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TESTING FOR TEMPORAL ASYMMETRY

IN THE METAL PRICE-STOCK RELATIONSHIP

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Abstract: In this paper, we investigate temporal asymmetry in the metal price-stock relationship based on threshold cointegration and asymmetric non-causality tests. Using international data from a sample of four representative metals (aluminium, copper, lead and zinc), we find strong evidence of threshold cointegration between price and stocks in the metals market. Asymmetric causality tests also indicate a feedback link between prices and stocks for most metal markets. The results obtained in this study also show that negative changes for both prices and stocks have stronger effects than positive changes. Several theoretical explanations of asymmetry rationalize the findings.

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1. Introduction

Often construed as an abuse of market power, many consumers complain that metal prices at retail level respond faster to international metal markets price increase (resp. stock decrease) than to a decrease (resp. to an increase). The perceived asymmetry points to a potential gap in price and/or stock theory (Peltzman, 2000). Despite this perception, the economic and econometric analyses of international metal stock adjusments have received only minimal attention. One reason for this has been the lack of suitable inventory or stock data. Another has been the lack of any uniform theory of inventory behaviour or inventory-price relationship to serve as a basis for research. Attempts have been made to study this relationship in metal markets but with only mixed results.

This study is based on recent advances in asymmetric time series analysis (e.g., see Enders, 2001; Enders and Granger, 1998; and Enders and Siklos, 2001) which are applied to analyze price and/or stock transmission at various stages in the production and distribution chain in international metal markets. We depart from previous research in identifying the potential role metal future markets may play in asymmetric price or stock transmission. Our analysis is based on newly available inventory data for aluminium, copper, lead and zinc.

This paper is organized as follows. In Section 2 the role of stocks and prices in market equilibrium is examined. Research methodology appears in Section 3. Section 4 reports and discusses threshold cointegration results, while Section 5 deals with asymmetric causality and implications of the findings for metals markets. Section 6 concludes the paper.

2. Stocks and market equilibrium

Among received studies that have attempted to analyse the inventory and price relationship, not many exist. The principal effort has been to determine why commodity producers, consumers and dealers hold stocks and in what quantities. Among early efforts, Brennan (1958) employed agricultural commodity data to evaluate the supply of storage theory developed by Working (1949), Kaldor (1939) and Keynes (1930). Other theories that have been proposed to explain inventory behaviour are the accelerator, asset returns, buffer stocks, quadratic or S, production smoothing, and optimization, e.g. see reviews by Chikan (1984), Williams and Wright (1991) and Wright and Williams (1982). Supply of storage implies that firms will adjust their stock levels until the marginal revenue of holding stocks equals their marginal cost, the latter determined by coverage and stockout yields as well as cost of storage. Tests of this theory have followed, for example, in studies by Brennan (1958,1991) on precious metals, lumber, butter, eggs, wheat, oats and wheat; Telser (1958) on cotton and wheat, Weymar (1966) on cocoa; and Considine and Heo (2000) on crude oil and its products.

Earlier studies dealing with supply of storage that focus particularly on metals inventory behaviour include Burrows (1971) and Ghosh et al. (1987). Among more recent applications, Bresnahan and Suslow (1985) have investigated copper market dynamics in the context of the London Metal Exchange (LME) by concentrating on the asset character of stocks. They examine the rate of return of holding copper and the implications of inventory stockouts on this rate of return. Thurman (1988) interprets the supply of storage by estimating a structural model of stock equilibrium that utilizes in addition direct measurements of stocks. Pindyck (1994) concentrates on the cost aspect by studying how consumers and producers balance the costs of adjusting consumption and production with the costs of decreasing inventory holdings as a reaction to metal price fluctuations. As the spread between spot and futures prices varies, the costs of drawing down inventories determine just what quantity of inventories will be held. His conclusions confirm production cost smoothing behaviour, in which inventories are used to shift production to periods in which costs are low and in which inventories are used to avoid stockouts and to reduce scheduling costs. The attempt of Labys and Lord (1991) was to explain fluctuations in inventories as a response to how markets move in and out of equilibrium. They employ the Granger and Lee (1989) multi-cointegration approach and focus more on production and consumption than price adjustments.

Fama and French (1988) have employed spot and futures prices to examine whether the convenience yield on inventory falls at a decreasing rate as inventory increases for aluminium, copper, lead, and zinc. The implication of the theory examined is that futures prices exhibit less variation than spot prices with inventories are low. But when inventories are high, spot and futures prices should have roughly the same variability. They further test this theory by relating inventory levels to positive demand shocks induced by business cycle activity.

What is lacking in most studies is that actual metals inventory data have not been employed in reaching any conclusions about the metals inventory and price relationship, with the exception of Pindyck (1994) and Thurman (1988). In fact, inferences about the nature of inventory levels and behaviour have mostly been made in terms of the basis or the inter-temporal price spread and related cost and financial factors. What we know, however, is that the behaviour of a variety of different forms of inventories are important for understanding metal industry price movements. Consumers acting as manufacturers or processors carry operational stocks to smooth activity of their production and distribution systems to serve as a buffer against interruptions, such as for maintenance, and to facilitate transactions accounting for periodic or cyclical variations. The convenience for manufacturers comes from avoiding plant closings and maintaining reasonable cost coverage against finished product price quotations. While these activities require only low stock holdings, additional stocks will be held only with the expectation of a return or risk premium for doing so.

Producers maintain stocks to facilitate sales and deliveries but sometimes are forced to carry inventories from periods of seasonally high production to periods of low production because of cyclical turndowns in sales. Dealers or merchants carry stocks just to facilitate their everyday business of buying and selling metals, some of which call for forward delivery. Finally, the commodity exchanges themselves hold inventories not only to accommodate speculative transactions but also to better deal with market physical surpluses and deficits.

Even though the literature on the international metal market seems abundant, it is, however, important to note that most of these studies are only based on the description, behaviour and interaction between price and stocks variables in these markets. A number of questions arise immediately: (i) How do international metal markets' variables react when these markets are in equilibrium or dis-equilibrium? (ii) How do international metal markets' variables react when these markets are below or above a threshold equilibrium value? (iii) Do price levels respond faster to international metal markets disequilibrium than stocks? (iv) Do negative changes for both prices and stocks have stronger effects than positive changes in these international metal markets?

3. Econometric methodology

Our research approach depends on tests of cointegration and causality. Testing for cointegration and temporal causality between price and stock is based on a bivariate VAR representation involving two series, price and stock. The bivariate VAR representation in the presence of cointegration is written as,

$$\Delta x_t = \omega + \sum_{i=1}^{p} \beta_i \Delta x_{t-i} + \sum_{j=1}^{q} \gamma_j \Delta y_{t-j} + \theta_1 \mu_{t-1} + \zeta_{x,t} \quad (1)$$

$$\Delta y_t = \delta + \sum_{i=1}^{p} \varphi_i \Delta x_{t-i} + \sum_{j=1}^{q} \phi_j \Delta y_{t-j} + \theta_2 \mu_{t-1} + \zeta_{y,t} \tag{2}$$

where x stands for price variable and y for stock variable, μ_{t-1} is the error correction term, Δ is the lag operator, p and q are the lag lengths for price and stock respectively. A crucial limitation of equations (1) and (2) is that they do not capture the asymmetric nature of the responses of x and y to each other; nor they capture the asymmetric, if any, of the responses in the presence of asymmetric cointegration. These equations are not valid if the responses of prices to stocks depend on whether stocks are increasing or decreasing and vice versa. In addition, they are not valid in the presence of asymmetric cointegration. In order to capture these asymmetric effects, we decompose y and x and account for asymmetric cointegration as follows,

$$y_t^{+} = \begin{cases} y_t \ if \ \ y_t \geq 0 \\ 0 \ \ otherwise \end{cases} \tag{3}$$

$$y_t^{-} = \begin{cases} y_t \ if \ \ y_t \leq 0 \\ 0 \ \ otherwise \end{cases} \tag{4}$$

$$x_t^{+} = \begin{cases} x_t \ if \ \ x_t \geq 0 \\ 0 \ \ otherwise \end{cases} \tag{5}$$

$$x_t^{-} = \begin{cases} x_t \ if \ \ x_t \leq 0 \\ 0 \ \ otherwise \end{cases} \tag{6}$$

and incorporate these decompositions into equations 1 and 2.

In addition and as in Enders and Granger (1998), and Enders and Siklos (2001), we introduce asymmetric adjustments in the model by letting the deviation from the long-run equilibrium ($\widehat{\mu_t}$) behave as a Threshold Autoregressive (TAR) process. Thus, we have,

$$\Delta \widehat{\mu}_t = I_t \rho_1 \widehat{\mu}_{t-1} + \left(1 - I_t\right) \rho_2 \widehat{\mu}_{t-1} + \varepsilon_t \tag{7}$$

where I_t is the Heaviside indicator such that

$$I_t = \begin{cases} 1 \ \ if \ \ \widehat{\mu}_{t-1} \geq \tau \\ 0 \ \ if \ \ \widehat{\mu}_{t-1} < \tau \end{cases} \tag{8}$$

where τ is the value of the threshold; ρ_1 and ρ_2 are parameters to be estimated. Petrucelli and Wooldford (1984) show that the necessary and sufficient conditions for stationarity of μ_t are: $\rho_1 < 0$, $\rho_2 < 0$ and $(1+\rho_1)(1+\rho_2) < 1$. As the threshold value τ in the above equation is unknown (and there is no a priori reason to expect that it should be zero), the procedure suggested in Chang (1993) and Enders and Siklos (2001) was used to perform a grid-search. Specifically, the estimated residuals from the long-run equation were sorted in ascending order and called $\widehat{\mu}_1^{\tau} < \widehat{\mu}_2^{\tau} < \cdots < \widehat{\mu}_T^{\tau}$ where T is the number of usable observations. The largest and smallest 15 percent of the $\left\{\widehat{\mu}_t^{\tau}\right\}$ values were discarded and the remainder considered as possible thresholds. Then, for each possible threshold, the underlying model is estimated, and the preferred threshold value is selected as the one which minimizes the sum of squared residuals. In our study we consider a consistent estimation of the threshold (e.g., see Enders and Siklos, 2001). In addition, since the exact nature of the non-linearity may not be known, it is possible to allow the adjustment to depend on the change in $\widehat{\mu}_{t-1}\left(i.e., \Delta\widehat{\mu}_{t-1}\right)$ instead of the level of $\widehat{\mu}_{t-1}$. In this case, the Heaviside indicator in equation (8) becomes,

$$I_t = \begin{cases} 1 \ \ if \ \ \Delta\widehat{\mu}_{t-1} \geq \tau \\ 0 \ \ if \ \ \Delta\widehat{\mu}_{t-1} < \tau \end{cases} \tag{9}$$

Enders and Granger (1998) and Enders and Siklos (2001) show that this specification is especially relevant when the adjustment is such that the series exhibit more ‘momentum’ in one direction than the other; the resulting model is called momentum-threshold autoregressive

(M-TAR) model[1]. In equation (7) by allowing ρ_1 and ρ_2 to take different values, the model recognizes that positive and negative deviations from equilibrium can be corrected for at different speeds. In other words, having different values of ρ_1 and ρ_2 implies asymmetric adjustment (see Enders and Siklos, 2001; Chen et al., 2005). If cointegration exists, $\rho_1 < 0$ and $\rho_2 < 0$. Testing for cointegration is performed based on the t_{Max} and Φ tests proposed by Enders and Siklos (2001). The t_{Max} statistic is given by the larger t-statistic of ρ_1 and ρ_2. A significantly negative t_{Max} statistic would imply that ρ_1 and ρ_2 are both negative (e.g., see Enders, 2001 for improved critical values). The Φ test is an F-test examining the joint hypothesis of $\rho_1 = 0$ and $\rho_2 = 0$ (e.g., see Enders, 2001 for improved critical values).

If the errors in equation (7) are serially correlated, it is possible to use a TAR or an M-TAR model augmented with lagged values of $\Delta\mu_t$ for the residuals. Thus, equation (7) could be replaced by,

$$\Delta\hat{\mu}_t = I_t\rho_1\hat{\mu}_{t-1} + \left(1-I_t\right)\rho_2\hat{\mu}_{t-1} + \sum_{i=1}^{k}\gamma_i\Delta\hat{\mu}_{t-i} + \varepsilon_t \tag{10}$$

Assuming a classic TAR or M-TAR holds, Equations (1) and (2), can therefore, be rewritten as

$$\Delta x_t = \omega + \sum_{i=1}^{p}\beta_i\Delta x_{t-i} + \sum_{j=1}^{q}a_j\Delta y_{t-j}^{+} + \sum_{j=1}^{q}b_j\Delta y_{t-j}^{-} - \alpha_1^{+}I_t\hat{\mu}_{t-1} - \alpha_1^{-}\left(1-I_t\right)\hat{\mu}_{t-1} + \zeta_{x,t} \tag{11}$$

$$\Delta y_t = \delta + \sum_{i=1}^{p}c_i\Delta x_{t-i}^{+} + \sum_{i=1}^{p}d_i\Delta x_{t-i}^{-} + \sum_{j=1}^{q}\phi_j\Delta y_{t-j} - \alpha_2^{+}I_t\hat{\mu}_{t-1} - \alpha_2^{-}\left(1-I_t\right)\hat{\mu}_{t-1} + \zeta_{y,t} \tag{12}$$

In equation (1), the null hypothesis in the Granger causality test is that y does not cause x, which is represented by

$$H_0\text{: } \gamma_1 = \gamma_2 = \cdots = \gamma_q = 0 \text{ and } \theta_1 = 0 \tag{13}$$

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[1] Caner and Hansen (2001) present a statistical argument for M-TAR adjustment. If μ_{t-1} is a near unit root process, setting the Heaviside indicator using $\Delta\mu_{t-1}$ can perform better than the specification using a pure TAR adjustment.

while the alternative hypothesis is given by

$$\text{H}_1\text{: } \gamma_j \neq 0 \text{ for at least one j or } \theta_1 \neq 0 \tag{14}$$

Similar hypotheses can be formulated for equation (2) to test for causality in the reverse direction. Obviously, the value of the test statistic depends on p and q, which makes it necessary to use various information criteria to choose the optimal lag length.

If equation (11) is adopted, then to find out whether or not y_t^+ has any effect on x, the null and alternative hypotheses may be written respectively as,

$$H_0^+ : \ a_1 = a_2 = \cdots = a_q = 0 \ \text{ and } \alpha_1^+ = 0 \tag{15}$$

while the alternative hypothesis is given by

$$H_1^+ : \ a_j \neq 0 \text{ for at least one j or } \alpha_1^+ \neq 0 \tag{16}$$

To find out whether or not y_t^- has any effect on x, the null and alternative hypotheses may be written respectively as,

$$H_0^- : \ b_1 = b_2 = \cdots = b_q = 0 \ \text{ and } \alpha_1^- = 0 \tag{17}$$

while the alternative hypothesis is given by

$$H_1^- : \ b_j \neq 0 \text{ for at least one j or } \alpha_1^- \neq 0 \tag{18}$$

Similar hypotheses can be formulated for testing the effects of x^+ and x^- on y based on equation (12).

To find out whether y^- has a stronger effect on x than y^+ the following null and alternative hypotheses are used in conjunction with (11),

$$H_{a0} : \ b_j \geq a_j \text{ for } \ j = 1, 2, \cdots, q \text{ and } \alpha_1^- \geq \alpha_1^+ \tag{19}$$

and,

$$H_{a1} : b_j \ngeq a_j \ \text{ for } \ j = 1,2,..., q \text{ or } \ \alpha_1^- \ngeq \alpha_1^+ \tag{20}$$

To examine the above restrictions three different but asymptotically equivalent Granger-causality causality tests are considered. More specifically, we use the Wald (W), Likelihood ratio (LR) and the Lagrange multiplier (LM) tests based on constrained and unconstrained estimates. The main reason for the simultaneous use of all three tests is that, even though they are asymptotically equivalent, the three tests may yield different results, especially in finite samples such as ours. For example, although the inequality $W > LR > LM$ can be demonstrated theoretically and is often widely cited, particularly in the case of linear regressions, it is still possible that the three statistics could lead to different inferences. Unfortunately, this problem that is often encountered in finite samples is difficult to resolve because there is no theoretical justification for the superiority of any one technique over the others. Appendix II provides a detailed description of the three causality-testing procedures.

4. Empirical tests for threshold cointegration

This study concentrates on inventory and price linkages in relation to the London Metal Exchange (LME) for aluminium, copper, lead and zinc. Exchange transactions are recorded in the form of spot or cash (settlement) prices and near and more distant futures prices. The use of LME spot prices possesses the advantage that there are simultaneous spot and forward prices designated for fixed maturities and transacted every business day.

A related consideration is the selection of the stock or inventory variables to be used. Stock data are normally available on a country by country case but the evaluation of the storage curves for individual countries may not be fruitful. Instead a global analysis is performed based on the availability of such data from the World Bureau of Metal Statistics. These data measure total world stockholding and are also disaggregated its components: LME warehouse stocks, world consumer stocks, world producer stocks and world merchant stocks.

Brennan (1958) makes no such distinction in his analysis of agricultural commodities and relies on data from obvious storage such as grain elevators. In the present case, we initially examine the relation of the basis to each of the given categories and report those results in Table 1. As with other attempts to analyse commodity stocks, reported or measured stocks are not always synonymous with actual or total stocks. The LME (2004) has compared reported stocks with official market balances for these metals and has found reasonable correlations between the two to make the present use of the reported or measured stocks credible.

The data used in the present study are monthly and cover the period 1990 to 2003. Data sources and definitions appear in the appendix. The inventory data were compiled by Labys and mostly Xiarchos and are reported in Xiarchos (2006).

Next, before performing cointegration analysis, we used the efficient DF-GLS and the ERS unit root tests (e.g., see Elliot et al., 1996; Cheung and Lai, 1995) together with the KPSS unit root test which relies on the hypothesis that a series is a strong mixing variable to check for nonstationarity in individual price and stock series for both aluminum, copper, lead and zinc. In both unit root tests, the lag length is determined using the Akaike Information Criterion (AIC) and the Bayesian Information Criterion (BIC) approach with a maximum lag order of 8 allowed. While the unit root hypothesis could not be rejected for all level series, the first differenced price and stock series were always found to be stationary. The test results support the hypothesis that metal prices and stocks for both aluminum, copper, lead and zinc are each integrated of order one or I(1). This integration property readily lends itself to cointegration analysis.

______________________[INSERT TABLE 1 AROUND HERE]______________________

Next, in our analysis of price and stock asymmetric relationship, we employ Enders and Siklo’s (2001) test for threshold cointegration, which extends Engle and Granger’s (1987) procedure to encompass possible asymmetric adjustments to disequilibrium. Table 2 reports cointegration test results on bilateral price-stock relationships for aluminium, copper, lead and zinc assuming threshold and momentum adjustment. The table reports value of the adjustment coefficients ρ_1 and ρ_2, their t–values, and the Φ_μ and Φ_μ^* - statistics for the null hypothesis of a unit root in μ_t (no cointegration) against the alternative of cointegration with asymmetric adjustment. The lag length is selected such that the Akaike Information Criterion (AIC) and the Bayesian Information Criterion (BIC) are minimized. The F-test for symmetric adjustment $\rho_1 = \rho_2$, the underlying long run relations, the consistent estimate of the threshold as well as the value of AIC and BIC are also presented in the table.

The estimated Φ_μ and Φ_μ^* -statistics for the relationship between prices and stocks are 3.97 and 7.17 for aluminium, 5.51 and 8.64 for copper, 2.00 and 2.89 for lead and 3.03 and 13.95 for zinc. We compare those values with the critical values reported in Enders and Siklos (2001). We conclude that there is strong evidence in favour of cointegration with asymmetric adjustment between prices and stocks for aluminium, copper, lead and zinc. Even though there is evidence of cointegration with both TAR adjustment and M-TAR adjustment, clearly the AIC and BIC favor the M-TAR specification over that of TAR.

Notice that the F-statistics for the null hypothesis of symmetric adjustment ($\rho_1 = \rho_2$) reject symmetric adjustment for both TAR and M-TAR specifications at conventional significance levels in all models (aluminum, copper, lead and zinc). Also, since the AIC and BIC selects the M-TAR specification for all models, in what follows we emphasize the M-TAR models.

The point estimates for ρ_1 and ρ_2 suggest substantially faster convergence for negative (below threshold) deviations from long run equilibrium than positive (above threshold) deviations for copper and zinc while the reverse is true for aluminum and lead. For example, in the zinc price-stock model, the point estimates of ρ_1 and ρ_2 suggest that negative deviations from long run equilibrium resulting from decreases in copper stocks or increases in copper prices (such that $\Delta\hat{\mu}_{t-1} < -0.001$) are eliminated at a rate of 18 percent per month while positive deviations are eliminated at only 0.6 percent per month. Of the four models considered, the largest discrepancy between the elimination of below and above-threshold deviations occurs for the zinc model where negative deviations are eliminated at the rate of 46 percent per month, while positive deviations are eliminated at 2 percent per month. These results are mostly consistent with asymmetric adjustment on metal markets, and supportive of the so-called 'rockets and feathers' story. However, the asymmetry here is not defined in terms of positive versus negative deviations from a long run equilibrium as in the 'rockets and feathers' literature; rather it is defined in terms of the rate of change of the deviations from long run equilibrium that are below or above a certain threshold.

[INSERT TABLE 2 AROUND HERE]

5. Empirical tests for causality and implications

Next, for the purpose of testing for causality on the basis of equations (11) and (12), we need to ensure that p and q in those equations are adequate and close to the lags of the true models and therefore that lags are not included unnecessarily. For this purpose, we use the AIC and BIC defined above. After trying many combinations of the values of p and q between 1 and 8, we decided that the optimal values p = q = 1 for aluminium and lead, p = q = 2 for copper, and p = q = 3 for zinc.

The results of causality testing under the assumption of asymmetry, based on Wald, LR and LM tests, are reported in Table 3. The superscripts + and – indicate positive and negative changes respectively. These results show that causality runs from positive and negative price changes to stock and vice versa for only aluminium. For copper, causality runs from positive and negative price changes to stocks and from negative stock changes to prices.

For lead, causality runs from positive and negative stock changes to prics only. For zinc, it is only negative volume changes which cause prices. Finally, Table 4 reports results that reinforce the results reported in Table 3. These results show that negative price and stock changes have stronger effects than positive changes.

______________________[INSERT TABLE 3 AROUND HERE]_______________________

______________________[INSERT TABLE 4 AROUND HERE]_______________________

Before we present the economic interpretation of the results, and in order to place them into perspective, we need to highlight some features of metals markets, based on previous research. These features pertain to the behaviour and roles of various market participants. First, Labys and Granger (1971) and Labys (1980) demonstrate that the behaviour of hedgers in metal markets resembles the behaviour of speculators in the sense that they act on the same decision variables. The models presented in these papers do not make a distinction between hedgers and speculators acting in both markets because they have similar demand for and supply of contracts functions.

The second feature is that the market is dominated by speculators, who cause the prices to deviate from the level determined by the cost-of-carry relationship. Labys and Granger (1971) argue that one important difference between speculators and hedgers is that the latter are more likely to be the market participants, who actually require the physical commodity (for example, industrial companies), whereas speculators are the participants who are not interested in the physical commodity per se but rather in generating profit from holding ownership in that commodity (for example, financial institutions). They further argue that since financial activity dominates real activity, it is plausible to conclude that the bulk of commodity trading can be traced back to speculators and not as much to those who actually need the physical commodity. Furthermore, Labys and Granger (1970) present several intuitive explanations as to why speculators react to the arrival of new information by buying and selling in the futures market rather than the spot market. The conclusion that we want to arrive at on the basis of this discussion is that the price-inventory relationship in the metal markets is influenced largely by the behaviour of speculators, since they are the dominant players in this market. Hence, whether or not the relationship is symmetric depends on the behaviour of speculators.

So, what is the economic interpretation of our results? The finding of: (i) bidirectional causality for aluminium and to some extend for copper markets and (ii) causality from stocks to prices for lead and zinc seems to be consistent with the noise trading model of De Long et al. (1990). This model postulates that the activity of noise speculators is not based on economic fundamentals, and hence it results in a temporary mispricing. The price, however, moves towards its mean value in the long run in the absence of transitory component. Hence, the model predicts that a positive causal relationship running from prices to inventories is consistent with a feedback trading strategy of noise speculators who base their decisions on past price movements. Moreover, the model predicts that a positive causal relationship from inventories to prices is consistent with the hypothesis that price changes are caused by the actions of noise speculators.

The finding of asymmetry may sound strange, given that there is a nice intuitive explanation as to why the price-inventory relationship should be symmetric in metal market. In fact following Karpoff (1988) we may argue that the constraints on short selling (which may take the form of either a prohibition or differential costs of short and long positions) make the relationship asymmetric, as indicated by high positive correlation between inventory and prices. In the absence of these constraints, which is the case in the futures markets, there should be a symmetric relationship as indicated by zero correlations between the two variables. This intuitive explanation is, however, not inconsistent with our findings for a number of reasons. First, this explanation pertains to the first definition of symmetry given in

Enders and Granger (1998) or Enders and Siklos (2001). According to this definition, symmetry implies that inventories in a declining market tend to be different from that associated with a bear market, and so the ratio of inventories to (absolute) price change tends to vary between bull and bear markets. While this definition of symmetry is about temporal asymmetry, which is defined to exclude contemporaneous behavior, our evidence pertains to a different concept of symmetry. Second, using another definition of asymmetry, for instance, contemporaneous asymmetry in metal markets and applying these tests directly to the inventory-price ratio could have changed our results.

But even if we accept this relationship, for the sake of argument, and even if we assume that the short-selling constraints explanation is valid for temporal asymmetry, there is no reason to assume that this factor plays an exclusive role in determining the nature of the price-inventory relationship. Surely, there are other explanations and factors that play a role. It is these explanations and factors that validate our finding of asymmetry. While temporal symmetry is indicative of a market where trades have a linear unbiased response to information, asymmetry is based on trader or speculator heterogeneity and hence on the distinction between bulls and bears, between optimists and pessimists, and between informed and uninformed traders or speculators. These distinctions are the basis of the theories proposed by Epps (1975), Copeland (1976), Jennings et al. (1981) and Wang (1994).

The results obtained in this study show that negative changes have stronger effects than positive changes. Unlike the prediction of the Epps model, this result implies that the bears'demand function is steeper than that of bulls, and hence a greater level of inventory will be associated with a negative price change than with a positive price change. This proposition makes a lot of sense if we put forward the proposition that bears are quick in reacting to negative price changes, motivated by the desire to cut losses, while bulls are 'greedy', preferring to wait for further price rises before they sell (see Moosa et al., 2003). The same proposition appears to be valid if we base the reasoning on the distinction between informed and uninformed traders or speculators.

The results can also be explained by putting forward the proposition that expectations are stabilizing in a bull market and destabilizing in a bear market (see Moosa et al., 2003). Hence, when the price rises in a bull market, traders or speculators believe that it will not rise any further, and so they sell and dampen the price rise and the subsequent trading inventories. On the other hand, when the price falls in a bear market, traders or speculators believe that it will fall further and so they sell, leading to further price falls and rising trading inventories.

6. Conclusions

This paper investigated international metals markets and presented some evidence for the presence of temporal asymmetry in the price-inventory relationship. Using a sample of monthly observations over the period 1990 – 2003, the following results were found:

(i) Results of threshold cointegration clearly indicate cointegration with asymmetric adjustment between price and stock for aluminium, copper, lead and zinc;

(ii) Causality with asymmetry is found in most metals markets.

These results are explained in terms of several hypotheses, but they appear to be in contrast with the Epps (1975) hypothesis. Specifically, the results imply that the bears' demand function is steeper than the bulls'demand function. The results also indicate that expectations are stabilizing in a bull market and destabilizing in a bear market.

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Appendix I: Definitions of metal stock and price variables

Stock variables appear from World bureau of Metal Statistics. Various issues. *Metal Statistics*. London; and London Metal Exchange (2004) LME Stocks and Market Balances. London Metal Exchange, London. Price variables appear from London Metal Exchange. Various years. Metal Price Data. London Metal Exchange, London.

Stock variables
(i) Aluminium
awtots = world total stocks
(ii) Copper

cwtots = world total commercial stocks
(iii) Lead
lwtots = world total commercial stocks
(iv) Zinc
zwtots = total country stocks
Price variables
(i) Aluminium
alcp = LME cash price for aluminium
(ii) Copper
cucp = LME cash price for copper
(iii) Lead
lecp = LME cash price for lead
(iv) Zinc
zncp = LME cash price for zinc

Appendix II: Three asymptotically equivalent tests for Granger-causality

This appendix presents three asymptotically equivalent tests (Wald, Lagrange multiplier and likelihood ratio) for examining Granger-causality in vector autoregressive (VAR) processes.

(i) Wald tests for Granger-causality

Let $\theta = vec\left[\Pi_{ij}\right]_{i,j=1,2}$ be the vector of all VAR coefficients. Non-causality restrictions can be formulated as $R\theta = 0$ with suitably chosen restrictions matrix R having full row rank. Thus, testing for Granger-causality means testing the following null hypothesis,

$$H_0 : R\theta = 0 \text{ against } H_1 : R\theta \neq 0$$

Suppose an asymptotically normally distributed estimator $\hat{\theta}$ of θ is available, that is, $\sqrt{T}\left(\hat{\theta} - \theta\right) \xrightarrow{d} N\left(0, \Sigma_{\hat{\theta}}\right)$, where T is the sample size, $\xrightarrow{d}$ denotes convergence in distribution and $\Sigma_{\hat{\theta}}$ is the covariance matrix of the asymptotic distribution. Then $\sqrt{T}\left(R\hat{\theta} - R\theta\right) \xrightarrow{d} N\left(0, \Sigma_{\hat{\theta}} R'\right)$. Thus, the standard Wald statistic for testing H_0 is

$$\lambda_W = T\hat{\theta}' R' \left(R\hat{\Sigma}_{\hat{\theta}} R'\right)^{-1} R\hat{\theta}$$

where $\hat{\Sigma}_{\hat{\theta}}$ is a consistent estimator of $\Sigma_{\hat{\theta}}$.

If H_0 is true, it then follows that,

$$\lambda_W = T\hat{\theta}'R'\left(R\hat{\Sigma}_{\hat{\theta}}R'\right)^{-1}R\hat{\theta} \xrightarrow{d} \chi^2_{[J]}$$

where J is the number of restrictions.

(ii) Lagrange multiplier tests for Granger-causality

Again, let $\theta = vec\left[\Pi_{i,j}\right]_{i,j=1,2}$ be the vector of all VAR coefficients. Non-causality restrictions can be formulated as $R\theta = 0$ with suitably chosen restrictions matrix R having full row rank. Thus, testing for Granger-causality means testing the following null hypothesis,

$$H_0 : R\theta = 0 \text{ against } H_1 : R\theta \neq 0$$

To conduct the LM test for Granger-causality, we need the restricted maximum likelihood estimator, θ_R, which can be obtained by maximizing the constrained likelihood, or by solving the first order conditions from the Lagrangian function,

$$\phi(\theta, \quad \eta) = L(\theta) + \eta' R\theta$$

where $L(\theta)$ is the log-likelihood function, and η is a J – dimensional vector of Lagrange multipliers. The first-order conditions are given by,

$$d_R + F_R\eta_R = 0$$

where $d_R = \left.\frac{\partial L}{\partial \theta}\right|_{\theta=\theta_R}$, $F_R = \left.\frac{\partial (R\theta)'}{\partial \theta}\right|_{\theta=\theta_R}$; and η_R is the optimal solution for η. Furthermore, if H_0 is true, we would expect θ_R to be close to $\tilde{\theta}$ and d_R to be close to zero. Under suitable regularity conditions,

$$\frac{d}{\sqrt{T}} = \frac{1}{\sqrt{T}}\frac{\partial L}{\partial \theta} \xrightarrow{d} N\left(0, \quad \lim\left(\frac{I(\theta)}{T}\right)\right)$$

If H_0 is true, it then follows that,

$$\lambda_{LM} = d'_R\, I(\theta_R)^{-1} d_R \xrightarrow{d} \chi^2_{[J]}$$

where $I(\theta_R)$ is a consistent estimator of the information matrix based on the restricted estimator θ_R. Another formulation of the LM test is that,

$$\lambda_{Lm} = \eta_R^{'} F_R^{'} I(\theta_R)^{-1} F_R \eta_R \xrightarrow{d} \chi^2_{[J]}$$

From $\lambda_{Lm} = \eta_R^{'} F_R^{'} I(\theta_R)^{-1} F_R \eta_R \xrightarrow{d} \chi^2_{[J]}$ and $\lambda_W = T\hat{\theta}' R' \left(R\hat{\Sigma}_{\hat{\theta}} R'\right)^{-1} R\hat{\theta} \xrightarrow{d} \chi^2_{[J]}$, it is clear that the LM test is based only on the restricted ML estimator, while the Wald test is based only on the unrestricted ML estimator.

(iii) Likelihood ratio tests for Granger-causality

The LR tests for Granger-causality uses both the restricted and unrestricted estimators. Thus, testing for Granger-causality means testing the following null hypothesis,

$H_0 : R\theta = 0$ against $H_1 : R\theta \neq \theta$

and the testing procedure is based on the following statistic,

$$\lambda_{LR} = 2\left[L\left(\tilde{\theta}\right) - L(\theta_R)\right] \xrightarrow{d} \chi^2_{[J]}$$