PRICE DISCOVERY IN THE ALUMINUM MARKET

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An extended version of the S. Beveridge and C. R. Nelson (1981) decomposition and a latent variable approach are used to examine how the noise content, and therefore the informativeness, of four aluminum prices that have been quoted at various times since 1970—the (now defunct) U.S. producer price, a transactions price reported in a trade journal, and the LME and Comex exchange prices. It was found that the start of aluminum futures trading in 1978 resulted in greater price transparency in the sense that the information content of transactions prices increased. LME prices quickly came to be more informative than published transactions prices. Although the initial Comex aluminum contract failed to attract liquidity and had low information content, the 1999 contract, trading currently, is as transparent as the LME contract. © 2005 Wiley Periodicals, Inc. Jrl Fut Mark 25:967–988, 2005

INTRODUCTION

What is the impact of exchange futures trading on transactions prices for a physical commodity? This is the principal question addressed in this article. In the absence of exchange trading, the possibility of price

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discrimination provides those transactors who have market power with the incentive to conceal actual transactions prices. This can result in similar transactions taking place at different prices. Centralized exchange trading should increase market transparency, because all actual and potential transactors will have equal access to exchange prices. Further, by facilitating speculative access, increased transparency should result in increased market efficiency, in the weak and semistrong form senses defined by Fama (1976).

Once trading takes place on an organized and liquid exchange, the exchange price becomes a common reference price for all transactions. The exchange market aggregates the information available to different transactors and, in the ideal situation, the exchange price becomes a sufficient statistic for that information (Bray, 1981). Differences may exist between the prices of different transactions, but these differences will now be clearly related to location, grade, delivery conditions, and other specifiable factors. Traders negotiate on premia and discounts relative to the exchange price rather than on the price itself.

The majority of empirical studies of price discovery are confined to the analysis of cash and futures market data in the presence of futures trading—see, for example, Garbade and Silver (1983) and, in relation to equity index futures, Booth, So, & Tse (1999) and Frino, Harris, McInish, & Tomas (2004). This focuses attention on the market efficiency issues of unbiasedness (are futures prices unbiased predictors of future cash prices?) and predictability (to what extent do futures prices enable more accurate prediction of future cash prices?). Yang, Bessler, and Leatham (2001) provide a useful survey of that literature.

Price discovery has also been discussed in market microstructure models, resulting in two alternative measures of the contribution to price discovery: information shares, as defined by Hasbrouck (1995), and the common factor component weights of Gonzalo and Granger (1995). Both the information shares (IS) metric of Hasbrouck (1995) and the permanent-transitory (PT) metric of Gonzalo and Granger (1995) are formulated within a vector error-correction model (VECM). Baillie, Booth, Tse, & Zabotina (2002) and de Jong (2002) discuss the relationship between the IS and PT metrics—see also Lehmann (2002). The results obtained from the two approaches are similar so long as the residual correlations in the VECM are small. In the more likely case that these correlations are large, PT measures do not take account of these correlations while in the IS metric, the residual covariance must be ascribed to one or other price through the ordering of the prices in the Cholesky factorization of the error-variance matrix. What this amounts to in practice is that, when there is high contemporaneous correlation in price changes across markets, there is no way of knowing in which of the two markets price discovery is taking place. Effectively, one has a single market.

Monthly average data are used in this study and so high contemporaneous correlations are to be expected, and indeed are found. The microstructure methods are therefore not helpful in the analysis of these data, and it is preferable to utilize the univariate Beverage-Nelson (BN) decomposition methodology. However, similarly to the VECM-based procedures, cointegration is required. If each price series is decomposed into a permanent and a transitory component, the information content of transactions prices before and after the introduction of futures trading can be compared. This gives a direct measure of the impact of futures trading on price discovery.

The aluminum market, which was dominated for most of the 20th century by a small number of transnational smelting companies who set prices on an administered (list) basis, is examined. Although list prices were clear and widely disseminated, there was substantial and variable (but secret) discounting from these prices (Radetzki, 1990, p.81). The list prices were therefore an unreliable guide to actual transaction prices. Exchange trading of aluminum started on the London Metal Exchange (LME) in October 1978, and although liquidity of the new contract was initially low, the industry effectively moved by the mid-1980s to pricing on the basis of these exchange prices. U.S. producers abandoned the practice of selling on a list-price basis at the end of 1985. Comex (now part of NYMEX) traded an aluminum contract from December 1983, but this contract failed to achieve liquidity and was abandoned in February 1989. A second Comex aluminum contract was launched in June 1999. The secondary objective of this article is to compare the informativeness of the LME aluminum contract with that of the two Comex contracts.

Prior to exchange trading of aluminum, reference prices were published in trade journals, based on averages or estimates of the prices at which recent transactions were made. These related to specific circumstances (location, grade etc.), which may not have been representative of the market as a whole. Sample sizes may have been small and reported transactions prices may therefore have been noisy. The prices initially referred to as relating to "certain other transactions" reported in the London *Metal Bulletin* are specifically considered. These prices were based on reported transactions prices for northern European delivery.

The focus of this paper is on the information content of the various prices (the producer list price, and the *Metal Bulletin,* LME, and Comex prices). The main tool used to analyze informativeness is the Beveridge-Nelson (BN, 1981) decomposition. The BN technique of decomposition of a nonstationary series into a random walk component and a transient, mean reverting component is used. If the price in question were generated within an efficient market, the transient component would be zero. The BN technique is therefore well adapted to the study of futures markets. The measure of the informativeness of each price series is the noise-to-signal ratio—the (square root of) the ratio of the variances of the transitory and permanent price components.

The BN procedure relies on the price series being nonstationary. Standard methods are used to test for this. Where there is more than a single price for the same period, it is required that these prices be cointegrated. If two prices relate to the same product, differences between them should be mean reverting (although not necessarily to zero); otherwise the prices would diverge increasingly over time, allowing limitless arbitrage possibilities. See also Gonzalo and Granger (1995), Hasbrouck (1995), and Baillie et al. (2002), who all require cointegration. Cointegration suggests a latent variable approach to modeling informativeness, which is also used.

THE ALUMINUM FUTURES MARKET

Aluminum is currently the most heavily traded nonferrous metal future. This may be seen from Figure 1, which shows the average daily volume and open interest for the six LME metals—aluminum (Al), copper (Cu), nickel (Ni), lead (Pb), tin (Sn), and zinc (Zn)—over the period January 1999 to June 2003. The figure aggregates contract numbers for the LME and Comex markets for aluminum and copper.[1] Although copper had historically been the most important of the LME contracts, LME aluminum open interest overtook that of copper in mid-1996 and volume did so in 1997.[2] The remaining four nonferrous metals are much less important on either criterion.

[1]Source: LME and NYMEX. The LME figures include options excluded from the Comex figures. LME and Comex contract sizes differ for aluminum and copper (25 tons, equal to 55,115 lbs, for the LME; 44,000 lbs for Comex) and across LME contracts (6 tons in nickel, 5 tons in tin, otherwise 25 tons). Aluminum alloy and silver are omitted, both of which either are or have been traded on the LME over the period concerned. Figures for tin are to December 2002. Aluminum is also traded on the Tokyo Commodity Exchange (TOCOM) and on the Shanghai Metal Exchange, which also trades copper.
[2]Briefly also in 1992 in each case.

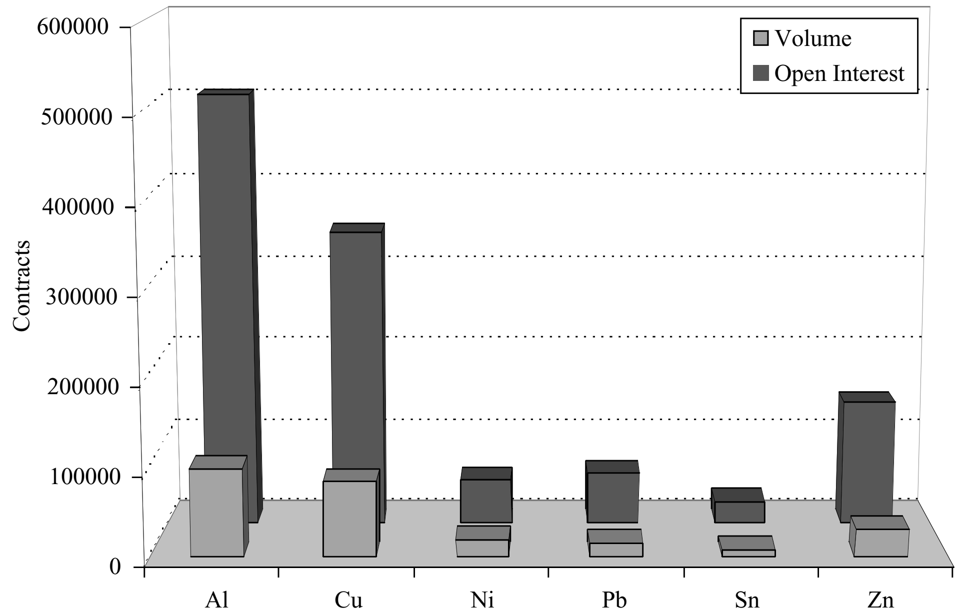

FIGURE 1
Average daily futures trading volume and open interest, nonferrous metals, January 1996–June 2003.

Figure 2 gives the same figures as a share of world refined metal consumption.[3] This shows futures trading in copper to be slightly more important than that in aluminum, although this may simply reflect offsetting positions on the two exchanges. Futures volume in tin also exceeds that in aluminum as a share of world consumption, and the figures for nickel and zinc are comparable. The much greater uniformity in Figure 2 relative to Figure 1 indicates that the importance of aluminum and copper in nonferrous metals trading is primarily due to the size of these two industries.

It is also interesting to compare trading activity in the 1985 Comex aluminum contract with that in the 1999 contract. The earliest year for which Comex open interest data could be obtained was 1987;[4] the last year in which the 1985 contract traded actively. Open interest in that year averaged 1471 contracts, slightly higher than the average in the initial months of the 1999 contract, but lower than subsequently. In the first six months of 2003, Comex aluminum open interest averaged 8820 contracts. Comparison between these two periods suggests that insufficient

[3]Source for consumption data: World Bureau of Metal Statistics, *World Metal Statistics* (various issues).
[4]Source: CFTC, *Commitments of Traders* reports, 1988. Unfortunately, the LME does not make available open interest figures prior to 1990.

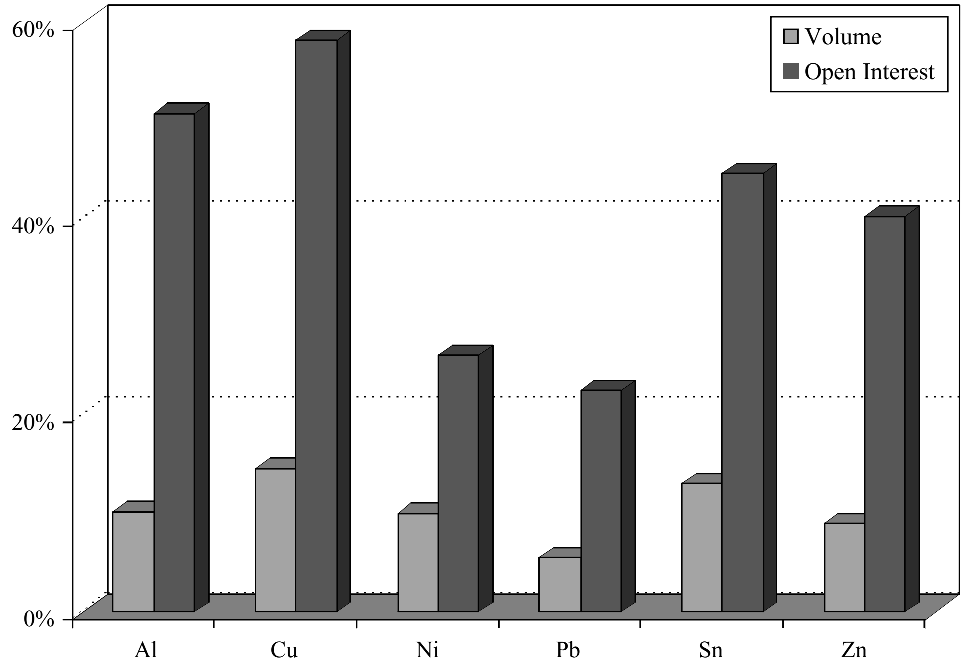

FIGURE 2
Average daily futures trading volume and open interest in relation to world refined consumption, nonferrous metals, January 1996–June 2003.

liquidity may indeed have accounted for the failure of the 1985 contract, but does not allow any inference as to why this was the case.

THE BEVERIDGE-NELSON DECOMPOSITION

Beveridge and Nelson (1981, henceforth BN) proposed a particular decomposition methodology for a nonstationary time series. This involved identification of the trend with the permanent component of the series with the consequence that the residual, which is by definition transitory, is identified as a combination of cycle and noise. If there are two alternative price measures, relating to either the same or different time periods, the series that exhibits the greater transitory variance can be considered less informative about the underlying trend, and in that sense, less efficient. On this view, transitory variance is associated with "noise."

BN noted that any I(1) series y_t may be decomposed into three components: a random walk μ_t, a stationary component e_t, and an initial condition $(y_0 - \mu_0)$—see also Hamilton (1994, p. 504). Consider the AR(m) representation

$$a(L)\,\Delta y_t = \varepsilon_t \tag{1}$$

where L is the lag operator, $a(L)$ is a lag polynomial or order m and the ε_t are IID by construction. Equation (1) can be inverted to obtain the infinite MA representation

$$\Delta y_t = \alpha(L)\varepsilon_t \tag{2}$$

where $\alpha_0 = 1$. This allows y_t to be written as

$$\begin{aligned} y_t &= y_0 + \varepsilon_t + (1 + \alpha_1)\varepsilon_{t-1} + \cdots + (1 + \alpha_1 + \cdots + \alpha_{t-1})\varepsilon_1 + \kappa \\ &= y_0 + \sum_{j=0}^{t-1}\left(\sum_{i=0}^{j}\alpha_i\right)\varepsilon_{t-j} + \kappa \end{aligned} \tag{3}$$

where κ depends on preinitial condition disturbances ε_t, with $t < 1$. The proposed decomposition is

$$y_t = \mu_t + e_t, \quad \text{where } \mu_t = \mu_{t-1} + \nu_t \quad \text{and} \quad E[\nu_t | y_{t-1}, \ldots, y_0] = 0 \tag{4}$$

This decomposition is achieved by setting

$$\nu_t = (1 + \alpha_1 + \cdots + \alpha_{t-1})\varepsilon_t \tag{5}$$

For sufficiently large t

$$\mu_t = \mu_0 + \alpha(1)\sum_{j=0}^{t-1}\varepsilon_{t-j}, \qquad \text{where } \alpha(1) = \sum_{i=1}^{\infty}\alpha_i \tag{6}$$

From Equations (2) and (3), the transient error e_t may be expressed as

$$\begin{aligned} e_t = y_t - \mu_t &= \sum_{j=0}^{t-1}\left[\sum_{i=0}^{j}\alpha_i - \sum_{i=0}^{t-1}\alpha_i\right]\varepsilon_{t-j} + (y_0 + \kappa - \mu_0) \\ &= -(\alpha_1 + \alpha_2 + \cdots + \alpha_{t-1})\varepsilon_t - (\alpha_2 + \cdots + \alpha_{t-1})\varepsilon_{t-1} \\ &\quad - \cdots - \alpha_{t-1}\varepsilon_2 + (y_0 - \mu_0) \\ &= \sum_{j=0}^{t-1}\gamma_j\varepsilon_{t-j} + (y_0 - \mu_0), \qquad \text{where } \gamma_j = -\sum_{i=j+1}^{t-1}\alpha_i \end{aligned} \tag{7}$$

For sufficiently large t, therefore, the transitory variance σ_e^2 is

$$\sigma_e^2 = \left(\sum_{j=0}^{\infty}\gamma_j^2\right)\sigma_\varepsilon^2 \tag{8}$$

where $\sigma_\varepsilon^2 = E\lfloor\varepsilon_t^2\rfloor$. The variance σ_p^2 of the change in the permanent component follows from Equations (4) and (5) as

$$\sigma_p^2 = \alpha(1)^2\sigma_\varepsilon^2 \tag{9}$$

It follows that the noise–signal ratio is given by

$$\frac{\sigma_e}{\sigma_p} = \frac{1}{\alpha(1)}\sqrt{\left(\sum_{j=0}^{\infty}\gamma_j^2\right)} \tag{10}$$

Further, from Equation (4)

$$\Delta y_t = \nu_t + \Delta e_t \tag{11}$$

and Equations (7) and (8) imply

$$Var(\Delta e_t) = 2\left(\sum_{j=0}^{\infty}\gamma_j^2 - \sum_{j=1}^{\infty}\gamma_j\gamma_{j-1}\right)\sigma_\varepsilon^2 \tag{12}$$

The ratio $\frac{Var(\Delta e_t)}{Var(\Delta y_t)}$ measures the proportion of the variance of the price change attributable to the transitory component of the price.[5] In what follows, the standard deviations of log prices are reported, as these are more easily interpreted than variances, and also the ratio of the standard deviation of the change in the transitory price component to that of the change in the log price [i.e., the square root of the expression in Equation (12)].

The interest here is in comparing the transitory variances of alternative price measures that relate, in some sense, to the same underlying price. This motivates multivariate generalization of the BN procedure in which the AR Equation (1) is replaced by the corresponding VAR (Vector AutoRegression); see, for example, Tse and Zabotina (2001). It is further required that the various prices entering the multivariate decomposition should be cointegrated, with the implication that divergences between these prices cannot grow indefinitely. For this to be the case, the decomposition should be applied to a CVAR (cointegrated VAR), which restricts the price series to have a single common trend. Although in principle these generalizations are straightforward, it was found that both the VAR and CVAR estimates suffered from a high degree of collinearity with the effect that the decomposition results were sensitive to sample and lag length specifications.[6] For this reason, the univariate BN decompositions were chosen. However, the analysis is restricted to those prices that are cointegrated.

[5]This ratio can exceed unity. Because $Cov(\nu_t, e_t) = \alpha(1)[\alpha(1) - 1]$, this will happen if $\alpha(1) > 1$.

[6]Tse and Zabotina (2001) estimated a bivariate VAR defined with respect to price changes and a bid–ask indicator. This specification would give rise to a low degree of collinearity, and their estimates should not be vulnerable to the problem discussed.

DATA

There are four sources of price data:

1. North American aluminum producer list prices are available on monthly basis from 1970 to the end of 1985 when the practice of issuing list prices was abandoned.[7]
2. The London *Metal Bulletin* (henceforth MB) transactions price series, originally referred to as relating to "certain other transactions," from January 1970 until February 1989. This price relates to European prices. The quotation basis for this series is not entirely consistent over time, and there are some gaps, particularly in the 1970s, during which the series was either not published, or published quotations were not revised from previous issues of the journal.
3. The complete daily LME cash settlement prices from October 1979 to December 2003.[8]
4. Comex aluminum settlement prices from December 1983 to February 1989, when the initial contract ceased trading,[9] and the settlement prices for the second Comex aluminum contract from June 1999 to December 2003.

For consistency, monthly averages of each of these prices are considered.

The four price series are charted in Figure 3. Movements in the producer price show little relationship with those in the MB price, and later, with the LME price. However, the MB, LME, and Comex prices do move closely over 1983–89, although Comex was less volatile, and again from 1999 where Comex has typically been at a premium to the LME.

ANALYSIS

Five periods are considered:

Period I: January 1970–December 1978. These were the closing years of the producer pricing period prior to the start of LME trading in aluminum.

Period II: January 1979–December 1983. This was the initial period of the LME aluminum contract.

[7]Source: *Non Ferrous Metals Data* (various issues).
[8]Sources: 1989–1988, World Bureau of Metal Statistics, *World Metals Statistics* (various issues); 1989–2003, http://www.lme.co.uk/HistData.htm.
[9]Source: Comex.

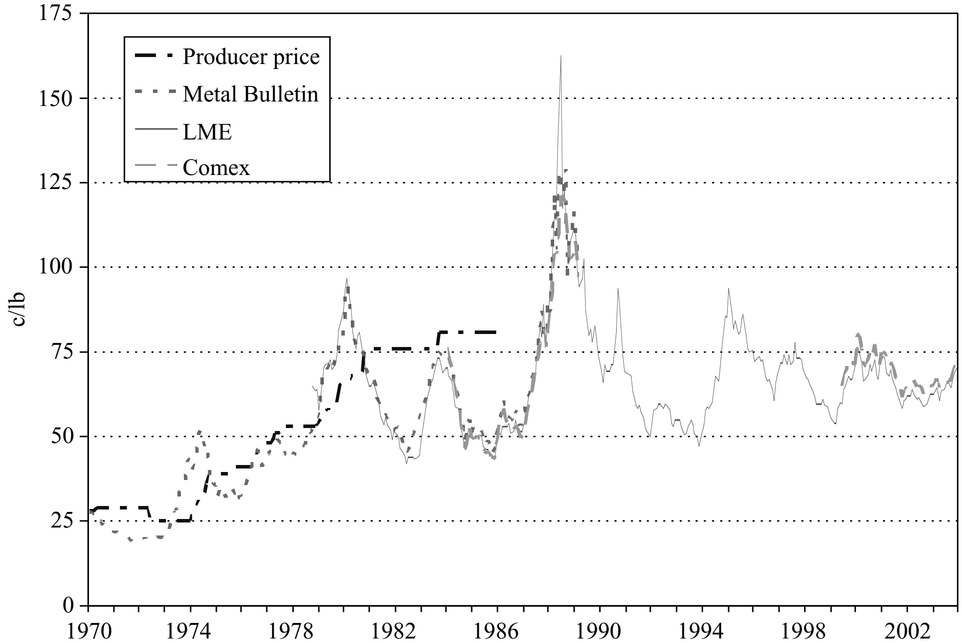

FIGURE 3
Aluminum prices, 1970–2003.

Period III: January 1984–February 1989. This was the period in which the LME competed with Comex in aluminum and succeeded in establishing itself as the pricing basis for world trade in aluminum.

Period IV: March 1989–May 1999. In this period, the LME was the clear pricing basis for world aluminum.

Period V: June 1999–December 2003. This period saw renewed competition between the LME and Comex in aluminum.

Table I lists the prices considered in each period. Producer prices after January 1984 are no longer considered, as Comex started to trade aluminum on 8 December 1983. Aluminum prices in the United States

TABLE I
Aluminium Price Regimes

	Producer price (PP)	Metal Bulletin *price (MB)*	*LME (LM)*	*Comex (CM)*
Period I: Jan 1970–Dec 1978	Yes	Yes	No	No
Period II: Jan 1979–Dec 1983	Yes	Yes	Yes	No
Period III: Jan 1984–Feb 1989	No	Yes	Yes	Yes
Period IV: Mar 1989–May 1999	No	No	Yes	No
Period V: June 1999–Dec 2003	No	No	Yes	Yes

Note. The table lists the five pricing regimes used in the Beveridge-Nelson decompositions that follow.

TABLE II
ADF(3) Test Statistics

Period	*Producer price*	Metal Bulletin *price*	*LME*	*Comex*	*95% critical value*
I	−0.23	−0.91			−2.89
II	−2.46	−1.82	−1.46		−2.91
III		−0.83	−0.88	−1.04	−2.91
IV			−2.45		−2.88
V			−1.41	−1.22	−2.92

Note. Tests are performed for the price-regime combinations listed in Table I. A test statistic with a negative value lower than the stated critical value would reject the null hypothesis of nonstationarity in favor of stationarity.

effectively moved onto an exchange basis. The initial period of Comex aluminum trading terminates in February 1989–28 February 1989 was the last date on which the CFTC published open interest and commitment of traders figures for that contract. The MB price after the end of Period III in February 1989 is also not considered, because by that time European aluminum was transacted at a price related to the LME price by a differential for European delivery.

First the integration and cointegration properties of the various prices in each regime are established. In each case, the logarithm of the (undeflated) price is considered. Table II lists ADF(3) nonstationarity test statistics, together with the 95% critical values. A test statistic with a lower negative value than the listed critical value would reject the null hypothesis of nonstationarity in favor of stationarity. None of the listed statistics gives a rejection, and the tests are therefore consistent with each of the series being nonstationary in each regime.[10] Although unsurprising, this outcome is important in that it allows tests of cointegration to commence.

Cointegration is important because two (or more) series that measure the same price must be cointegrated. This is because differences between different measures of the same underlying price cannot diverge indefinitely. Table III gives the Johansen cointegration test results for the prices listed in Table I—see Johansen (1988) and Johansen and Juselius (1990). The first two columns of the table give the test outcomes for the null that both prices (Periods I and V) and all three pieces (Periods II and III) are cointegrated. The hypothesis is rejected for Periods II and III, but not for Periods I and V.

[10]The same results obtain using different lag lengths. Although shorter lag specification increases test power, it would leave open the question of pretest bias. Therefore a uniform and conservative lag specification is preferred.

TABLE III

Johansen Cointegration Test Statistics

Period	*All prices cointegrated*	*Tail probability*	*Price pairs*	*Test statistic*	*Tail probability*
I	$\chi^2(1) = 0.18$	67.5%			
			PP,MB	$\chi^2(1) = 6.01$	1.42%
II	$\chi^2(4) = 13.8$	0.79%	PP,LM	$\chi^2(1) = 6.47$	1.10%
			MB,LM	$\chi^2(1) = 2.21$	13.7%
			MB,LM	$\chi^2(1) = 0.58$	44.6%
III	$\chi^2(4) = 22.7$	0.01%	MB,CM	$\chi^2(1) = 0.70$	40.4%
			LM,CM	$\chi^2(1) = 0.62$	43.3%
V	$\chi^2(1) = 2.30$	12.9%			

Note. The table gives the Johansen reduced rank test statistics for VAR(3) models for the variables listed in Table I. The initial two columns include all the prices considered in the regime in question, and the final column gives the pairwise comparisons for Periods II and III. No cointegration test is given for Period IV when only a single price is considered. A tail probability of less than 5% indicates rejection of the null of cointegration.

For Periods II and III pairwise comparisons are also used. Period II gives an unambiguous result—the MB and LME prices are cointegrated, but neither is cointegrated with the producer price. It is concluded that, over this period, the producer price was no longer providing an indication of aluminum transactions prices; it is therefore omitted from the decompositions that follow. The results for Period III are less clear. The data clearly reject the null of two cointegrating vectors, but are compatible with each price pair being cointegrated. These test outcomes do not provide any basis for concluding that any single one of these prices is responsible for the rejection of two cointegrating vectors. Therefore all three prices in this regime continue to be considered.

The discussion now turns to the BN decomposition results given in Table IV. The table lists the standard deviation of the (logarithmic) price changes, the standard deviation of the transitory component of the log price, the standard deviation of the first difference of these transitory components, and the last of these as a percentage of the first.

First comparisons across the alternative prices in each pricing regime are considered.

Period I was the producer pricing period, prior to the introduction of futures trading on the LME. The log standard deviation of the producer price was around half that of the MB transactions price. The transitory component of the producer price was less than half that of the MB price, and changes in the MB price were predominantly transitory. In this period, the producer price appears to have provided

TABLE IV
Beveridge-Nelson Decomposition Results

Period		*SD of price change*	*SD of transitory component*	*SD of change in transitory component*	*Percentage transitory*
I	Producer price	2.54%	2.72%	1.89%	74.4%
	Metal Bulletin	5.14%	6.28%	4.63%	90.1%
II	*Metal Bulletin*	4.54%	7.07%	5.13%	113.0%
	LME	5.07%	4.36%	3.95%	77.9%
III	*Metal Bulletin*	8.10%	4.52%	3.37%	41.6%
	LME	8.27%	4.17%	2.85%	34.5%
	Comex	5.56%	7.12%	5.46%	98.2%
IV	LME	5.17%	1.08%	1.62%	31.3%
V	LME	3.48%	1.37%	1.26%	36.2%
	Comex	3.25%	1.41%	1.31%	40.4%

a more accurate measure of the underlying price than the MB transactions price reported.[11]

Period II corresponds to the initial years during which the LME had started trading aluminum. Note that, over this period, the producer price was not cointegrated with the LME exchange price of the MB transaction price, and therefore cannot be taken as measuring the same underlying price. Effectively, the existence of a futures market made the producer price irrelevant as pricing moved to be basis the futures price and not, as previously, the producer list price. Comparison of the MB and LME prices shows that the latter had the substantially lower transitory variance. Even in these initial years, therefore, the LME gave a superior guide to the underlying price than did the MB transactions price.[12]

Period III relates to the period over which the initial Comex aluminum contract traded. The Beveridge-Nelson decomposition shows that

[11]This result is subject to the qualification that the BN technique assumes homoscedasticity and is therefore not entirely appropriate for data, such as the producer list price, which change only occasionally. Further, the conclusion does not have the implication that the producer price gave a more accurate representation of the price that producers could expect to obtain, as distinct from ask, and that consumers could have expected to pay in any period. It does imply that the producer price provided superior information on the underlying price trend.

[12]The MB price provides an instance in this period in which the standard deviation of the change in the transitory price component exceeds that of the price itself—see footnote 5. Premia over or discounts under the exchange price may have been serially correlated. It would therefore be incorrect to take this result as implying that the exchange price provided a better guide to transactions prices than the MB price. As in the previous footnote, the implication is for informativeness about the underlying price trend.

the Comex price had a substantially higher transitory variance than either the LME or the MB prices. This is despite the much lower variability of the Comex price over this period, as most of this variability turns out to have been transitory. The comparison of the LME and MB prices is similar to the same comparison in Period II, although the two transitory variances were much closer in the later period.

Period V again saw the LME and Comex competing against each other. The Comex price has been slightly less variable than the LME price, but a higher proportion of this variability is seen as transitory. The net result is that the transitory variances of the two prices are nearly indistinguishable. Effectively, both prices have been equally informative about underlying price developments.

Comparisons over time are more problematic, as they require controls for movements of volatility over time. It is notable, however, that the proportion of the price-change standard deviation due to transitory movements has declined from over 70% in the 1970s and early 80s to between 30% and 40% since the mid-1980s. This signifies a major increase in market transparency. Although it is not possible to infer that this increased transparency was caused by the advent of futures trading in aluminum, it is sufficient to note that the increased transparency has occurred as futures trading became liquid.

THE LME AND *METAL BULLETIN* PRICES

In the previous section it was argued, on the basis of the BN decomposition, that the LME price was consistently more informative about the fundamental price of aluminum than was the MB transactions price. This judgment was based on the lower transient variance of the exchange price. However, that analysis suffered from the limitation of treating each period as homogeneous. In this section, a latent variable procedure is used to examine the evolution of the relative informativeness of the two prices over time. The procedure used is a variant of the method due to Griliches (1977) and Zellner (1977). It is, however, not possible to apply this technique to the Comex price because of the relatively short periods over which Comex has traded aluminum.

On the latent-variable hypothesis, both the exchange and the trade journal prices are regarded as measuring this latent price subject to a measurement error. The analogy is with Friedman's permanent income model of consumption expenditures (Friedman, 1957) in which measured income differs from permanent income by transitory income, which may be regarded as a measurement error. This approach allows the relative importance of these two measures to evolve over time.

Let y_t now be the vector consisting of the logarithms of the LME and MB prices. The cointegrated vector error-correction model (VECM)

$$\Delta y_t = a_0 + \sum_{i=1}^{3} A_i \Delta y_{t-i} + a_4 (1\ -1) y_{t-4} + e_t \tag{13}$$

is estimated, where a_0 and a_4 are 2×1 vectors and $A_i (i = 1, \ldots, 4)$ are 2×2 matrices. The term $(1\ -1)y_{t-4}$ is the lagged log difference of the two prices (the error-correction term) and imposes a unit cointegrating vector on the joint price process. As noted above, cointegration requires that divergences between the two prices should be mean reverting, and is equivalent to stating that the two prices both measure the same underlying fundamental.

The VECM specified in Equation (13) is estimated by ordinary least squares over the complete sample for which there is information on both prices, i.e. Periods II and III combined (February 1979–February 1989, where one extra month has been lost through lagging). The estimated model is listed in Table V. Heteroscedasticity-robust t statistics are

TABLE V

Estimated VECM

Dependent Variable	Δ ln MB$_t$		Δ ln LME$_t$	
Regressors	Δ ln MB	Δ ln LME	Δ ln MB	Δ ln LME
Lag 1	0.1224	−0.0484	0.7583	−0.3786
	(0.80)	(0.37)	(4.24)	(2.07)
Lag 2	−0.1297	0.4028	0.2224	−0.1410
	(0.88)	(2.83)	(1.18)	(0.66)
Lag 3	0.2587	−0.3230	0.6752	−0.5186
	(1.50)	(2.15)	(2.45)	(2.56)
Intercept	0.0040 (0.58)		−0.0079 (1.00)	
ln MB$_{t-4}$ − ln LME$_{t-4}$	−0.0300 (0.28)		0.2157 (1.63)	
R^2	0.5171		0.6197	
Standard error	5.74%		5.51%	
Residual correlation		0.6926		
Residual diagnostics				
LM(7) serial correlation	$F(7,106) = 0.71$ [66.5%]		$F(7,106) = 1.37$ [22.5%]	
LM heteroscedasticity	$F(14,98) = 2.49$ [0.46%]		$F(14,98) = 6.93$ [0.00%]	
Normality	$\chi^2(2) = 1.82$ [40.4%]		$\chi^2(2) = 9.57$ [0.83%]	

Note. Equation specification: (13). Sample: February 1979 to February 1989.
Heteroscedasticity-consistent t statistics in parentheses; tail probabilities in brackets.

given in parentheses. The equations are well determined. Although the error-correction terms are not individually significant, a Wald test of their joint significance gives a value of $\chi^2(2) = 11.1$, associated with a tail probability of 0.4%. The interest will be in the error terms of the VECM. The standard LM test confirms that the residuals are serially independent. The presence of heteroscedasticity is to be expected and this enters into the subsequent modeling exercise.

The latent variable procedure decomposes the error term e_t in Equation (12) into a common (scalar) component z_t and a 2×1 vector of idiosyncratic errors ε_t:

$$e_t = \begin{pmatrix} 1 \\ \lambda \end{pmatrix} z_t + \varepsilon_t \tag{14}$$

This will be recognized as a single factor model within the factor-analysis class.[13] Because idiosyncratic factors are definitionally uncorrelated, one may write $E(\varepsilon_t \varepsilon_t') = \begin{pmatrix} \omega_1^2 & 0 \\ 0 & \omega_2^2 \end{pmatrix}$ and $E(z_t^2) = \zeta^2$. Again, by definition the common factor z_t is independent of the idiosyncratic factors ε_t, and both inherit serial independence from the CVAR. It follows that

$$E(e_t e_t') = \begin{pmatrix} \sigma_1^2 & \rho\sigma_1\sigma_2 \\ \rho\sigma_1\sigma_2 & \sigma_2^2 \end{pmatrix} = \begin{pmatrix} \omega_1^2 + \zeta^2 & \lambda\zeta^2 \\ \lambda\zeta^2 & \omega_2^2 + \lambda^2\zeta^2 \end{pmatrix} \tag{15}$$

Define the noise-to-signal ratios $\theta_j = \omega_j^2/\eta_j^2\zeta^2$ $(j = 1, 2)$, where $\eta = \begin{pmatrix} 1 \\ \lambda \end{pmatrix}$. Equation (15) relates the three observed quantities in the estimated error variance matrix to four unknowns $(\omega_1^2, \omega_2^2, \zeta^2, \text{and } \lambda)$. Solving leads to

$$\theta_1 = \frac{\lambda\sigma_1^2 - \rho\sigma_1\sigma_2}{\rho\sigma_1\sigma_2} \quad \text{and} \quad \theta_2 = \frac{\sigma_2^2 - \lambda\rho\sigma_1\sigma_2}{\lambda\rho\sigma_1\sigma_2} \tag{16}$$

Equation (16) leaves the noise–signal ratios unidentified, but it is shown below how the estimates can nevertheless be bounded.

With the use of the residuals from the VECM the variances σ_1^2, σ_2^2 and ρ are estimated with the use of centered 25-month moving averages. Conditional on a value for λ, this would allow the two noise-to-signal ratios, θ_1 and θ_2, to be estimated. Considering the complete set of rolling estimates, the restriction $\lambda\sigma_1^2 - \rho\sigma_1\sigma_2 > 0$ forces a value of $\lambda > 0.655$,

[13]In a variation on standard factor analysis notation the first factor loading is normalized instead of the variance of the common factor. This allows a variance estimate for the common factor (relative to this normalization) to be recovered.

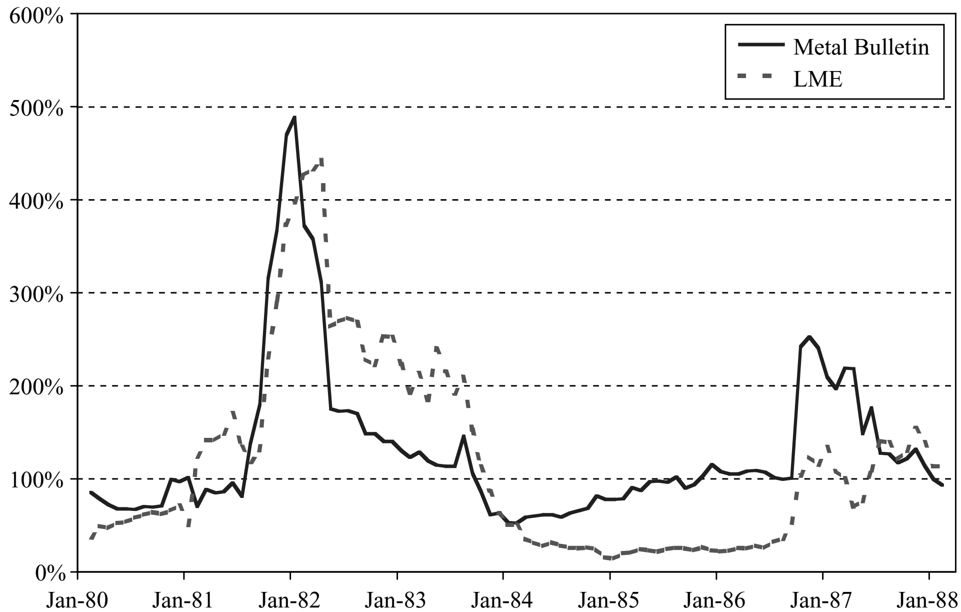

FIGURE 4
Estimated noise-to-signal ratios.

whereas the restriction $\sigma_2^2 - \lambda\rho\sigma_1\sigma_2 > 0$ forces a value of $\lambda < 0.875$. Therefore λ was set at its midpoint of 0.765. Figure 4 plots the resulting noise–signal ratios. Although the precise numbers differ, the general story told by the estimates remains the same throughout the admissible range $0.765 < \lambda < 0.875$.

Figure 4 shows that, if this model is accepted as a valid representation of the relationship between the two prices, the LME and the MB prices had comparable noise-to-signal rations over the initial years of LME trading in aluminum futures to 1982, and the LME may have been more noisy than the MB price in 1983 and 1984. However, from 1984 through to 1987, the LME price became much less noisy both absolutely and relative to the MB price. Then, from 1987 as metals price volatility increased sharply, both prices showed increased noisiness.

THE LME AND COMEX PRICES

In this section, the relationship between the two exchange prices over the two periods (III and V) in which Comex has traded aluminum is briefly considered. Whether either of the two markets can be regarded as leading the other is also examined. For consistency with the earlier discussion, the monthly average data used in the earlier analysis of the relationship of the exchange prices to the MB transaction price are used.

However, a comprehensive discussion of the relationship between the two prices would need to employ data on either a daily or an intradaily frequency.

The first tool used is that of Granger causality tests, introduced by Granger (1969)—see Hendry (1995, p. 179) for a discussion. A variable z is said to Granger-cause a second variable y if, knowing the history of y, knowledge of the history of z assists in predicting y. This test is implemented by regression of y on lag distributions of itself and z:

$$y_t = \alpha_0 + \sum_{i=1}^{k} \alpha_i y_{t-i} + \sum_{i=1}^{k} \beta_i z_{t-i} + \varepsilon_t \tag{17}$$

The test is then the standard Wald exclusion test on the lag distribution x:

$$H_0 : \beta_1 = \cdots = \beta_k = 0 \tag{18}$$

In the tests used z and y are the monthly changes in the logarithms of the average Comex and LME prices.

The samples available to us are short (59 observations in Period III and 52 in Period V prior to loss of observations through differencing and lagging). In order for the tests to have power, the length of k, the lag distributions, needs to be severely curtailed. Tests using lag lengths of $k = 2$ − longer lag distributions give inconclusive results, while restriction to a lag length of 1 results in misspecification. Results are reported in Table VI. The notation $C \not\rightarrow L$ is to be read as “changes in Comex prices do not Granger-cause changes in LME prices” and correspondingly for the notation $L \not\rightarrow C$. In either case, a significant test statistic, indicated by a tail probability of less than 5%, rejects the null of “Granger noncausality,” establishing the presence of a Granger-causal link.

The tests indicate that in Period III, the Comex and LME price processes could be regarded as having been independent. In reality, this

TABLE VI
Granger Causality Tests

	Period III	*Period V*
$C \not\rightarrow L$	$F(2,54) = 1.46$ [24.2%]	$F(2,47) = 0.38$ [68.3%]
$L \not\rightarrow C$	$F(2,54) = 0.28$ [75.4%]	$F(2,47) = 3.43$ [4.1%]

Note. The notation $A \not\rightarrow B$ is to be read as “changes in A prices do not Granger-cause changes in B prices.” Test statistics relate to the null hypothesis (18) defined in relation to Equation (17). Tail probabilities are given in brackets.

may simply imply that the data do not show sufficient variability to establish the directions in which the prices influenced each other. By contrast, the test results for Period V show LME prices as Granger-causing Comex prices but not *vice versa*. This suggests that one may regard the LME as being the price leader over this, more recent, period.

In an earlier section it was shown that the LME and Comex prices were cointegrated in Period V and that the evidence was also consistent with their being cointegrated in Period III. The Granger-causality tests fail to account for cointegration. To look further at the issue of leadership, the Johansen reduced rank regression procedure is used. Let x be the vector consisting of the log LME and Comex prices; that is, $x = \begin{pmatrix} y \\ z \end{pmatrix}$, in the notation of Equation (17). The VECM($k + 1$) for x written in the form

$$\Delta x_t = \gamma_0 + \sum_{i=0}^{k} \gamma_i \, \Delta x_{t-i} + \alpha\beta' x_{t-k-1} + \varepsilon_t \tag{19}$$

is considered. Attention focuses on the error-correction matrix $\alpha\beta'$, which forces cointegration of the components of x. In this case, with two variables, $\alpha\beta'$ has the reduced rank of one; that is, α and β are each 2×1 vectors and $x\beta$ defines the error-correction term. By convention, the first component of β is normalized to unity. If the LME and Comex prices are measuring the same fundamental, the log premium of the LME price over Comex should be expected to tend to a constant. If this is the case, $\beta = \begin{pmatrix} 1 \\ -1 \end{pmatrix}$. Within this framework, the estimated α coefficients are informative about the extent to which each of the prices reacts to divergences between them, that is, to widening or narrowing of the LME premium over Comex.

For consistency with the Granger-causality results, reduced rank VAR(3) models for each of Periods III and V are estimated; that is $k = 2$ is set in Equation (19). Estimation results are reported in Table VII. The hypothesis $\beta = \begin{pmatrix} 1 \\ -1 \end{pmatrix}$ is rejected for Period III but accepted for Period V. The table therefore reports the unrestricted estimates of α and β for Period III and the restricted estimates for Period V. The estimated α coefficients are of comparable magnitude in Period III, but in Period V the reaction almost all appears to take place in the LME price.

Looking first at in Period III, the failure to accept homogeneity of the two exchange prices casts further doubt on the reliability of the Comex price over that period and is in line with the ambiguous results obtained earlier. Turning to Period V, the Granger-causality tests suggest

TABLE VII
Reduced Rank Regression Results

	Period III		*Period V*	
	ln LME	ln Comex	ln LME	ln Comex
α	−0.2674	0.2330	−0.3636	0.0244
β	1.0000	−1.0993	1.0000	−1.0000
H_0: $\beta' = (1, -1)$	$\chi^2(1) = 5.66$ [1.7%]		$\chi^2(1) = 0.59$ [44.4%]	

Note. The table reports the estimated α and β coefficients from the reduced rank regression (19) with $k = 2$. The first coefficient of β is normalized to unity. The likelihood ratio test of $\beta' = (1, -1)$ relates to the unrestricted regression. Tail probabilities are given in brackets. The table reports the unrestricted regression for Period III and the restricted regression for Period V.

that it is the LME that leads Comex, whereas the reduced rank regression results show that it is largely the LME that accommodates to widening or narrowing of its premium over Comex. This is the precise situation anticipated by Bailie et al. (2002, p. 310). According to the PT methodology, price discovery should be ascribed purely to Comex. This conclusion is paradoxical in the light of the much lower trading volume and open interest on Comex—see the second section. By contrast, the IS methodology would take the high contemporaneous error correlation as implying that each market contributes to price discovery. Within that framework, it is not possible on the basis of the price data alone to state that one market is more important than the other. However, the volume and open interest information would indicate priority of the LME. However, a comprehensive discussion of this relationship would require a move to higher frequency data.

CONCLUSIONS

An attempt has been made to quantify the relative information content of the different exchange and nonexchange pricing regimes that have been prevalent in the aluminum industry. Specific attention was paid to U.S. producer prices until their abandonment in the mid-1980s, transactions prices published in a trade journal (the *Metal Bulletin,* MB), and LME and Comex exchange prices over the periods which they were traded.

Prior to the advent of exchange trading of aluminum, producer prices were more informative than reported transactions prices. With the start of exchange trading on the LME, the producer price quickly lost its information content. These estimates show that the information content of the MB transactions price increased once futures trading started, although the LME price was the more informative of the two prices, at least after the initial 2–3 years of LME trading. The answer to the principal question posed in the introduction is therefore that futures trading in aluminum was clearly associated with an increase in the transparency of transactions prices in the market for physical aluminum.

In the mid-1980s, the LME faced competition in aluminum from Comex, but the Comex contract failed to attract liquidity and was less informative than either the LME or MB prices. Comex trading is a new aluminum contract started in 1999, and this contract has been more successful. These estimates show the Comex and LME markets to be equally transparent and that information appears to flow in both directions between the two markets. However, there is limited evidence of the LME acting as price leader.

These results go beyond those of many previous articles that have tackled price-discovery issues in futures markets by utilizing a data set that covers prices both before and after the introduction of futures trading. They illustrate how futures prices increase the information content in transaction prices such that, eventually, the transaction prices move to a futures basis. Aluminum has been one of the greatest successes for commodity futures trading over the past two decades. Although it would be optimistic to expect this degree of success to be replicated in every new commodity futures market, this increased market transparency underlines the promise that commodity futures trading holds.

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