metals and metal products;
United Kingdom: wholesale prices--total (excluding food, drink and tobacco).

Source (for both industrial production and producer price indices): OECD, Historical Statistics 1960-1975 and Main Economic Indicators, various issues. I am grateful to Sompheap Sem for assembling this data.

2/ The estimated equations are (t-statistics in parentheses)

$$\Delta \ln A_t = \underset{(1.22)}{0.0027} + \underset{(3.65)}{0.56} \Delta \ln A_{t-1} \qquad R^2 = 0.31 \quad DW = 2.03$$

$$rt = \underset{(1.96)}{0.0049} + \underset{(7.59)}{0.79} r_{t-1} \qquad R^2 = 0.66 \quad DW = 1.86$$

$$\Pi_t = \underset{(1.62)}{0.0026} + \underset{(5.43)}{0.73} \Pi_{t-1} \qquad R^2 = 0.50 \quad DW = 1.75$$

The equations were also estimated by OLS over the sample 1978q1-1985q4.

IV. ECONOMETRIC STRATEGY

20. We noted in the previous section that financial market data is prone to exhibit heteroscedasticity. If this is not taken into account, invalid inferences may be drawn. We adopt two procedures in the estimations reported in this paper. The first is to compute, for the OLS regressions, heteroscedasticity-consistent standard errors using the White (1980) procedure. The OLS estimator $\hat{\beta}$ in the regression $y = X\beta + u$ with $Euu' = \Omega$ has variance $(X'X)^{-1}(X'\Omega X)(X'X)^{-1}$. White showed that, in the case in which Ω is diagonal, the ijth element w_{ij} of the matrix $W = X'\Omega X$ may be consistently estimated as

$$\bar{\hat{w}}_{ij} = T^{-1} \sum_{t=1}^{T} x_{it} u_{jt} u_{jt} x_{jt} \tag{22}$$

21. The second approach we adopt is to suppose that the error u follows an ARCH (autoregressive conditional heteroscedastic) process. 1/ Confining attention to first order ARCH processes, one posits

$$E_{t-1} u_t = 0$$

(23)

and $$E_{t-1} \sigma_t^2 = \sigma^2 + \phi^2 u_{t-1}^2$$

1/ The ARCH model was proposed by Engle (1982). Developments and applications are surveyed in Engle et al. (1985) and Engle and Bollerslev (1986).

In this process the errors are uncorrelated but not independent, since they are related through their second moments. The conditional and unconditional error variances differ, the conditional variance being given as (23) and the unconditional variance as $(1-\phi^2)^{-1}\sigma^2$.

22. Engle (1982) discusses maximum likelihood estimation of ARCH models. Alternatively, one may consider a two stage (Aitken) estimator where the ARCH parameter(s) is estimated from an initial OLS regression and used to make a subsequent GLS correction. Since the information matrix is block diagonal [Engle (1982, p. 997)], this Aitken estimator will be asymptotically efficient [Cox and Hinckley (1974, p. 308)]. Thus, if the OLS residuals are denoted as $[e_t]$ one estimates ϕ from the regression

$$(24) \qquad e_t^2 = \sigma^2 + \phi^2 e^2_{t-1} + v_t$$

and then scales the observations by the factor $(\hat{\sigma}^2 + \hat{\phi}^2 e^2_{t-1})^{1/2}$ in the second stage regression.

23. An advantage of the data set exploited in this paper is that it relates to identical contracts on all seven LME metals. It seems plausible, and indeed is the case, that the price innovations are correlated across these metals. A demand surprise, for example, that results in an upward revision in the price of copper will have the same effect on the price of aluminum. This suggests that efficiency gains may be obtained (in those regressions in which the regressor sets differ across metals) by the use of the SUR (Seemingly Unrelated Regression) estimator. Furthermore, SUR estimation has the additional advantage that it allows straightforward imposition of cross equation restrictions.

24. Consider an m equation model, and write the ith equation as $y_i = X_i\beta_i + u_i$. Writing $E(u_{it}u_{it}') = \Sigma$, and supposing $E(u_{is}u_{it}') = 0$ for $s \neq t$, the SUR estimator of the stacked coefficient vector $\beta = (\beta_1',\ldots,\beta_m')'$ is

$$(25) \qquad \hat{\beta} = [X'(\Sigma^{-1} O I_T)X]^{-1}X'(\Sigma^{-1} O I_T)^{-1}y$$

where X and y are the conformably stacked data matrices and I_T is the identity matrix of order T (the sample size). 1/ In view of the concern expressed above about heteroscedasticity, the question arises as to whether and how (25) may be generalized to allow for heteroscedasticity.

25. The answer to this question is somewhat mixed. There does not appear to be any straightforward way of generalizing the White (1980) procedure to systems estimation since the White procedure allows the stacked error variance matrix $\Sigma O I$ to be unstructured while the SUR estimator requires a matrix which may, in an appropriate normalization, be taken to be constant. However, the ARCH procedure, which does impose a specific structure on the error variance matrix, generalizes fairly easily to systems estimation.

1/ Because of "missing observations" for the aluminum and nickel price series, the SUR procedure was slightly modified. The covariance matrix Σ was estimated from the subsample for which complete information was available allowing β to be estimated from the entire sample with the "missing" entries for y and the corresponding entries for X set to zero. A two-stage procedure was employed with the initial estimate of Σ taken from the OLS estimates (estimated over the entire sample for which data on each metal was available).

26. The error variance matrix Σ_t contains $\frac{1}{2}m(m-1)$ distinct elements, and in principle one might allow each of these to follow an ARCH process. This gives rise to a problem with regard to ensuring that Σ_t remains positive definite at each date; but, in any case, is ruled out in our sample by degrees of freedom considerations. An alternative, which only expends 2m degrees of freedom, is to suppose that Σ_t exhibits a constant correlation structure. Write the ijth element of Σ_t as σ_{ijt} and suppose

$$\sigma_{iit} = h_{it}^2 = \sigma_{ii} + \phi_i^2 \sigma_{ii,t-1}$$

(26) and

$$\sigma_{ijt} = h_{it} h_{jt} \rho_{ij} \quad (i \neq j)$$

which is a natural generalization of (23). Then, defining $\rho_{ii} = 1$, $R = (\rho_{ij})$ is the constant correlation matrix. Now write $h_t = (h_{1t}, \ldots, j_{mt})'$ and $H_t = \text{diag}(h_t)$. This allows us to express the SUR estimator of β as

$$(27) \qquad \hat{\beta} = [X^{*\prime}(R^{-1} \otimes I_T) X^*]^{-1} X^{*\prime}(R^{-1} \otimes I_T) y^*$$

where

$$X^* = (H^{-\frac{1}{2}} \otimes I_m) X \text{ and } y^* = (H^{-\frac{1}{2}} \otimes I_m) y \text{ implying } X_j^* = H^{-\frac{1}{2}} X_j \text{ and } Y_j^* = H^{-\frac{1}{2}} y_j.$$

The estimation procedure is therefore the same as in the ARCH-OLS model (24): one can obtain consistent β coefficients from the SUR estimator (25) which allows one to estimate the scaling matrix H from (24) and to proceed to estimate the ARCH-SUR estimator (27). 1/

1/ As with the SUR, estimator, R was estimated from the sample for which complete data were available, the initial estimate being taken from the ARCH-OLS estimates.

27. In the results reported in Section 5 we employ four estimators: (i) the OLS estimator with heteroscedasticity consistent standard errors; (ii) the ARCH-OLS estimator defined by (24); (iii) the SUR estimator; and (iv) the ARCH-SUR estimator defined by (24), (25) and (27).

V. RESULTS

28. In Table 1 we report estimates of the metals' price adjustment equations (15) but without imposition of either the unit restriction on the exchange rate innovation coefficient or the equality restriction on the interest rate and the inflation rate coefficients. The generalization is therefore to

$$(28) \qquad \nabla \ln p_{jt} = -\alpha_{j1}\nabla \ln X_t - \alpha_{j2}\nabla r_t + \alpha_{j3}\nabla \Pi_t + \alpha_{j4}\nabla \ln A_t + \varepsilon_{jt}$$

$$(j=Ag,\ldots,Zn)$$

(∇r, $\nabla \Pi$ and ∇A are scaled innovations--see Section 3).

29. The coefficient estimates are broadly in line with the discussion of Section 2--the exchange rate innovation enters with the predicted negative coefficient in all estimates and appears to be distributed around unity. The interest rate coefficients, which are very poorly determined, are more often negative than positive, and are roughly half the size of the inflation innovation coefficients. The activity coefficients, which are much better determined, are as expected. Substantial ARCH effects are only apparent for silver and zinc--in the case of silver this is almost entirely due to the spectacular volatility exhibited during the 1979-80 Bunker Hunt corner. Overall, the fit is only modest, but this is perhaps to be expected in relationships which omit most of the fundamental determinants of prices.

30. It would be possible to impose restrictions on the Table 1 estimates ($\alpha_1 = 1$, $\alpha_2 = \alpha_3$). It is, however, preferable to look for departures from

efficiency prior to the imposition of restrictions which are predicated on efficiency obtaining. If the markets are efficient the residuals from (28) should be serially independent, and so one test is obtained by consideration of the Durbin-Watson (DW) and Lagrange Multiplier (LM) 1/ tests in Table 1. While none of these give definitive answers, the DW statistics for the aluminum and lead equations are worryingly low.

31. In Table 2 we report the results of adding one variable at a time to (28). This "simple to general" approach 2/ is dangerous since a variable may appear insignificant in this test when it would be jointly significant with a second variable; and because a variable may appear significant as the consequence of correlation with a second variable, in the presence of which it would be insignificant. However, degrees of freedom constraints prevent our adopting the preferable full 'general to simple' approach. Nevertheless, it should be borne in mind that the results of the tests reported in Table 2 are, at best, indicative and not decisive.

32. There are two sets of tests and three sets of additional regressors. The first set of additional regressors comprises the intercept and lagged dependent variable. Addition of these variables corresponds to the weak form efficiency tests reported in Gilbert (1986a), and asks (intercept) whether there is a consistent bias in the metal forward price and (lagged dependent variable) whether the price change is forecastable from its previous history. The second set of additional regressors comprises the lagged price innovations. Here one asks whether the response to the financial market

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1/ Godfrey (1978).

2/ See Gilbert (1986b).

innovations is noninstantaneous. Finally we consider addition of noninnovational variables. This final set of tests is not altogether distinct from those resulting from the addition of the lagged innovations since any stationary variable may be expressed as a sum of lagged innovations (the so called moving average representation). The first set of tests for each set of additional regressors gives the t-statistics from the addition of each variable to the OLS and ARCH-OLS regressions; the second (likelihood ratio test) considers exclusion of the entire vector of variables from the corresponding system (SUR and ARCH-SUR) estimates.

33. Inspection of the results indicates clear evidence of weak form inefficiency through lagged dependence (aluminum; possibly also copper, lead and zinc); somewhat weaker evidence of bias (possibly aluminum, nickel, tin and zinc); evidence that the lagged forward premium (backwardation) is informative (aluminum and lead); and evidence that exchange rate changes (possibly also changes in the rate of inflation) are not transmitted efficiently to metals prices (aluminum; possibly also silver, lead, tin and zinc).

34. Tables 3-6 report the results of a data-based attempt to isolate the locations of these inefficiencies. We further generalized (28) to

$$(29) \qquad \nabla \ln p_{jt} = \beta_{j0} + \beta_{j1}\nabla \ln p_{j,t-1} - \beta_{j2}(\ln p^{f}_{j,t-1} - \ln p_{j,t-1})$$

$$- \beta_{j3}\nabla \ln X_t - \beta_{j4}\nabla \ln X_{t-1} - \beta_{j5}\nabla r_t + \beta_{j6}\nabla \Pi_t$$

$$+ \beta_{j7}\nabla \ln A_t + \varepsilon_{jt} \qquad (j=Ag,\ldots,Zn)$$

which nests (28). The choice of these additional regressors was motivated by the results of the tests reported in Table 2. For each metal, we then attempted to find simplifications of (29) based on the estimated equations (not reported). The simplifications took the form of eliminating variables which were estimated as (insignificantly) taking the opposite sign from that predicted by the discussion of section 2 together with those variables which were associated with very low t-statistics. Also, noting that

$$\nabla \ln p_{jt} = \ln p_{jt} - \ln p^{f}_{j,t-1} \tag{30}$$

$$= \Delta \ln p_{jt} - (\ln p^{f}_{j,t-1} - \ln p_{j,t-1})$$

the parameter β_{j2} was restricted either to be zero or unity, the latter value implying that the dependent variable of the regression becomes $\Delta \ln p_{jt}$. (The unit restriction was appropriate only in the case of lead). In total, this amounts to 23 restrictions. Tables 3, 4, 5 and 6 report respectively the OLS, ARCH-OLS, SUR and ARCH-SUR estimates of these restricted equations.

35. An essential part of any simplification exercise is to test the proposed simplifications against the unrestricted hypotheses [Gilbert (1986b)]. In this case, this implies testing against the equations defined by (29). Tables 3 and 4 give the standard F tests taking each equation independently; and Tables 5 and 6 report the likelihood ratio tests for the

entire set of restrictions. 1/ All these tests accepted the proposed simplifications.

36. Nevertheless, other simplifications would also be possible which might result in somewhat different interpretations. It is doubtful that encompassing tests of rival simplifications would be very powerful on this data given the relatively low degree of explanation obtained. The results reported in these four tables are therefore more conjectural than those reported in Tables 1 and 2 since they are conditional on the congruency of the proposed simplifications. Despite this, these results do offer one plausible interpretation of the departures from efficiency uncovered in Table 2.

37. The ARCH error variance process is only evident for silver, and it is thus for that metal that the ARCH and conventional estimates differ most. However, the relatively high residual correlations between silver and both copper and nickel generalize the effects of the ARCH transformation to these equations in the ARCH-SUR estimates.

38. Aluminum, nickel and zinc are seen as having a negative bias of the order of 1½-2% over this sample, implying a preponderance of short hedging (typically associated with producer sales), while tin provides evidence of a

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1/ The "missing observations" for the aluminum and nickel prices complicate the likelihood calculations. The log-likelihood L was computed as $L = \Sigma L_t$ where

$$L_t = -\tfrac{1}{2}n_t \ln(2\pi) - \tfrac{1}{2}\ln|\Sigma_t| - \tfrac{1}{2}e_t'\Sigma_t^{-1}e_t$$

where e_t is the vector of residuals for period t and n_t is the number of metals for which data are available. Where data were unavailable, the associated rows and columns of Σ_t were eliminated.

positive bias of the same order. This may be a consequence of the very heavy forward support activities operated by the International Tin Council (ITC) over this period. 1/ The lagged price innovations appear as significant in five of the estimated equations, confirming the results reported in Table 2, the effect being particularly acute in the aluminum relationship. 2/

39. The exchange rate terms all take the predicted sign, although the estimated coefficient for aluminum is very sensitive to the presence of the ARCH adjustment. We were not very successful in isolating the lagged exchange rate effects implied by Table 2. The long run elasticities are distributed fairly widely around the theoretical value of unity: aluminum, nickel and zinc all show much lower elasticities which is consistent with producer pricing on a dollar basis (these three markets are dominated by powerful producers); while silver, lead and tin sample which ended in 1978 found significant lagged price change appear oversensitive to changes in the value of the dollar. Only copper has an estimated elasticity close to the theoretical value of unity.

40. The interest, inflation and activity innovation effects remain relatively poorly defined, except in the copper relationship. It is notable that the activity variable is absent from the tin equation; this again may demonstrate the effects of the ITC support operation.

1/ See Anderson and Gilbert (1986).

2/ Gilbert (1986a), considering four of the LME metals over a sample which ended in 1978 found significant lagged price change effects for lead (not apparent in this sample) and zinc (sign reversed in these tests). The biases were also positive on the earlier sample. These contrasts warn of the danger of taking these estimated relationships as structural.

VI. CHOICE OF EXCHANGE RATE INDEX

41. Equation (3) shows that the weight assigned to a particular currency (say that of country j) in the construction of the Ridler-Yandle exchange rate index X_i specific to commodity i should, in principle, depend on both the commodity i supply and demand elasticities in j and on j's share of world production and consumption of the commodity. However, Gilbert (1987) shows that in a multicommodity model the weights will be complicated functions of all own and cross demand and supply elasticities and all market share parameters. This suggests that the construction of commodity specific indices may be of limited value. However, that study used aggregate commodity indices, and it was therefore not possible to test that view on that data set.

42. The results reported in Section 5 are based on regressions which employ a GNP-weighted exchange rate index. We also experimented by substituting (in the same regressions) 1/ (i) the IMF MERM index in which the weights reflect the weights of each currency in US overseas trade; 2/ (ii) commodity-specific indices employing weights proportional to, respectively, consumption of the commodity in each country, production of the commodity and production plus consumption. 3/ Referring to equation (3), use of consumption weights amounts to assuming that demand elasticities are uniform across

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1/ The use of equation specifications chosen on the basis of estimates which employ the GNP-weighted index will, in principle, imply a (probably small) bias in favor of that index.

2/ IMF, International Financial Statistics, (various issues).

3/ I am grateful to Sompheap Sem for assembling the data required for these reweighting exercises.

countries and that supply elasticities are negligible by comparison; the use of production weights requires the converse assumption; and the use of the average weights requires the assumption that supply and demand elasticities are equal and uniform across countries.

43. The results of these tests are summarized in Table 7. An obvious question is which choice of index gives the best fit. Since these models are nonnested, no formal test is possible simply on the basis of these regressions; but it is in any case doubtful whether any of the available nonnested tests would be able to discriminate effectively between alternative indices of this sort in regressions in which the fit is relatively poor. However, an impressionistic test is provided by comparison of the model log-likelihoods which are listed in the final rows of the table. Here it is notable that the commodity-specific indices give higher likelihood values than the nonspecific indices implying that, despite the theoretical considerations advanced in Gilbert (1987), there is merit in adopting a commodity-specific weighting. Within each class, weighting by consumption shares appears superior to weighting by either production shares or by an average of the two; and GNP weights are superior to trade weights.

44. A second comparison relates to the size of the estimated long run exchange rate elasticities. The theory of section 2 implies that these elasticities should be unity, and the multicommodity generalization proposed in Gilbert (1987) implies that these elasticities should be within the interval

$[0, (1-v_1)^{-1}]$. 1/ It is notable that the choice of index can have a significant effect on the size of these elasticities, and that, in particular, use of the MERM index appears to give rise to high estimated elasticities.

45. The high elasticities associated with the MERM index reflect the fact that the use of trade weights gives a particularly high weight to currencies (in particular to the Canadian dollar) 2/ which vary relatively little in relation to the US dollar. The more closely a country is linked through trade to the United States, the less its currency will move in relation to the dollar. 3/ Since the MERM index shows relatively less variation than the other indices, the variation that is shown must take a higher weight in the estimated regressions. This may go some way towards explaining the difficulties experienced by Gilbert (1987) in obtaining estimated elasticities which satisfy the theoretical restrictions.

1/ [0,1] in Gilbert (1987) where the MERM index is used. The weights in the MERM index sum to unity over the set of countries 2...n which exclude the United States. In this study we scaled the MERM index by the factor $[1-v_1]^{-1}$ where v_1 is the US weight in the GNP weighted index. This should ensure that the elasticities obtained from the MERM index in this study are comparable with those obtained from the other indices.

2/ The Canadian dollar has a weight of 0.129 in the MERM index which is to be compared with 0.036 in the GNP-weighted index. The production-weighted index also suffers from a problem relating to the Canadian dollar which is apparent in the positive elasticities estimated for nickel where Canada is the major producer--the Canadian dollar has a weight of 0.568 in the nickel-specific production weighted index.

3/ The standard deviation of the MERM index is 0.103 over the sample 1978q1-1985q4. Over the same sample, the GNP-weighted index has standard deviation 0.116.

VII. CONCLUSIONS

46. This paper has two purposes. The first is to extend standard market efficiency tests to intermarket efficiency by looking at the extent to which changes on the forex, t-bill and final product markets are arbitraged across into metals markets. Secondly, we are concerned to quantify the response of metals prices to exchange rate changes.

47. The results of the market efficiency tests have been predominantly positive. The seven LME metals all respond as predicted to exchange rate and inflation innovations and the evidence for noninnovational responses to these variables appears inconclusive. On the other hand, except in the case of copper, the interest rate responses are poorly defined and insignificant.

48. An additional benefit of the use of these financial market variables in the efficiency tests is that this substantially increases the power of the standard weak form efficiency tests by reducing the equation standard errors. The power of these tests is further increased by the use of systems estimators. As a consequence we have been able to isolate much clearer evidence of bias and lagged dependence than in previous work on the same markets. In the light of these tests, none of the LME metals markets appears to conform exactly to the efficient market paradigm.

49. The implications of this finding for policy are slight. Market inefficiency, as evidenced by bias and lagged dependence, is generally a consequence of thin trading; and this increases the danger of market manipulations. The tin experience is relevant here, and there is some suggestion that the British market regulators were insufficiently vigilant of tin trading on the LME. 1/ The evidence of inefficiency provided in this paper

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1/ See Anderson and Gilbert (1986).

reinforces the need for regulatory vigilance. The evidence of market inefficiency does imply that structural econometric commodity price models may to some extent generate predictions superior to those estimated by simply taking the published forward price, and this provides some encouragement to the econometric modeling program.

50. On the question of the long-run exchange rate elasticity, the estimates reported in this paper suggest wide variation across metals. The average elasticity across all seven metals is in line with theoretical predictions, but only a single metal (copper) exhibits an elasticity of the predicted magnitude. One group of metals (aluminum, nickel and zinc) appears very insensitive to changes in the value of the dollar, and this may reflect the influence of producer pricing on a dollar basis. However, this tendency is offset by a second group (lead, silver and tin) which appears overresponsive to changes in the value of the dollar. Finally, we found that the estimated elasticities are quite sensitive to the weights used in construction of the exchange rate index; that consumption-weighted commodity specific indices appear to perform best; and that the MERM trade-weighted index appears less satisfactory in this context.

Table 1: ESTIMATES ON THE NULL HYPOTHESIS

Dependent variable $\nabla \ln p$

		$\nabla \ln X$	∇r	$\nabla \Pi$	$\nabla \ln A$	ϕ	df	R^2	DW	LM
Ag	OLS	-2.37	-1.53	3.37	3.88		28	0.206	2.00	0.54
		(1.81)	(0.32)	(0.82)	(1.36)					(0.71)
	ARCH-OLS	-2.00	0.25	1.96	1.13	0.31	28	0.142	2.13	0.52
		(1.39)	(0.08)	(0.52)	(0.54)	(1.77)				(0.72)
Al	OLS	-1.15	-1.03	2.98	1.65		22	0.240	1.40	1.04
		(2.36)	(0.60)	(2.49)	(2.47)					(0.41)
Cu	OLS	-0.93	-1.76	3.24	1.86		28	0.326	2.28	0.68
		(2.19)	(1.41)	(2.06)	(1.90)					(0.61)
	ARCH-OLS	-0.84	-1.69	3.18	1.39	0.22	28	0.315	2.33	0.81
		(1.48)	(1.30)	(2.32)	(1.76)	(1.22)				(0.53)
Ni	OLS	-0.68	-0.81	2.57	1.86		19	0.162	1.82	0.54
		(1.20)	(0.51)	(1.55)	(3.18)					(0.71)
Pb	OLS	-1.07	0.95	1.85	1.58		28	0.147	1.47	0.66
		(1.30)	(0.61)	(1.22)	(1.52)					(0.63)
	ARCH-OLS	-1.11	0.98	1.77	1.56	0.02	28	0.147	1.46	0.68
		(1.38)	(0.56)	(1.03)	(1.50)	(0.13)				(0.62)
Sn	OLS	-1.39	-0.45	0.89	0.20		28	0.233	2.08	0.37
		(2.00)	(0.43)	(1.06)	(0.33)					(0.83)
Zn	OLS	-0.21	-0.51	1.81	1.63		28	0.101	2.04	1.05
		(0.32)	(0.48)	(1.36)	(1.68)					(0.40)
	ARCH-OLS	-0.33	-1.27	2.47	1.96	0.33	28	0.090	2.02	0.79
		(0.43)	(0.74)	(1.54)	(1.98)	(1.74)				(0.54)

Notes: t statistics in parentheses under coefficients--heteroscedasticity-consistent in case of OLS estimates. The LM statistic tests for serial correlation of order up to 4 and is distributed F(4,df-4); the marginal significance (area under the right hand tail) is given in parentheses. The ARCH-OLS regressions are only computed if Φ is estimated as positive. The statistics for the ARCH-OLS regressions are estimated from the unscaled residuals.

Table 2: TESTS FOR DEPARTURES FROM MARKET EFFICIENCY

		Ag	Al	Cu	Ni	Pb	Sn	Zn	$\chi^2(7)$
Intercept									
	OLS	-0.03	-1.19	-0.45	-0.94	-0.99	1.23	-0.68	10.34
	ARCH-OLS	-0.32		-0.76		-1.32		-0.45	11.35
Lagged dependent variable									
	OLS	-0.90	4.67	-1.99	-0.18	1.42	-0.76	-1.49	8.65
	ARCH-OLS	-0.64		-1.90				-1.19	8.54
Lagged Innovations									
$\nabla \ln X(-1)$	OLS	1.42	-0.69	-0.20	-0.43	-0.47	1.04	1.49	12.21
	ARCH-OLS	0.79		-0.58				0.78	9.42
$\nabla r(-1)$	OLS	-0.23	-0.80	0.52	0.64	-0.11	-0.56	0.75	5.08
	ARCH-OLS	-1.64		0.02	-	0.10		0.42	8.41
$\nabla \Pi(-1)$	OLS	-0.66	1.05	-0.82	0.42	1.27	1.21	-0.09	8.92
	ARCH-OLS	-0.33	0.85	-0.78		0.93		-0.64	9.17
$\nabla \ln A(-1)$	OLS	-1.74	-0.69	-1.62	-0.88	-1.32	-0.30	-0.73	5.76
	ARCH-OLS	-1.15		-1.19		-1.06		-0.41	4.84
Noninnovational variables									
fprem(-1)	OLS	0.15	-1.86	0.43	0.64	-2.08	0.86	0.26	12.76
	ARCH-OLS	-0.34	-0.03	-1.95				0.38	12.03
$\Delta \ln X$	OLS	-0.35	2.29	-0.06	-0.43	-0.36	1.25	-0.19	11.24
	ARCH-OLS	0.25	1.79	0.36				-0.30	11.64
Δr	OLS	-0.39	0.67	-0.13	0.64	0.38	-0.77	-0.70	4.73
	ARCH-OLS	0.48		0.32		-0.10		-0.30	3.12
$\Delta \Pi$	OLS	-0.95	-1.32	-0.09	0.42	-0.42	-1.06	-0.18	6.43
	ARCH-OLS	-0.99	-1.14	0.36		-0.41		0.36	7.35
$\Delta \ln A$	OLS	-1.28	-1.27	-0.94	-0.88	-0.21	-0.79	-0.44	5.75
	ARCH-OLS	-0.91		-1.02		-0.24		-0.23	5.28

Notes 'fprem' is the forward premium $\ln p^f_{t-1} - \ln p_{t-1}$. Quoted statistics are t-statistics (heteroscedasticity-consistent in the OLS case) from addition of specified variable to the null regressions reported in Table 1. Degrees of freedom: as Table 1, except one degree lost in LDV regressions for Al and Ni.

The $\chi^2(7)$ statistic tests exclusion of the entire vector of variables in respectively the SUR and ARCH-SUR estimates. The 95% critical value for $\chi^2(7)$ is 14.07. ARCH corrections (OLS and SUR) were only made where ϕ was estimated as positive.

Table 3: PRICE INNOVATION RELATIONSHIPS--OLS ESTIMATES

	Ag	Al	Cu	Ni	Pb	Sn	Zn
Dependent variable	∇lnp	∇lnp	∇lnp	∇lnp	Δlnp	∇lnp	∇lnp
Intercept (x100)		-1.36 (0.77)		-2.68 (0.98)		2.04 (1.39)	-2.14 (0.82)
∇lnp$_{-1}$	-0.23 (0.99)	0.43 (4.84)	-0.37 (1.99)			-0.14 (1.18)	-0.34 (1.72)
∇lnX	-3.26 (2.02)	-0.34 (0.68)	-1.52 (3.26)	-0.20 (0.27)	-0.77 (1.00)	-1.70 (2.85)	-0.23 (0.27)
∇lnX$_{-1}$					-0.51 (0.76)		
∇r		-0.53 (0.32)	-1.65 (1.62)			-0.27 (0.27)	
$\nabla\Pi$	5.25 (1.30)	1.20 (1.04)	4.82 (2.99)	1.18 (0.72)	1.53 (1.69)	1.02 (1.33)	2.47 (2.38)
∇lnA	3.39 (1.58)	1.48 (2.49)	1.41 (2.38)	1.59 (1.89)	1.90 (2.09)		1.59 (2.16)
df	28	19	27	18	28	27	27
R^2	0.233	0.419	0.410	0.143	0.238	0.282	0.205
DW	1.71	1.23	1.89	1.84	1.65	2.04	1.59
LM F(4,df-4)	0.45	0.39	0.18	0.73	1.05	0.43	0.79
ms	(0.77)	(0.81)	(0.94)	(0.58)	(0.40)	(0.78)	(0.54)
Restrictions h	4	2	3	4	4	3	3
F(h,df-h)	0.21	0.59	0.21	0.64	0.22	0.31	1.38
ms	(0.93)	(0.56)	(0.89)	(0.65)	(0.92)	(0.81)	(0.27)
Implied long-run exchange rate elasticity	-2.65	-0.59	-1.11	-0.20	-1.28	-1.49	-0.17

Notes: Heteroscedasticity-consistent t-statistics in parentheses under coefficient estimates. 'ms' (marginal significance is the area under the right-hand tail of the distribution.

Table 4: PRICE INNOVATION RELATIONSHIPS--ARCH-OLS ESTIMATES

	Ag	Al	Cu	Ni	Pb	Sn	Zn
Dependent variable	∇lnp		∇lnp		Δlnp		∇lnp
Intercept (x100)							-1.76 (0.71)
∇lnp$_{-1}$	-0.13 (0.56)		-0.37 (1.90)				-0.30 (1.45)
∇lnX	-2.24 (1.48)		-1.46 (2.26)		-0.63 (0.79)		-0.27 (0.34)
∇lnX$_{-1}$					-0.49 (0.67)		
∇r			-1.72 (1.38)				
$\nabla\Pi$	3.25 (0.83)		4.81 (3.25)		1.29 (0.86)		2.41 (1.50)
∇lnA	0.68 (0.33)		1.41 (1.81)		1.84 (1.88)		1.72 (1.70)
ϕ	0.39 (2.33)		0.06 (0.31)		0.04 (0.21)		0.08 (0.35)
df	28		27		28		27
R^2	0.077		0.389		0.236		0.199
DW	1.94		1.88		1.64		1.65
LM F(4,df-4) ms	0.60 (0.66)		0.16 (0.87)		1.09 (0.38)		0.76 (0.56)
Restrictions h	4		3		4		3
F(h,df-h) ms	0.23 (0.92)		0.21 (0.89)		0.21 (0.93)		1.31 (0.29)
Implied long-run exchange rate elasticity	-1.98		-1.07		-1.12		-0.23

Notes: t-statistics in parentheses under coefficient estimates. 'ms' (marginal significance is the area under the right-hand tail of the distribution. ARCH-OLS estimates were only computed where ϕ was estimated as positive.

Table 5: INNOVATION RELATIONSHIPS--SUR ESTIMATES

	Ag	Al	Cu	Ni	Pb	Sn	Zn
Dependent variable	∇lnp	∇lnp	∇lnp	∇lnp	Δlnp	∇lnp	∇lnp
Intercept (x100)		-2.09 (1.81)		-1.92 (1.04)		2.63 (2.09)	-2.08 (1.22)
∇lnp$_{-1}$	-0.24 (2.60)	0.44 (4.04)	-0.27 (2.87)			-0.20 (1.51)	0.32 (2.52)
∇lnX	-3.13 (2.82)	-0.01 (0.03)	-1.33 (2.89)	-0.20 (0.32)	-0.47 (0.73)	-1.81 (3.69)	-0.25 (0.41)
∇lnX$_{-1}$					-0.88 (1.69)		
∇r		-0.37 (0.59)	-1.64 (2.57)			-0.10 (0.11)	
$\Delta\Pi$	5.53 (2.47)	0.66 (0.72)	4.22 (4.58)	0.37 (0.29)	1.28 (1.04)	0.97 (1.03)	2.27 (1.85)
∇lnA	3.35 (2.46)	1.16 (2.37)	1.42 (2.75)	0.96 (1.19)	1.90 (2.51)		1.52 (2.01)
Implied long-run exchange rate elasticity	-2.52	-0.02	-1.04	-0.20	-1.35	-1.51	-0.19
Residual correlations Ag		0.52	0.56	-0.31	0.02	0.16	0.06
Al			0.44	0.25	-0.19	0.14	0.04
Cu				0.26	0.15	0.41	-0.05
Ni					0.15	0.11	0.24
Pb						0.36	0.39
Sn							0.21
Log-likelihood				-321.1			
Likelihood ratio test against unrestricted model			χ^2 (23) = 30.1		(ms = 0.15)		

Notes: t-statistics in parentheses under coefficient estimates. The SUR estimates were estimated over the complete sample 1978q1-1985q4 with absent data for Al and Ni (and the values for the corresponding X variables) replaced by zeros. The error covariance matrix Σ was estimated over the restricted sample 1980q3-1985q4, using a two stage procedure in which the initial estimate of Σ was calculated from the residuals from the regressions reported in Table 3.

Table 6: PRICE-INNOVATION RELATIONSHIPS--ARCH-SUR ESTIMATES

		Ag	Al	Cu	Ni	Pb	Sn	Zn
Dependent variable		∇lnp	∇lnp	∇lnp	∇lnp	Δlnp	∇lnp	∇lnp
Intercept (x100)			-1.50 (1.30)		-1.58 (0.85)		2.67 (2.13)	-1.63 (0.99)
∇lnp$_{-1}$		-0.18 (1.51)	0.40 (3.61)	-0.33 (3.26)			-0.19 (1.45)	-0.22 (1.76)
∇lnX		-2.40 (2.28)	-0.22 (0.48)	-1.42 (3.06)	-0.59 (0.91)	-0.54 (0.82)	-1.81 (3.71)	-0.32 (0.52)
∇lnX$_{-1}$						-0.82 (1.57)		
∇r			-0.64 (0.98)	-1.81 (2.77)			-0.18 (0.20)	
$\nabla\Pi$		5.02 (2.09)	1.44 (1.62)	4.92 (5.21)	1.34 (1.05)	1.38 (1.12)	1.16 (1.24)	2.34 (2.00)
∇lnA		1.62 (1.13)	1.27 (2.71)	1.29 (2.45)	1.22 (1.51)	1.89 (2.51)		1.67 (2.25)
ϕ		0.38 (2.24)		0.07 (0.39)		0.01 (0.06)		0.08 (0.36)
Implied long-run exchange rate elasticity		-2.04	-0.36	-1.06	-0.59	-1.36	-1.52	-0.26
Residual correlations	Ag		0.50	0.56	-0.29	0.04	0.24	0.13
	Al			0.38	0.25	-0.19	0.16	0.06
	Cu				0.27	0.12	0.42	-0.05
	Ni					0.17	0.09	0.25
	Pb						0.37	0.38
	Sn							0.19
Log-likelihood					-320.8			
Likelihood ratio test against unrestricted model				χ^2 (23) = 27.0		(ms = 0.26)		

Notes: t-statistics in parentheses under coefficient estimates. The SUR estimates were estimated over the complete sample 1978q1-1985q4 with absent data for Al and Ni (and the values for the corresponding X variables) replaced by zeros. The error covariance matrix Σ was estimated over the restricted sample 1980q3-1985q4, using a two stage procedure in which the initial estimate of Σ was calculated from the residuals from the regressions reported in Table 4. The ARCH coefficient ϕ, which was restricted to being nonnegative, was estimated from the residuals from the Table 5 SUR estimates.

Table 7: CHOICE OF EXCHANGE RATE INDEX

LONG-RUN EXCHANGE RATE ELASTICITIES AND LOG-LIKELIHOOD

		Nonspecific Indices		Commodity-specific Indices		
		GDP Weights	MERM Weights	Consumption Shares	Production Shares	Consumption + Production Shares
Ag	SUR	2.52	-3.26	-2.79	-3.20	-2.75
	ARCH-SUR	-2.04	-2.66	-2.27	-3.06	-2.57
Al	SUR	-0.02	-0.24	-0.30	-0.04	-0.00
	ARCH-SUR	-0.36	-0.26	-0.68	-0.31	-0.06
Cu	SUR	-1.04	-1.26	-1.05	-1.24	-1.08
	ARCH-SUR	-1.06	-1.29	-1.05	-1.36	-1.08
Ni	SUR	-0.20	-0.14	-0.32	1.12	0.05
	ARCH-SUR	-0.59	-0.61	-0.66	0.86	0.04
Pb	SUR	-1.35	-1.37	-1.20	-0.72	-1.03
	ARCH-SUR	-1.36	-1.41	-1.19	-0.94	-1.11
Sn	SUR	-1.51	-1.79	-1.66	-1.75	-1.87
	ARCH-SUR	-1.52	-1.81	-1.67	-1.81	-1.86
Zn	SUR	-0.19	-0.21	-0.21	-0.10	-0.14
	ARCH-SUR	-0.26	-0.30	-0.27	-0.14	-0.14
Average	SUR	-0.98	-1.18	-1.08	-0.92	-0.97
	ARCH-SUR	-1.03	-1.19	-1.11	-0.97	-0.97
Log-likelihood	SUR	-321.1	-321.9	-320.1	-321.6	-320.9
	ARCH-SUR	-320.8	-321.6	-319.8	-319.6	-320.2

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