Futures Market Efficiency: Evidence from Cointegration Tests

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INTRODUCTION

This article has two major aims: first, the "efficient market" hypothesis is examined for four nonferrous metals—copper, lead, tin, and zinc—traded in the London Metal Exchange (LME). A futures market is efficient relative to an information set such that only new unanticipated information leads to a price change. Second, the recently developed cointegration theory is employed to test efficiency in these markets.[1] A problem in testing market efficiency is that financial price series are generally not stationary.[2] Consequently, conventional statistical procedures are no longer appropriate for testing market efficiency, because they tend to bias toward incorrectly rejecting efficiency. The use of a cointegration approach properly accounts for the nonstationary behavior of futures and spot price series. Cointegration between these two variables implies that they never drift far apart. The market efficiency hypothesis, on the other hand, requires that the current futures price and the future spot price of a commodity are "close together." If these two price series are not cointegrated, they will tend to deviate apart without bound, which is contrary to the market efficiency hypothesis. On the other hand, if, for example, the spot prices in two different markets are cointegrated, then one of the prices must help in forecasting the other. Since market efficiency implies that the price at each point in time should include all available information and, given past prices, no other information should improve prediction of futures price, then cointegration of two speculative markets for two different assets implies inefficiency.

The notion of market efficiency is of considerable importance to any investor who wishes to use these markets to hedge against price risk. Profitable trading strategies cannot exist in an efficient market. Consequently, agents can engage in efficient markets at lower transaction costs than in markets which require extensive

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[1]Hakkio and Rush (1989) employ a similar test of market efficiency in the foreign exchange market.

[2]See Ma and Hein (1990) for more details on this issue.

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information search. Moreover, Turnovsky (1979) shows that an efficient futures market can have important implications for international commodity agreement, domestic stabilization schemes, and form of governmental intervention.

In recent years, a number of studies investigate the degree and source of inefficiency in commodity futures markets. These studies employ a conventional approach for testing the efficiency of futures markets for a variety of commodities. Following Fama (1970), they employ either weak form tests of efficiency (e.g., Cargill and Rausser, 1975; Goss, 1981; Kofi, 1973; Solt and Swanson, 1981; Tomek and Gray, 1970) or semi-strong form tests of efficiency (e.g., Garcia et al. 1988; Goss, 1983, 1988; Gupta and Mayer, 1981; Leuthold and Hartman, 1979).[3] The weak, form test relies on the historical sequence of prices and involves regressing cash price at contract maturity on a previous futures price. If the intercept coefficient is zero and the slope is one, the market is regarded as efficient, and the futures price is considered to be an unbiased predictor of the subsequent cash price. However, this test is criticized on the ground that the coefficient estimates are based on the ex post knowledge of the data that is not available to the agents in the market forcasting ex ante. Hence, this test uses more knowledge than efficient speculators are likely to possess and, thus, do not constitute a valid test of market efficiency.[4]

As an alternative, subsequent research centers on whether future prices fully reflect all publicly available information at the time of contracting; this tests for market efficiency in a semi-strong form sense. In this case, an econometric model is employed to compare the forecast error of the model with that of futures price. However, results from this test are contradictory. While several authors find evidence in support of the market efficiency hypothesis (Goss, 1988; Gupta and Mayer, 1981), others find rather weak results (Goss, 1983).

Recent developments in the theory of cointegration provide new methods for testing market efficiency. Hakkio and Rush (1989) and Shen and Wang (1990) demonstrate that, in an efficient market, the spot price and futures price should be cointegrated; whereas Granger (1986) shows that cointegration between prices in two different markets implies inefficiency. This article employs the cointegration technique to test the market efficiency hypothesis for four metals traded in the LME.

LITERATURE REVIEW

Several studies test for market efficiency in various metal markets using both weak and semi-strong form tests. Gupta and Mayer (1981) conduct a semi-strong form test of the efficiency of five commodity markets—copper, tin, sugar, cocoa, and coffee. They identify ARIMA models and use the mean squared error to compare the forecasts from the estimated models and from the futures markets with the actual spot price. Their results fail to reject the hypothesis that the futures markets for copper and tin are semi-strong efficient.

Goss (1981) tests the hypothesis that futures prices (and lagged spot prices) are unbiased predictors of subsequent spot prices for copper, lead, tin, and zinc at the LME for the period 1971–1978. Using OLS and instrumental variable estimates, his

[3]Kaminsky and Kumar (1990) employ both a weak form as well as a semi-strong form test of market efficiency.

[4]See Gupta and Mayer (1981), Burns (1983), and Garcia et al. (1988) for more details on this issue.

results provide unqualified support for efficiency in the tin market, and tentative support for efficiency in the copper and zinc market. However, results from the lead market do not provide support for the efficiency hypothesis. Goss (1985) revises his 1981 article by introducing joint tests and extending the sample period to 1966–1984. The results indicate that, at the 5% level of significance, the revised version of the efficient market hypothesis is rejected for copper and zinc, but cannot be rejected for tin or lead. At the 1% level, the hypothesis cannot be rejected for any of the four metals.

Goss (1983) tests the hypothesis that the market for copper, tin, lead, and zinc at the LME is efficient in a semi-strong form sense. Using data for the 1971–1979 period, his results support the view that none of these markets is efficient in the semi-strong form sense.

Canarella and Pollard (1986) analyze the efficiency of the LME for the period 1975–1983 using both overlapping and nonoverlapping data and three different estimation methods. The results conform to the predictions of the unbiased predictor hypothesis regardless of the methodology and/or data employed.

Goss (1988) conducts a semi-strong form test of the efficiency in the copper and aluminum markets at the LME. The test consists of a comparison of the predictive powers of the futures markets. Based on the "mean squared error" criterion, the results indicate that the hypothesis of an efficient copper and aluminum market cannot be rejected.

MacDonald and Taylor (1989) investigate the presence of time-varying risk premia in the copper, lead, tin, and zinc price series in the LME, conditional on the assumption of rationality. Their results indicate that forward prices contain information on futures spot prices over and above the information contained in current spot prices. Sephton and Cochrane (1990) raise concerns about the data used in MacDonald and Taylor's (1989) study, and examine the market efficiency hypothesis with respect to six metals—aluminum, copper, lead, nickel, tin, and zinc—traded on the LME. Using overlapping data and both single and multimarket models, they find evidence contrary to the tenet that the LME is an efficient market.

The review of the existing literature indicates the lack of consistency in the results of various studies. Each of these studies employs a traditional hypothesis-testing procedure and makes no attempt to test the nature of stationarity of the price series employed. As discussed in detail in the next section, the issue of the nonstationarity behavior of various spot and futures price series raises concern regarding the use of conventional statistical procedures in these studies. This is more true given the recent volatility of commodity spot prices. For example, during 1983–1986, prices of metals fell by about 20% then rose by about 50% from 1986–1988. Hence, it is of interest to reinvestigate the issue of efficiency in these metal markets using the cointegration technique. This article employs this technique to test efficiency in four nonferrous metal markets—copper, lead, tin, and zinc over the period 1971–1988. This alternative statistical approach properly accounts for the nonstationary behavior of various price series, and thus tries to resolve a major concern with most of the previous studies. Moreover, very few studies examine the data for the 1980s, a highly volatile period.

MARKET EFFICIENCY AND COINTEGRATION

Following Koppenhaver (1983), futures market equilibrium is defined in terms of expected prices and a nonnegative risk premium. The relationship between any futures price and the expectation of its realization at the delivery date can be expressed as

$$FP(t - j; t) = E[FP(t; t) \mid I(t - j)] - R(t - j) \tag{1}$$

where $FP(t - j; t)$ is the futures contract price quoted at time $t - j$ for delivery at time t, E is the expectations operator, $FP(t; t)$ is the futures contract price at maturity, $I(t - j)$ is the information set at time $t - j$, and $R(t - j)$ is a risk premium conditional on the systematic risk of holding a position in the futures market, $R(t - j) > 0$.[5]

Next, assuming that market efficiency requires the futures price to be a fair game with respect to the information set and, that a zero maturity basis exists for the futures market relative to the spot market, eq. (1) can be written as

$$E[SP(t)] = FP(t - j; t) + R(t - j) \tag{2}$$

where $SP(t)$ is the spot price at time t and $SP(t) = FP(t; t)$.[6]

Eq. (1) implies that the expected futures price at time t, given the information set at time $t - j$, is greater than or equal to the currently observable futures price $FP(t - j; t)$. If $R(t - j) = 0$ in eq. (1), then the process is defined as a martingale process. Again, if $R(t - j) = 0$ and the price changes $E[FP(t; t)] - FP(t - j; t)$ are independently and identically distributed, then the process is known as a random walk. If $I(t - j)$ is limited to only spot and futures prices extending from time $t - j$ into the past, then testing the nonnegative price change represents a weak-form efficiency test. A semi-strong form test implies that $I(t - j)$ includes other publicly available information.

In general, a survey of the existing literature shows two different methods for testing efficiency. First, the spot price, $FP(t - j; t)$, is regressed on the futures contract price at maturity, $FP(t; t)$, as shown below

$$FP(t - j; t) = \alpha + \beta FP(t; t) + \varepsilon(t) \tag{3}$$

In this case, market efficiency requires that $\alpha = 0$ and $\beta = 1$.[7] An alternative method for testing efficiency is to regress the rate of depreciation on the forward premium as shown below

$$FP(t - j; t) - FP(t - j - 1; t) = \alpha + \beta[FP(t; t) + FP(t - j - 1; t)] + \varepsilon(t) \tag{4}$$

In this case also, market efficiency requires $\alpha = 0$ and $\beta = 1$.[8] However, both these methods for testing parameter restrictions are criticized on various grounds. Maberly (1985) argues that the OLS estimates of eqs. (3) and (4) are biased as a result of an ex post censored data problem related to the forecast error. Ex post, the forecast error and the forecast are negatively correlated. On the other hand, Elam and Dixon (1988) suggest that if the futures or spot price changes follow a random walk, eqs. (3) and (4) turn in a regression model with a lagged dependent variable

[5]This equation highlights the importance of examining the effects of risk on market efficiency. The conventional approach of market efficiency is limited by its failure to consider possible source of bias in using futures prices as expectation of subsequent spot prices. Risk aversion, irrational market behavior, imperfect capital markets, transaction costs, and information costs lead to differences between current futures prices and subsequent spot prices. See Garcia et al. (1988) for more details on this issue.

[6]This assumes that cash and futures prices converge at maturity.

[7]For an application of this method, see Frenkel (1979) and the references therein.

[8]See Hansen and Hodrick (1980) and Huang (1984) for a description of this method.

which, in turn, causes an excessively higher number of rejections of the efficiency hypothesis.

However, a more fundamental problem, as mentioned in Hein, Ma, and MacDonald (1990), is that the validity of hypothesis tests is seriously complicated by the nonstationarity of the time series. Since spot and future prices are either assumed to have unit roots (Elam and Dixon, 1988), or found to be nonstationary (Ma and Hein, 1990), it raises a serious problem for the unbiasedness testing in eqs. (3) or (4). Engle and Granger (1987) show that, since nonstationary variables have infinite variances that make the F-tests or t-tests invalid, the standard hypothesis testing does not apply to time series with unit roots.

Hakkio and Rush (1989) and Shen and Wang (1990) show that recent developments in the theory of cointegration provide new methods for testing market efficiency. These methods can properly account for the nonstationarity in the price series. If two variables are cointegrated, they are said to have a common trend, causing them to move together in the long run. The presence of cointegration between the spot price and the futures contract price is a necessary condition for market efficiency. Market efficiency implies that the futures price does not consistently over- or underpredict the spot price. If these two prices are not cointegrated, then they will tend to deviate apart without bound. This is contrary to the market efficiency hypothesis. Moreover, market efficiency implies testable restrictions in the form of $\alpha = 0$ and $\beta = 1$. Stock (1987) suggests methods for testing such restrictions in a cointegrating framework.

The cointegrating approach is used here to test efficiency in two different ways. First, in the case of the four metal markets considered—copper, lead, tin, and zinc—it is determined whether either the spot or futures price in any one of these markets is cointegrated with the corresponding spot or futures price in any of the other three markets.[9] Following Granger (1986), it can be argued that the presence of cointegration between two speculative markets for two different assets implies predictability which, in turn, indicates that one market is caused by the other and, hence, inefficiency exists.

Second, tests for cointegration are conducted between the spot and futures price within each of these four markets. Hakkio and Rush (1989) argue that "...market efficiency implies that even if the spot and forward rates are non-stationary, they never drift apart so that they will be cointegrated" (p. 76). If they are not cointegrated, then their difference $SP(t) - FP(t - j; t) = U(t)$ is nonstationary—say, a random walk—and the variables can drift infinitely apart. In this case, the futures prices' prediction of the next period's spot price can be improved by using available information. Hakkio and Rush (1989) show that if, for example, $SP(t) = FP(t - j; t) + n(t - j)$ where $n(t - j) = n(t - j - 1) + e(t - j)$, then, by substitution, $SP(t) = FP(t - j; t) + n(t - j - 1) + e(t - j)$. Thus, the futures price does not include all available information and the market is inefficient. On the other hand, in an efficient market, the futures and spot prices never drift far apart so that they will be cointegrated.[10] In this case, if $SP(t)$ and $FP(t)$ are cointegrated

[9]Hsieh and Kulatilaka (1982) and Sephton and Cochrane (1990) also suggest the need for a similar multi-market test of market efficiency, arguing that the metals under study are traded in the same location, and hence a wider view of efficiency is required to establish whether forecast errors across metal markets aid in predicting current forecast errors in other markets in the Exchange.

[10]Hakkio and Rush (1989) show that cointegration is necessary but not sufficient for market efficiency. For the market to be efficient, the cointegrating vector must be one. Moreover, cointegration requires the residuals of the cointegrating regression to be only stationary, while market efficiency requires the residuals to be white noise.

with the cointegrating vector being one, their difference $SP(t) - FP(t - j; t) = U(t)$ is stationary.

COINTEGRATION TESTS

Recent works by Granger and Weiss (1983), Granger (1986), Engle and Granger (1987), Johansen (1988), and others, introduce a relatively new statistical concept of cointegration of time series. A vector of time series, each element of which is dominated by a unit root and, therefore, stationary only after appropriate differencing, may have linear combinations that are stationary without differencing. If this is the case, the variables in the vector are said to be cointegrated and can be expressed as a cointegrating vector. Cointegration implies a long-run equilibrium relationship between the variables, which will often have an economic interpretation.

Consider two time series, say, X_t and Y_t. Suppose X_t and Y_t are nonstationary, and need to be differenced once to induce stationarity. Both these variables are said to be integrated of order 1 $[X, Y \sim I(1)]$. In this case, most linear combinations of X_t and Y_t, such as $X_t - aY_t = V_t$, are also nonstationary. Usually, if X_t and Y_t are $I(1)$, then V_t will also be $I(1)$. Consequently, regressing Y on X or X on Y leads to a nonconstant error variance. Even if there is no "true" relationship between X and Y, such regression will frequently result in relatively high R^2. This spurious regression problem arises because $I(1)$ variables appear to trend together.[11]

However, there may be a number "d" such that $X_t - dY_t = U_t$ is stationary, i.e., $U \sim I(0)$. In this special case, X_t and Y_t are said to be cointegrated of order (1, 1), with a cointegrating vector of "d." The regression $X_t = dY_t + U_t$ is known as the cointegrating regression.[12]

Tests for cointegration are based on tests for unit roots in the residuals of the cointegrating regression. Since the original work of Dickey and Fuller (1979), a number of unit root tests have been proposed in the literature to determine if two variables, X and Y, are cointegrated. Engle and Granger (1987) use Monte Carlo studies to examine the power of several of these tests in the bivariate case. They recommend the use of augmented Dickey–Fuller (ADF) test. The augmentation allows for the possibility that the residuals from the cointegrating regression follow higher-order autoregressive process.[13]

The testing procedure here follows a two-stage estimation. A test of the hypothesis that the variables, X and Y, are $I(1)$ series is conducted. If this hypothesis cannot be rejected, a test for cointegration is performed by testing the residuals from the cointegrating regression

$$X_t = c + dY_t + u_t \tag{5}$$

[11]Elam and Dixon (1988) argue that most of the previous studies on market efficiency suffer from this problem.

[12]According to Hakkio and Rush (1989), the first term of the order of cointegration shows the number of times it is necessary to difference the individual data series to attain stationarity. The second term is the reduction in the number of times it is necessary to difference the linear combination to achieve stationarity. For example, taken separately, X_t and Y_t must be differenced once to be stationary, but the linear combination $X_t - dY_t$ must be differenced zero times to be stationary, a reduction of 1 from what was required for X_t and Y_t.

[13]Engle and Yoo (1987) show that the power of the ADF test is not much affected when the true model is assumed to have higher order autoregressive components, but is actually $AR(1)$. On the other hand, the power of the original Dickey–Fuller test can fall significantly if the true model has higher-order autoregressive components.

First differences of the residuals are regressed on the first lagged-level and lagged first differences of the residuals. The variables are cointegrated only if one can reject the null hypothesis that the t-statistic for the lagged-level term is zero. The critical values for these hypothesis tests are given in Engle and Granger (1987).

EMPIRICAL RESULTS

This section addresses the hypothesis that the markets for copper, lead, tin, and zinc in the LME are efficient. For each market, the price variables employed are the futures prices for three-month delivery and subsequent spot price when the contract matures. The data used are monthly average spot and futures (three months) prices from the LME for the sample period July 1971 to June 1988. On the LME, futures prices are quoted for three-month delivery, and that particular futures contract is not quoted officially again until settlement. The LME is considered the world's leading metals futures market and trades, inter alia, in futures contracts for these metals, which are officially quoted for spot and three-month delivery only.[14] Nonoverlapping data for the period July 1971 to June 1979 are taken from Goss (1981, 1983), while data for the subsequent period are collected from various issues of *Metallgesellschaft*.[15] Sixty-four observations are used for estimation purposes.[16]

Before proceeding to test for cointegration, it is essential to determine the individual time series properties of various price variables. The presence of a unit root is often a theoretical implication of models that postulate the rational use of available information by economic agents (see, e.g., Baillie and Bollerslev, 1987; Meese and Singleton, 1983). Nelson and Plosser (1982) find strong evidence in favor of unit root nonstationarity in a number of economic time series. Although there has been very little attempt in the literature to test for nonstationarities in spot and futures prices in commodity markets, it is assumed that these prices are also likely to be nonstationary over time.[17] Moreover, nonstationarity tests are required for cointegration. Hence, two different test procedures are used first to test for the degree of integration in the spot and futures price in each of the commodity markets.

Table I reports the results of the test for the presence of a unit root in the univariate representation of each variable. Two approximately equivalent procedures are employed for detecting a unit root—the ADF test and the Phillips–Perron Z_t statistic. The ADF test accounts for temporally dependent and heterogeneously distributed errors by including lagged innovation sequences in the fitted regression. On the other hand, the Phillips–Perron test accounts for nonindependent and identically distributed processes using a nonparametric adjustment to the standard Dickey–Fuller test. The first column in Table I shows the results of the ADF test. This is estimated by regressing the first difference of the variable on the lagged value of its level and significant lagged values of the first differences. Schwert

[14]See Fama and French (1988) for a discussion of the advantages of using data from the London Metal Exchange.

[15]Hein, Ma, and MacDonald (1990), among others, suggest the use of nonoverlapping data. They show that serial correlation, resulting from the overlap between the delivery cycle and the forecast horizon, may generate biased estimates as well as misleading conclusions.

[16]One of the problems with cointegration analysis is the number of observations required to achieve reliability in the tests. A number of studies in this area use 100 or more observations. Unavailability of consistent nonoverlapping data prior to 1971 precludes the use of a longer time series in this study.

[17]However, Gupta and Mayer (1981) fit the ARIMA model on the raw series of copper and tin.

Table I
TEST FOR A UNIT ROOT IN THE SPOT AND FUTURES PRICE

Metal	Price	t-Statistic for Augmented Dickey–Fuller Test	Z-Statistic for Phillips–Perron
Copper	Spot	−2.016	−2.489
	Futures	−1.077	−1.762
Lead	Spot	−2.116	−2.710
	Futures	−2.302	−2.645
Tin	Spot	−0.967	−1.088
	Futures	−2.302	−2.645
Zinc	Spot	−1.960	−2.842
	Futures	−2.414	−2.910

Model: $X_t = c_0 + a_0 X_{t-1} + \Sigma a_i X_{t-i} + \varepsilon_t$.

H_0: X_t contains a unit root, i.e., it is nonstationary. Reject H_0 if the t-statistic for a_0 is significantly negative.

From Fuller (1976, Table 8.5.2), the critical value for 100 observations at the 5% significance level is −3.45.

(1987) points out that unit root tests are adversely affected by the presence of moving average errors, and suggests using long lag lengths to avoid this problem. Accordingly, his formula, which gives a lag length of 12, is used. The lagged terms are individually tested for significance, beginning with the longest lag (12), dropping and sequentially testing until either significance is found or all lags exhausted. The null hypothesis of the existence of a unit root is then tested by comparing the t-statistic on the coefficient of the lagged level to the critical value. The test statistic is of the form of the usual Student's t-test, but its distribution is nonnormal, even asymptotically. Fuller (1976) reports the appropriate cumulative distribution and the critical values. The null hypothesis that each of the variables (the spot and futures price for copper, lead, tin, and zinc) has a unit root cannot be rejected. The values of the ADF statistic for each variable is less than the 5% critical value of 3.45.[18]

The Studentized version of the Phillips–Perron Z_t-statistic is reported in the second column of Table I.[19] The Z_t statistic is a simple modification of the standard t-statistic for the lagged-level term employed in the ADF test. The modification accounts for the nonindependent and identically distributed nature of the error sequence. Both the Dickey–Fuller and the Phillips–Perron tests have the same asymptotic distribution, hence, the same critical value applies as well. The results using the Phillips–Perron test are similar to those of the ADF test; in all cases, they are unable to reject the null hypothesis of a unit root. However, both the ADF and the Phillips–Perron tests on the first difference of each of the variables (not reported here) reject the unit root hypothesis. This indicates that the spot and futures prices in each of these markets are integrated of order one.

Given the presence of unit root nonstationarity in the variables considered one at a time, the existence of efficiency in the copper, lead, tin, and zinc market is tested

[18]Unit root tests are also conducted using a time trend. The results (not reported here) are similar to those given in Table I. The time trend variables are found to be insignificant.

[19]Following Schwert (1987) and Phillips and Perron (1988), respectively, the results are estimated using 10 lags of autocovariances and triangular weights.

in two ways. First, it is determined whether the spot and futures price in each of these markets are cointegrated with the corresponding spot and futures price in each of the other markets. Second, a test for cointegration between the spot and futures price is performed by examining the hypothesis that there is a stationary linear combination of these two variables in each market. The existence of market efficiency requires that the null hypothesis of no cointegration be accepted in the former case, while rejected in the latter case.

In the first case, cointegration tests are performed with the spot price in each market as the dependent variable. Since it is well-known that cointegration tests are symmetrical, only the relevant results are reported in Table II,[20] i.e., the cointegrating regressions and the ADF test statistics are shown. From Engle and Granger (1987), the relevant critical value for the 5% significance level is −3.17. In Table II, the ADF statistics uniformly reject the null hypothesis of no cointegration in all cases. The absolute value of the test statistic in each case is greater than the critical value. Hence, in general, the results suggest that spot prices in each of these markets are cointegrated. In other words, information about spot prices in one market can be used in predicting spot prices in other markets.

Table III reports similar results for futures prices in these four markets. Note that the ADF statistics fail to reject the null hypothesis of no cointegration in two cases: the copper and lead market, and the lead and zinc market. However, in all other cases, the results indicate that the futures prices in various markets are cointegrated. An analysis of Tables II and III, therefore, provide overwhelming evidence in favor of cointegration between spot and futures prices in each market with the corresponding spot and futures prices in other markets. This can be interpreted as evidence in favor of market inefficiency.

Next, an examination is conducted to determine whether the spot and futures prices from within the same market are cointegrated. The cointegrating regression and the test statistics are given in Table IV. The results from the lead, tin, and zinc

Table II
RESULTS FROM BIVARIATE COINTEGRATION TESTS OF SPOT PRICES IN DIFFERENT METAL MARKETS

Dependent Variable	Cointegrating Regression	R^2	t-Statistic for ADF Test
SP_C	$SP_C = -1.099 - 2.062SP_L$	0.78	−7.20
	$SP_C = 0.166 + 0.158SP_T$	0.56	−5.12
	$SP_C = -2.004 + 1.993SP_Z$	0.61	−4.12
SP_L	$SP_L = 1.166 + 2.339SP_T$	0.88	−4.10
	$SP_L = -1.009 - 0.733SP_Z$	0.91	−6.71
SP_T	$SP_T = 0.658 + 1.699SP_Z$	0.59	−3.66

Model: $u_t = b_0u_{t-1} + \Sigma b_iu_{t-i} + \varepsilon_t$, where u is the residual from the cointegrating regression $X_t = a_0 + a_1Y_t + u_t$.

H_0: X and Y are not cointegrated. Reject H_0 if the t-statistic on b_0 is statistically significant.

Variables: SP_C, SP_L, SP_T, and SP_Z are the spot price in the copper, lead, tin, and zinc markets, respectively. The ADF is the augmented Dickey-Fuller statistic. From Engle and Granger (1987), the critical value for ADF at the 5% significance level is −3.17. Although the R^2s are reported for completeness, it should be noted that they do not have the usual interpretation with $I(1)$ variables.

[20]A summary of the complete results are available from the author upon request.

Table III

RESULTS FROM BIVARIATE COINTEGRATION TESTS OF FUTURES PRICES IN DIFFERENT METAL MARKETS

Dependent Variable	Cointegrating Regression	t-Statistic for: R^2	ADF Test
FP_C	$FP_C = -3.056 - 3.042FP_L$	0.30	−1.61
	$FP_C = -1.563 - 2.151FP_T$	0.67	−4.12
	$FP_C = -0.115 + 0.366FP_Z$	0.50	−4.41
FP_L	$FP_L = -1.760 - 3.433FP_T$	0.71	−4.95
	$FP_L = -0.890 + 2.335FP_Z$	0.85	−1.94
FP_T	$FP_T = 4.000 + 1.069FP_Z$	0.80	−3.77

Variables: FP_C, FP_L, FP_T, and FP_Z are the futures prices in the copper lead, tin, and zinc markets, respectively.

The ADF is the augmented Dickey–Fuller statistic. From Engle and Granger (1987), the critical value for ADF at the 5% significance level is −3.17. For further description of this table refer to footnote 20.

market uniformly fail to reject the hypothesis that there is no cointegrating vector involving the spot and futures price. This finding is consistent with the hypothesis that there is no steady-state relationship between the futures and subsequent spot prices of lead, tin, and zinc. Thus, the futures price appears to be a biased predictor of the subsequent spot price in these markets. However, in the case of the copper market, the tests show some evidence of cointegration between the spot and futures price. In the cointegrating regressions with the spot price as the dependent variable, the ADF statistic rejects the null hypothesis of non-cointegration at the 10% level but not at the 5% level.[21]

Following the method suggested by Stock (1987), a test is conducted for the efficient pricing condition, $\alpha = 0$ and $\beta = 1$ in eq. (3), in each of the four markets. Failure to reject the joint null hypothesis $[\alpha = 0, \beta = 1]$ in a market implies efficiency. The results are given in Table V. The test statistic for testing this joint hypothesis has a chi-squared distribution with two degrees of freedom. The results indicate that the hypothesis is rejected at least at the 5% level in each of the four

Table IV

RESULTS FROM BIVARIATE COINTEGRATION TESTS IN EACH METAL MARKET WITH THE SPOT PRICE AS THE DEPENDENT VARIABLE

Metal	Cointegrating Regression	R^2	ADF
Copper	SP = −12.160 + 2.313FP	0.80	−3.10
Lead	SP = −2.773 + 1.965FP	0.91	−2.03
Tin	SP = 246.112 + 11.305FP	0.83	−0.91
Zinc	SP = 18.607 + 0.788FP	0.78	−1.35

Abbreviations: SP = spot price, FP = futures price, ADF = augmented Dickey-Fuller statistic.

From Engle and Granger (1987, Table 2), the critical value for the ADF statistic at the 5% significance level is −3.17. For further description of this table refer to footnote 20.

[21]Tests have been performed also for each market using the futures price as the dependent variable. The results (not reported here) are qualitatively similar to those reported in Table IV. This is not surprising given the well-known fact that cointegration tests are symmetrical.

Table V
RESULTS OF HYPOTHESIS TESTS ON α AND β

Market	$H_0: \alpha = 0$ and $\beta = 1$
Copper	12.19
Lead	15.33
Tin	8.96
Zinc	17.04

Model: $SP = \alpha + \beta FP + \varepsilon$.
The hypothesis test has a chi-squared distribution with two degrees of freedom. The critical values are: 5% significance level = 5.99; 1% significance level = 9.21.

markets. Hence, the futures price appears to be a biased predictor of the spot price. This, in turn, implies the absence of efficiency in each of these markets. These results are contrary to the findings reported in Gupta and Mayer (1981) and Goss (1988). Both these studies use semi-strong form tests. Gupta and Mayer report the acceptance of the efficiency hypothesis in the copper and tin market, while Goss finds efficiency in the copper market. The results appear to reinforce the findings of Goss (1983) and Sephton and Cochrane (1990). Goss reports the rejection of the efficiency hypothesis in the lead, tin, and zinc market, while Sephton and Cochrane find evidence contrary to the tenet that the LME is an efficient market.

SUMMARY AND CONCLUSIONS

It is now widely accepted that, in the presence of nonstationary time series, conventional hypothesis testing procedures are no longer appropriate. Empirical results over the period 1971–1988 for LME indicate the presence of unit roots in the spot and futures price of copper, lead, tin, and zinc. This implies nonstationarity in the price series of these metals, and thus raises serious concern regarding most of the previous studies that have tested for market efficiency using the levels of spot and futures prices of these metals. To account for the nonstationarity behavior of various price series, and thus resolve one of the major concerns with previous studies, this article uses the cointegration approach. The empirical results presented indicate the rejection of the "efficient market" hypothesis for four nonferrous metals—copper, lead, tin, and zinc—traded on the LME.

This article points out the problems of conventional hypothesis testing in the futures market literature and suggests how the cointegration approach can be used to circumvent some of these difficulties. The novelty of the alternative approach lies in its potential use as a method of testing efficiency in any asset market.

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