# Price Discovery in the European Aluminium Market

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**Abstract**

We use an extended version of the Beveridge-Nelson (1981) decomposition and the Kalman filter to examine how the noise content, and therefore the informativeness, of the LME aluminium price has evolved since the start of aluminium trading on the LME in 1978. We compare the LME price with the reference price published in a trade journal. We also consider whether futures trading in aluminium has been associated with any change in price volatility.

Key words: Price discovery, price volatility, commodity futures

# Price Discovery in the European Aluminium Market

# 1. Introduction

Price discovery is one of the most important functions provided by organized commodity exchanges. In the absence of an exchange, transactions will be made at different prices by different parties depending in part on the extent to which they are informed about market conditions and the extent to which they can find alternative counterparties. Once liquid trading takes place on an organized exchange, the exchange price becomes a common reference price for all transactions. Differences may exist between the prices of different transaction, but these will now be clearly related to location, grade, delivery conditions and other specifiable factors. 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).

This ideal is not always achieved. One reason for this is that speculative activity can induce some froth into the price process. Although informed speculation must be expected to be stabilizing, if noise traders are present in the market, it will not in general be possible to know whether a particular price movement originates from an informed or an uniformed trade. If informed traders have sufficiently long horizons, they will trade in a contrarian manner, but if they have short horizons, they may trade on a momentum basis, reinforcing what started as an arbitrary price movement (de Long *et al*, 1990; Shleifer and Vishny, 1997).

In the absence of exchange trading, industries develop less formal methods for gathering and disseminating price information. In particular, trade journals publish price information and market participants will be aware of these reported prices when they make their own transactions. Typically, this information will consist of averages or estimates of the prices at which recent transactions have been made. Nevertheless, the information content of these reported prices is not always clear. Even supposing that transaction prices are reported or estimated accurately, these transactions will relate to specific circumstances (location, grade etc) and may not be representative. Averaging will typically take place over prices relating to different circumstances on different occasions. Sample sizes may be small. Prices reported in trade journals will therefore also tend to be noisy.

Economists would generally expect a liquid market to generate more informed prices than those reported in trade journals in the absence of a market. However, it may take time for a new market to attract liquidity and many markets fail to take off. In any particular case, it must therefore be an empirical issue whether the exchange price or prices published in trade journals are more informative.

**Table 1: LME futures 3rd Wednesday Open interest**[1]

| | Aluminium | Copper | Zinc | Nickel | Lead |
|---|---|---|---|---|---|
| **1990** | 86,368 | 86,368 | 18,778 | 7,856 | 10,399 |
| **1992** | 167,303 | 127,040 | 58,739 | 27,209 | 20,877 |
| **1995** | 205,362 | 212,545 | 84,476 | 47,067 | 32,818 |
| **1997** | 258,091 | 144,022 | 86,.342 | 52,417 | 34,505 |
| **2000** | 302,455 | 180,724 | 95,590 | 50,158 | 34,719 |
| **2002** | 320,527 | 203,692 | 128,889 | 42,620 | 40,381 |

**Table 2: Turnover figures for LME contracts (lots)**

| | Aluminium | Copper | Zinc | Nickel | Lead |
|---|---|---|---|---|---|
| **1990** | 15,897 | 22,584 | 5,646 | 5,446 | 14,452. |
| **1992** | 32,385 | 28,890 | 15,839 | 5,679 | 9,230 |
| **1995** | 55,794 | 69,564 | 20,801 | 13,173 | 6,979. |
| **1997** | 88,870 | 59,683 | 29,211 | 18,292 | 9,299 |
| **2000** | 100,968 | 69,703 | 29,956 | 20,344 | 12,788 |
| **2002** | 88,966 | 66,052 | 32,271 | 12,698 | 13,590 |

In this paper, we consider the aluminium market which was dominated for most of the twentieth century by a small number of transnational smelting companies who set prices on an administered (list) basis. 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. However, reference prices were published in trade journals. Exchange trading of aluminium started on the London Metal Exchange (LME) in

[1]L ME open interest figures are based on the sum of all net long or all net short positions held by Clearing Members at the London Clearing House in respect of each forward delivery date, and do not include figures in respect of client positions.

October 1978 and although the initial liquidity of the new contract was low, the industry effectively moved by the mid nineteen eighties to pricing on the basis of these exchange prices. US producers abandoned the practice of selling on a list price basis in 1986. Tables 1 and 2 report average open interest and turn over figures for six chosen years over the 1990-2002 period.[2] They show that the aluminium contract is now, and has for some time, been the highest volume contract on the LME both in terms of number of contracts and of value. Copper is the second most important contract followed by zinc. Aluminium therefore provides a natural experiment allowing us to examine the effects of exchange trading on price informativeness.

Specifically, we consider the European aluminium market. Prior to trading of aluminium on the LME, market participants often made reference to the "certain other transactions" price published in the twice weekly London *Metal Bulletin*, and we compare the information content of this contract with that of the LME price.[3] The objective is to judge

- the relative informativeness of the LME and *Metal Bulletin* aluminium price series; and
- how this informativeness has evolved over time, in particular in relation to the growth in the importance of aluminium trading on the LME.

Price discovery is the process by which information becomes impounded in prices. Although we do not observe this directly, evolution of the degree of noisiness of the exchange and trade journal prices over time will allow us to judge how successfully that process has taken place.

We measure informativeness via the Beveridge-Nelson (1981) decomposition of the variance of price changes (henceforth, BN). BN show how a non-stationary time series may be decomposed into a permanent and a transient component with the properties that the permanent component purely represents the stochastic trend. Changes in the permanent component are therefore unforecastable. In this sense, we may interpret the transient component of a price series as a measure of its noisiness.

[2] Average Turnover figures were calculated from daily turnover data and open interest figures were calculated from monthly open interest data (source www.lme.co.uk)

[3] Radetzki (1990, p.81) states that publication of this price became discontinued since the introduction of the aluminium LME contract in 1979. In fact, publication continued but the term "certain other transactions" was dropped. There has never been a successful aluminium contract on a US exchange and transactions in north America are either based on the LME price or on the daily marker price published in *Metals Week*

The BN transient component will in general be serially correlated and therefore forecastable. To say that a series is noisy is not to imply that the noise component is uninteresting. Indeed, from the point of view of a trader, it is just the autocorrelation properties of the noise which allow her to make profits. But the noise component will also obscure the fundamental value of the commodity, which we can interpret in terms of the permanent component of the series.

In order to implement the BN procedure, we generalize their account in two respects:

- We implement the BN procedure in the context of a bivariate Vector AutoRegression (VAR) defined in terms of the exchange price and the trade journal price.
- We modify the VAR to a cointegrated VAR (ie a CVAR) to take into account the cointegration of the two price series.

The BN decomposition requires that the price trend should follow a random walk. This is attractive in terms of an efficiency interpretation of the trend but may at the same time be considered over-restrictive.

The latent variable approach decomposes a series, or in this case a vector process, into latent variable and a cyclical component. In our case, because the two price series are cointegrated, we require the latent variable to be common. By contrast with BN, the latent variable decomposition is performed recursively. Although the latent variable unforecastable *ex ante*, it may appear to have been forecastable *ex post*. A benefit of the latent variable approach is that we can retrieve a time series of estimates of the transitory variance.

Exchange trading in aluminium is now established and, to that extent, our discussion might appear to be now only of historical interests. Such an impression would be incorrect since the same issues arise in other markets in which exchange trading is not currently established. At the time of writing, the LME is giving active consideration to the introduction of a contract in Hot Rolled Coil (HRC) steel. Steel is currently prices on a list basis by steel producers, and there is a number of alternative prices reported in the trade press and by consulting companies to which the industry makes reference. In this respect, the steel market has obvious parallels with aluminium three decades ago. Proponents of exchange trading of steel make frequent reference to the success of the LME aluminium contract and suggest that this may be replicated in steel. We do not take a view on that issue in this paper, except to claim that a thorough understanding of changes in aluminium price discovery will better inform the continuing steel debate.

The remainder of this paper falls into seven sections. In Section 2 we discuss the price data and look at its statistical properties. In Section 3, we use all the available aluminium price series to ask whether the move to exchange trading of aluminium increased or reduced the variability of aluminium transactions prices. After this preliminary, we turn to price discovery. In section 4, we develop and apply the BN methodology within the CVAR framework. Then in section 5, we perform the moving window analyis. Section 6 concludes.

## 2. Data

We have a complete daily series of LME cash settlement prices from July 1979 to June 2000.[4] Aluminium was also traded in New York on Comex (now part of NYMEX) over the period 1982-1996. This gives us a second series of daily quotations over a subset of our sample.[5] Producer list prices are available on a monthly basis from 1970 to the end of 1985, when the practice of issuing list prices was abandoned.[6] Finally, we have two sources of data for transaction prices, the first relating to the European and the second to the North American market.

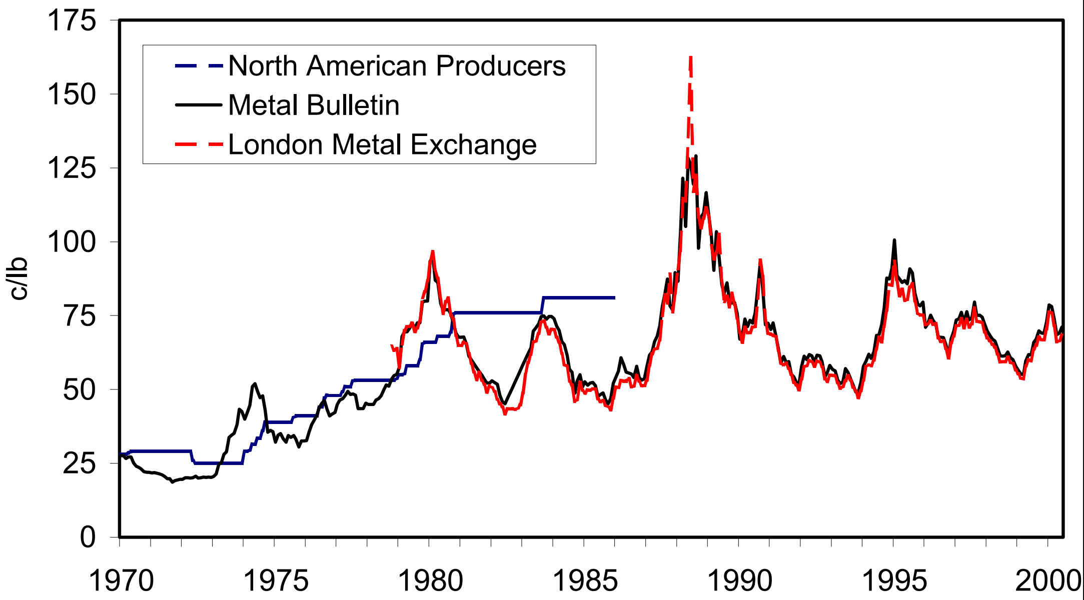
For Europe, transactions prices are quoted, originally as relating to "certain other transactions", in the twice weekly trade journal *Metal Bulletin*. This has allowed us to construct a series of monthly averages covering the period 1970-99.[7] The quotation basis is not entirely consistent over time, and there are some gaps, particularly in the nineteen seventies, during which the series was either not quoted or quotations were not revised. For North America, we have constructed a daily transaction price series over the period 1985 to 1997 from the trade journal *Metals Week*. This series has the advantage that it is available on a consistent daily basis, while the European series is intermittent and of variable consistency, but it was not

[4] Sources: 1979-1988 - World Bureau of Metal Statistics, *World Metals Statistics* (various issues); 1989-2000 – http://www.lme.co.uk/HistData.htm .
[5] Source: Comex.
[6] Source: *Non Ferrous Metals Data* (various issues).
[7] We gratefully acknowledge financial support for the data compilation from the Social Science Faculty Research Fund, Queen Mary, University of London.

quoted during the producer pricing period of the nineteen seventies and early eighties. The European series is therefore potentially the more informative in relation to the impact of exchange trading on the variability of aluminium prices.

In Figure 1 we plot monthly averages of the north American producer list prices, the LME aluminium prices and the *Metal Bulletin* transactions prices over the full sample, January 1970 to June 2000. These graphs indicate the LME price is clearly more variable than the two producer prices. However, it is not immediately apparent whether the variability of the transactions price has or has not increased.

**Figure 1**
**Aluminium Prices, 1970-2000**




In sections 4 and 5, we shall be specifically concerned with the LME and *Metal Bulletin* price series which both relate to the European market. They are also both available from the nineteen seventies through to the end of the sample. We first establish that both these prices are non-stationary, and then ask whether they are cointegrated. Cointegration is necessary if we are to regard the two series as measuring the same underlying price.

| Table 3: Stationarity and Cointegration Tests | | | | | |
|---|---|---|---|---|---|
| | **ln*LME*** | **Δln*LME*** | **ln*MB*** | **Δln*MB*** | **ln(*LME/MB*)** |
| **ADF(1)** | -2.39 | -10.10 | -1.78 | -12.19 | -8.05 |
| **ADF(3)** | -2.70 | -7.79 | -2.10 | -7.92 | -6.34 |
| The table gives the ADF(1) and ADF(3) unit root tests for the logarithms of monthly averages of the LME aluminium spot price, the Metal Bulletin (MB) price, their first differences, and the logarithm of the ratio between the two prices. The sample is January 1979 – June 2000 | | | | | |

Table 3 gives ADF(1) and ADF(3) statistics for the monthly averages of the logarithms of the LME and *Metal Bulletin* prices (ln*LME* and ln*MB* respectively) and their first differences over the sample of monthly data January 1979 – June 2000.[8] These tests establish that both prices are I(1) and therefore non-stationary. Table 3 also reports (column 5) the ADF test on the difference between the two prices, which is confirmed as being I(0).[9] As is visually apparent from Figure 1, the two price series are cointegrated with unit cointegrating vector.

## 3. Aluminium Price Volatility

The volatility of aluminium prices is of interest to us in that it provides some informal information as to whether exchange trading may have been associated with an increase or a reduction in the amount of noise in the price process. In Figuerola-Ferretti and Gilbert (2002), looking across the range of non-ferrous metals, we argued that exchange trading resulted in at most a small increase in price variability – see also Slade (1991).[10] However, the comparison in those papers was of exchange prices with producer list prices which may not have been representative of actual transactions prices. The data we have in this paper allows us to re-examine that conclusion for

[8] The ADF(3) statistics are included as a robustness check – there is no indication that anything more than a single lag is required to obtain serially independent errors.

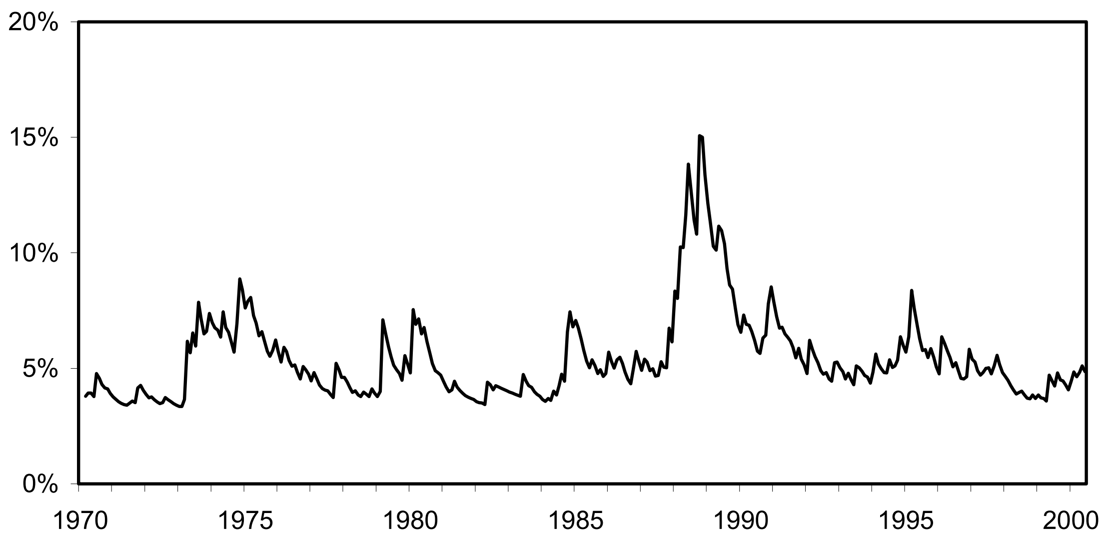
[9] We have checked the robustness of these results by performing a Phillips-Peron and Johansen cointegration test. Results are available under request. They support the hypothesis that both series are I(1) and cointegrated.

[10] One response is to suggest that the variability of prices is not *per se* important. Even if prices are more variable under the current exchange-pricing regime than they were under producer pricing, firms can now use liquid futures markets to offset this risk. In practice, however, hedging is costly, both because of commissions, roll costs, etc., but also because it requires management time, and, if not properly supervised, gives rise to new risks (Jorion, 1997, ch.2). If this is conceded, variability is indeed important.

aluminium using data which should provide a better measure of transactions prices.

It is arguable that there may be the possibility of "selection bias" which might vitiate this comparison.[11] The hypothesis to be tested is that LME trading was associated with an increase in the volatility of transaction prices for aluminium. Exchange trading requires a large pool of buyers and sellers and a considerable price variability in the commodity so that agents are prepared to use the market for hedging. This implies that aluminium LME trading would not have started unless prices had become more volatile. In other words, the direction of causality might be from price volatility to pricing regime. Reverse causality has been often claimed  see, for example, Telser (1981). According to this view, increased variability may have resulted in increased demand for insurance which in turn made hedging on commodity exchanges more attractive.[12] Whatever the merits of this view, it is not directly relevant to our concerns, since we are only attempting to discover whether aluminium price volatility increased over the time of the move to exchange trading, and not to establish a causal link.

**Figure 2: GARCH Estimates of *Metal Bulletin* Aluminium Price Volatility, 1970-2000**



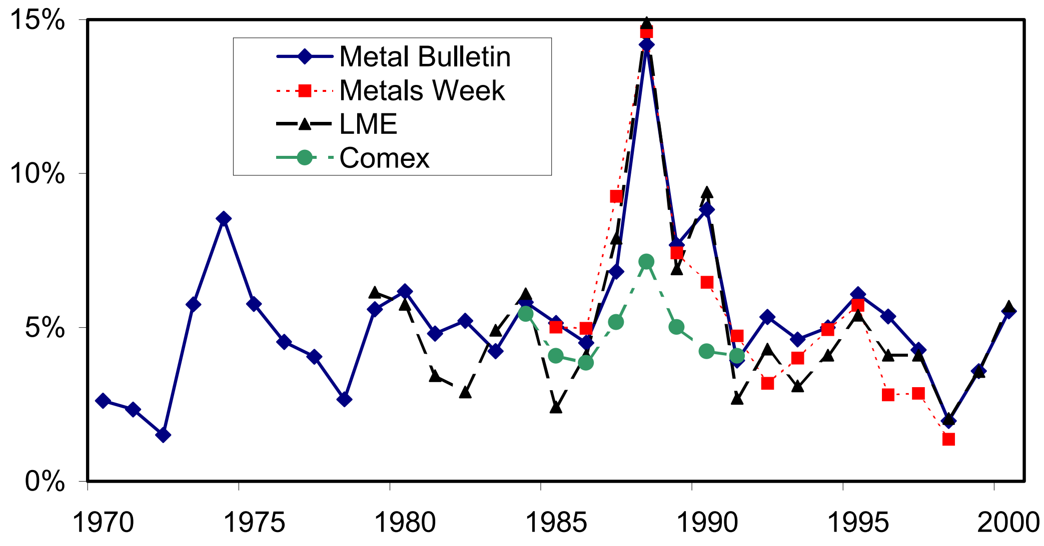

[11] We are grateful to Gordon Gemmill for this point.

### 3.1 Volatility Measures

We measure price volatility as the standard deviation of price returns over each month for which we have data for each of the Comex, LME, *Metal Bulletin* and *Metals Week* prices over the periods for which we have data. Since *Metal Bulletin* prices are not available on a daily basis, we used their estimated conditional variance from a GARCH(1,1) regression model.[13] The resulting annualized volatilities from the GARCH estimates are graphed in Figure 2. There is no discernible trend, but periods of high volatility are evident in the mid nineteen seventies, the early eighties, and (particularly) the late eighties.

In Figure 3, we plot the yearly standard deviations for the each of the COMEX, LME and *Metals Week* price series together the annual average of the GARCH estimate of the *Metal Bulletin* price volatility.[14] The four volatility estimates tell broadly the same story over the periods in which they are jointly available. The LME and *Metal Bulletin* volatilities appear to converge over the most recent years, while the *Metals Week* volatility diverges more. This may reflect the European focus of the

**Figure 3: Annual Aluminium Volatilities**

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[12] Slade (1991) tested for this possibility using data for a whole group of non-ferrous metals and found no evidence reverse causality.

[13] Engle (1982). See Bollerslev, Engle and Nelson (1994) for a summary. Estimation results available on request.

[14] The 2000 volatilities are based on six months data.

former pair of prices relative to the US focus of the latter. The COMEX volatility is the most divergent, probably as the consequence of the low liquidity achieved by this contract.

Although the volatilities of the four price series broadly move together, the correlations are only modest.[15] These are given in Table 4 for the sample of monthly data, 1985-1990, for which we have all four quotations.[16] Inspection of the time plots shows the LME and *Metals Week* prices as exhibiting very high volatility over 1987-88, while Comex and *Metal Bulletin* volatility rises less markedly.

**Table 4: Volatility Correlations, 1985-90**

| | LME | COMEX | *MB* | *MW* |
|---|---|---|---|---|
| LME | 1.0000 | | | |
| COMEX | 0.7629 | 1.0000 | | |
| *MB* | 0.4756 | 0.4058 | 1.0000 | |
| *MW* | 0.9519 | 0.7445 | 0.4842 | 1.0000 |

The table gives the correlation between the monthly volatilities of the four aluminium prices over the sample January 1985 – December 1990 for which all four prices are available.

### 3.2 Volatility Tests

We can test whether exchange trading was associated with an increase in the variability of the MB aluminium prices by performing an *F* test on the variances before and after the start of exchange trading. To do this, we are obliged to focus on the MB price series which is the only series we have which extends back prior to the start of exchange trading of aluminium. We do this first by using yearly standard deviations calculated from monthly returns and then we repeat the procedure using the monthly GARCH(1,1) volatility estimates.[17]

We split the entire sample into three periods:

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[15] Most striking of all is that the two transaction price series show very low correlations. This due to the fact that MB series is available on a monthly basis whereas MW comes on a daily basis.

[16] Although COMEX traded through 1991, this was only expiring contracts, and the market was very thin.

[17] We convert returns calculated from monthly averages to annual returns using a conversion factor of 3. This is based on the volatility of monthly average returns which, on the assumption of temporal independence, is related to the volatility of daily returns by the proportionality factor $(4n/3)^{½}$ , where $n$ is the number of trading days in the month. On the

Sample 1: January 1970 – December 1978, prior to the start of exchange trading

Sample 2: January 1979 – December 1985, the intermediate period in which the producer price functioned along side the exchange price.

Sample 3: January 1986 – June 2000 the post–producer pricing period.

Table 5 gives the estimated price volatilities of the *Metal Bulletin* price in these three sub-periods. Using both our measures, volatility is higher in Sample 3 than Sample 1. This appears to confirm that aluminium price volatility did indeed increase after the introduction of exchange trading. However, inspection of Figure 1 suggests that the high volatilities in the post-exchange trading period, Sample 3, may be entirely due to the high price volatility experienced in the late nineteen eighties. To investigate this, we divide sample 3 into

Sample 3a: January 1986-December 1990, and

Sample 3b: January 1991- June 2000.

On this basis, the increased volatility evident in Sample 3 appears to be entirely due to the very high volatility experienced in the late nineteen eighties (Sample 3a). Volatility in he nineteen nineties (Sample 3b) is only slightly greater than that in the nineteen seventies (Sample 1).

**Table 5: *Metal Bulletin* Price Volatilities**

| | Sample 1 1970-78 | Sample 2 1979-85 | Sample 3 1986-2000 | Sample 3a 1986-1990 | Sample 3b 1991-2000 |
|---|---|---|---|---|---|
| Annual | 17.69% | 16.81% | 22.33% | 30.56% | 16.57% |
| Monthly (GARCH) | 17.86% | 16.93% | 22.21% | 28.59% | 17.96% |

The table gives the volatility of the *Metal Bulletin* aluminium price calculated in the first row as the annual standard deviation of monthly price changes and in the second column as the square root of the GARCH estimate of the conditional variance estimated over the entire sample of monthly data, January 1970 – June 2000.

We may formally test these impressions using the standard Fisher variance equality test – see Table 6. The first column tests equality of the volatilities in Sample 1 with those in Sample 3, while the second column performs the same test for Sample 1 against Sample 3b, with p-values given in brackets. In both cases the results are

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same assumption, the volatility of annual returns is related to the volatility of daily returns by the proportionality factor $m^{1/2}$ , where $m$ is the number of trading days in the year.

clear-cut, confirming the impression deriving from the statistics in Table 6. The hypothesis of equal variances in Samples 1 and 3 is decisively rejected against the alternative of a higher variance in sample 3, but the hypothesis that the variances were equal in Samples 1 and 3b is not rejected. The late nineteen eighties (Sample 3a) appears to be an anomalous high volatility period. If we are prepared to ignore this period, aluminium price volatility appears neither higher nor lower than prior to the introduction of futures trading.

| **Table 6: Variance Equality Tests** | | |
|---|---|---|
| | Sample 3 versus Sample 1 | Sample 3b versus Sample 1 |
| Annual | $F_{13,9}$ = 1.51 [21.2%] | $F_{8,9}$ = 1.30 [26.8%] |
| Monthly (GARCH) | $F_{173,105}$ = 1.59 [0.49%] | $F_{113,105}$ = 0.88 [75.30%] |
| The table gives the outcomes and, in parentheses, the associated tail probabilities for the variance equality tests for the estimated *Metal Bulletin* volatilities reported in Table 3. | | |

Application of the Fisher F test to the estimated GARCH variances is not strictly valid since these estimated variances will not in general be independent. However, an alternative GARCH-X procedure gives the same result. We extend the GARCH(1,1) model by including three intercept dummies relating to samples 2, 3a and 3b:

$$\sigma_t^2 = h(1 + \phi_2 D_2 + \phi_{3a} D_{3a} + \phi_{3b} D_{3b}) + \alpha\varepsilon_{t-1}^2 + \beta\sigma_{t-1}^2 \quad (1)$$

The three dummies allow the mean of the GARCH process to vary across the four samples. Estimation results are given in Table 7

| **Table 7: GARCH-X estimates** | | | | | | |
|---|---|---|---|---|---|---|
| | $h$ | $\alpha$ | $\beta$ | $\phi_2$ | $\phi_{3a}$ | $\phi_{3b}$ |
| Coefficient | 0.000228 | 0.1209 | 0.7864 | -0.01500 | 1.75439 | 0.00583 |
| *t* ratio | 2.5162 | 2.8243 | 10.8522 | -0.0509 | 1.2068 | 0.0155 |
| The table gives the estimates of the GARCH-X model (3.18) applied to the *Metal Bulletin* monthly prices over the sample January 1970 to June 2000. | | | | | | |

The results show that none of the dummies are individually significant in explaining the conditional variance of *Metal Bulletin* prices. A likelihood ratio test (distributed as $\chi^2_3$) on their joint significance gave a value of 3.94 (tail probability

26.8%), implying failure to reject the null hypotheses of $\phi_2= \phi_{3a}= \phi_{3b}=0$. This test outcome is therefore consistent with a constant volatility process throughout the entire sample. However, the power of any test on GARCH estimates based on a relatively small number of observations may be low, and for this reason we prefer to rely on the traditional Fisher test.

# 4. The Beveridge-Nelson Decomposition and its Application to Aluminium

In this and the next section we apply two different methods to assess the relative efficiency of LME and MB prices in reflecting market conditions. In section 4, we apply the Beveridge-Nelson (1981) decomposition to estimate the informativeness of the two price series, and in section 5 we use a latent variable technique due to Grilliches (1977) and Zellner (1977).

### 4.1 The Univariate Beveridge-Nelson Decomposition

BN proposed a particular decomposition methodology for a non-stationary 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 the cycle. If we have two alternative price measures, either relating to the same of different time periods, we can think of the series which exhibits the greater transitory variance as being less informative about the underlying trend, and in that sense, less efficient. We will therefore compare the transitory variances of the two LME and *Metal Bulletin* price series to determine their relative efficiency.

BN showed that, in the context of a forecasting model for the first differences of the series, the trend may be thought of as the long run forecast of the level of the series and the cycle as the gap between the present level of the series and its long term forecast. The former may be interpreted as reflecting market fundamentals and the later the market noise inherent in the price series. By generalizing the BN decomposition to the bivariate case, we may apply it to the LME and MB price series. We can compare the estimated transitory variances and determine the most efficient price series in the sense we have defined.

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

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

where $L$ is the lag operator, $a(L)$ is a lag polynomial or order $m$ and the $\varepsilon_t$ are IID by construction. We may invert equation (1) to obtain the infinite MA representation

$$\Delta y_t = \alpha(L)\varepsilon_t \tag{3}$$

where $\alpha_0$ = 1. This allows us to write $y_t$ as

$$y_t = y_0 + \varepsilon_t + (1+\alpha_1)\varepsilon_{t-1} + ... + (1+\alpha_1+...+\alpha_{t-1})\varepsilon_1 + \kappa = y_0 + \sum_{j=0}^{t-1}\left(\sum_{i=0}^{j}\alpha_i\right)\varepsilon_{t-j} + \kappa \tag{4}$$

where $\kappa$ depends on pre-initial condition disturbances $\varepsilon_t$, with $t < 1$.

The proposed decomposition is

$$y_t = \mu_t + e_t \qquad where \qquad \mu_t = \mu_{t-1} + \nu_t \qquad and \qquad E[\nu_t \mid y_{t-1},...,y_0] = 0 \tag{5}$$

This decomposition is achieved by setting

$$\nu_t = (1+\alpha_1+....+\alpha_r+...)\varepsilon_t = \alpha(1)\varepsilon_t \tag{6}$$

implying $$\mu_t = \mu_0 + \alpha(1)\sum_{j=0}^{t-1}\varepsilon_{t-j} \tag{7}$$

From equations (3) and (4), we may express the transient error $e_t$ as

$$\begin{aligned} e_t = y_t - \mu_t &= \sum_{j=0}^{t-1}\left[\left(\sum_{i=0}^{j}\alpha_i\right) - \alpha(1)\right]\varepsilon_{t-j} + (y_0 + \kappa - \mu_0) \\ &= -(\alpha_1+\alpha_2+...+\alpha_r+...)\varepsilon_t - (\alpha_2+...+\alpha_r+...)\varepsilon_{t-1} - ... - \alpha_t\varepsilon_1 + (y_0 - \mu_0) \\ &= \sum_{j=0}^{t-1}\gamma_j\varepsilon_{t-j} + (y_0 - \mu_0) \quad where \quad \gamma_j = -\sum_{i=1}^{\infty}\alpha_i \end{aligned} \tag{8}$$

The transitory variance $\sigma_e^2$ is therefore

$$\sigma_e^2 = \left(\sum_{j=0}^{t-1}\gamma_j^2\right)\sigma_\varepsilon^2 \tag{9}$$

where $\sigma_\varepsilon^2 = E[\varepsilon_t^2]$. The variance $\sigma_p^2$ of the change in the permanent component follows from equations (5) and (6) as

$$\sigma_p^2 = \alpha(1)^2\,\sigma_\varepsilon^2 \tag{10}$$

### 4.2 Bivariate Generalization

We generalise the BN methodology to the bivariate case by taking the variable $y_t$ in (2) as the vector comprising the LME and *Metals Bulletin* prices. Explicitly, we replace (2) by the bivariate VAR representation

$$\Delta \ln MB_t = a_{1,0} + \sum_{j=1}^{3} a_{1,j} \Delta \ln MB_{t-j} + b_{1,j} \sum_{j=1}^{3} \Delta \ln LME_{t-j} + \varepsilon_{1t} \tag{11}$$

$$\Delta \ln LME_t = a_{2,0} + \sum_{j=1}^{3} a_{2j} \Delta \ln MB_{t-j} + b_{2,j} \sum_{j=1}^{3} \Delta \ln LME_{t-j} + \varepsilon_{2t}$$

Inverting to obtain the MA representation, selecting the MB equation, we obtain the analogous to equation (3) in the univariate case

$$\ln MB_t = \ln MB_0 + \sum_{j=0}^{t-1} \left( \sum_{i=0}^{j} \alpha_{1,i} \right) \varepsilon_{t-j}^{MB} + \sum_{j=0}^{t-1} \left( \sum_{i=0}^{j} \beta_{1,i} \right) \varepsilon_{t-j}^{LME} + \kappa \tag{12}$$

where $\alpha_{1,0} = 1$ and $\beta_{1,0} = 0$. We look for the same decomposition as in equation (5). To this end, we set

$$v_t = \alpha_1 (1) \varepsilon_t^{MB} + \beta_1 (1) \varepsilon_t^{LME} \tag{13}$$

implying $$\mu_t = \mu_0 + \alpha_1 (1) \sum_{j=0}^{t-1} \varepsilon_{t-j}^{MB} + \beta_1 (1) \sum_{j=0}^{t-1} \varepsilon_{t-j}^{LME} \tag{14}$$

and $$\begin{gathered} e_{1,t} = MB_t - \mu_t = \sum_{j=1}^{t-1} \gamma_{1j} \varepsilon_j^{MB} + \sum_{j=1}^{t-1} \delta_{1j} \varepsilon_j^{LME} + (\ln MB_0 - \mu_0) \\ \textit{where} \quad \gamma_{1j} = -\sum_{i=1}^{t-j} \alpha_{1i} \quad \textit{and} \quad \delta_{1j} = -\sum_{i=1}^{t-j} \beta_{1i} \end{gathered} \tag{15}$$

Write $$Var \begin{pmatrix} \varepsilon_t^{MB} \\ \varepsilon_t^{LME} \end{pmatrix} = \begin{pmatrix} \sigma_{MB}^2 & \rho \sigma_{mb} \sigma_{lme} \\ \rho \sigma_{MB} \sigma_{LME} & \sigma_{LME}^2 \end{pmatrix} \tag{16}$$

In this case, the transitory variance $\sigma_{e1}^2$ is

$$\sigma_{e1}^2 = \left( \sum_{j=1}^{t-1} \gamma_{1j}^2 \right) \sigma_{MB}^2 + 2 \left( \sum_{j=1}^{t-1} \gamma_{1j} \delta_{1j} \right) \rho \sigma_{MB} \sigma_{LME} + \left( \sum_{j=1}^{t-1} \delta_{1,j}^2 \right) \sigma_{LME}^2 \tag{17}$$

Analogously the transitory variance $\sigma_{e2}^2$ for the LME equation will be given by

$$\sigma_{e2}^2 = \left( \sum_{j=1}^{t-1} \gamma_{1j}^2 \right) \sigma_{MB}^2 + 2 \left( \sum_{j=1}^{t-1} \gamma_{2j} \delta_{1j} \right) \rho \sigma_{MB} \sigma_{LME} + \left( \sum_{j=1}^{t-1} \delta_{2,j}^2 \right) \sigma_{LME}^2 \tag{18}$$

From equation (13) it follows that the two permanent components have variances $\sigma_{p1}^2$ and $\sigma_{p2}^2$ given by

$$\sigma_{p1}^2 = \alpha_1(1)^2 \sigma_{MB}^2 + 2\alpha_1(1)\beta_1(1)\rho\sigma_{MB}\sigma_{LME} + \beta_1(1)^2 \sigma_{LME}^2 \tag{19}$$

and $$\sigma_{p2}^2 = \alpha_2(1)^2 \sigma_{MB}^2 + 2\alpha_2(1)\beta_2(1)\rho\sigma_{MB}\sigma_{LME} + \beta_2(1)^2 \sigma_{LME}^2 \tag{20}$$

Under normality, the VAR equations (11) may be efficiently estimated by Ordinary Least Squares (OLS). We may calculate the transitory variance of each of the price series by simulation of the estimated equations (10). By giving one standard deviation shock to each of the AR representations we calculate the $\alpha$ and $\beta$ coefficients of the inverted MA model specified in (12). We can then use equations (17-20) to compute the required variances.

### 4.3 Generalization to the Cointegrated Case

The bivariate BN generalization does not take into account the fact that the two price series we are analyzing are cointegrated. Failure to take into account cointegration implies that the estimates of the permanent components of the two series may not be mutually compatible. This is because, under cointegration, the two series must revert towards each other. The problem does not arise in the original univariate BN case because the permanent component of a single series cannot be mean reverting.

This makes it natural to generalize the bivariate VAR framework (10) to a cointegrated VAR, ie a CVAR. Hence (11) becomes

$$\begin{aligned}\Delta \ln MB_t &= a_{1,0} + \sum_{j=1}^{3} a_{1,j}\Delta \ln MB_{t-j} + b_{1,j}\sum_{j=1}^{3}\Delta \ln LME_{t-j} + c_1\left(\ln MB_{t-4} - \ln LME_{t-4}\right) + \varepsilon_{1t} \\ \Delta \ln LME_t &= a_{2,0} + \sum_{j=1}^{3} a_{2j}\Delta \ln MB_{t-j} + b_{2,j}\sum_{j=1}^{3}\Delta \ln LME_{t-j} + c_2\left(\ln MB_{t-4} - \ln LME_{t-4}\right) + \varepsilon_{2t}\end{aligned} \tag{21}$$

where we have imposed a unit cointegrating vector. By the Granger Representation Theorem (Granger, 1986; Engle and Granger, 1987), cointegration of ln*LME* and ln*MB* implies that one or both of $c_1$ and $c_2$ must be non-zero – ie one or both of the two prices must revert towards the other.

It is convenient to reparameterize equations (18) in terms of the change in one of the prices and the level difference between the two prices. Without loss of generality, we choose the change in the log *Metal Bulletin* price $\Delta \ln MB_t$ and the log price differential $\ln PD_t = \ln MB_t - \ln LME_t$. Equations (21) become

$$\begin{aligned} \Delta \ln MB_t &= A_{1,0} + \sum_{j=1}^{3} A_{1,j} \Delta \ln MB_{t-j} + B_{1,j} \sum_{j=1}^{4} \ln PD_{t-j} + \varepsilon_{1t} \\ \ln PD_t &= A_{2,0} + \sum_{j=1}^{3} A_{2j} \Delta \ln MB_{t-j} + B_{2,j} \sum_{j=1}^{4} \ln PD_{t-j} + \varepsilon_{2t} \end{aligned} \tag{22}$$

Equations (22) are now in the form of a standard bivariate VAR with the exception that the distributed lag on the levels term ln*PD* is longer by one period than that on the difference term Δln*MB*. Note in particular that the disturbances $\varepsilon_1$ and $\varepsilon_2$ on equations (22) are identical to those on equations (11).

Inverting equations (22) and selecting the *MB* equation, as previously, we obtain the Moving Average Representation of (22) as

$$\ln MB_t = MB_0 + \sum_{j=0}^{t-1} \left( \sum_{i=0}^{j} \alpha^*_{1,i} \right) \varepsilon^{MB}_{t-j} + \sum_{j=0}^{t-1} \left( \sum_{i=0}^{j} \beta^*_{1,i} \right) \varepsilon^{PD}_{t-j} + \kappa^* \tag{23}$$

where the asterisks indicate that inversion is relative to the reparamaterized equations. The transitory variance of the *MB* price $\sigma_{e1}^2$ is then

$$\sigma^2_{e1} = \left( \sum_{j=1}^{t-1} \theta^2_{1j} \right) \sigma^2_{MB} + 2 \left( \sum_{j=1}^{t-1} \theta_{1j} \phi_{1j} \right) \rho \sigma_{MB} \sigma_{PD} + \left( \sum_{j=1}^{t-1} \phi^2_{1,j} \right) \sigma^2_{PD} \tag{24}$$

where $\theta_{1j} = -\sum_{i=1}^{t-j} \alpha^*_{1i}$ *and* $\phi_{1j} = -\sum_{i=1}^{t-j} \beta^*_{1i}$ . The variance $\sigma_{p1}^2$ of the permanent component is

$$\sigma^2_{p1} = \alpha^*_1 (1)^2 \sigma^2_{MB} + 2\alpha^*_1 (1) \beta^*_1 (1) \rho \sigma_{MB} \sigma_{LME} + \beta^*_1 (1)^2 \sigma^2_{LME} \tag{25}$$

The variances $\sigma_{e2}^2$ and $\sigma_{p2}^2$ of the LME price can be calculated in the same way by reparameterizing equations (19) in terms of $\Delta \ln LME_t$ and $\ln PD_t$.

**4.4 Results**

We perform the decomposition over four sub-sample specifications.:

1970-78: This is the pre-LME sample.

1979-85: This is the period prior to widespread industry acceptance of the LME as providing a reference price for aluminium. We use the cointegrated VAR version of the decomposition.

1986-90: By this time, the LME was accepted as providing a price reference but LME volatility was unusually high.

1991-2000: The is the modern period with normal volatility levels.

For the initial period, 1970-78, we can only estimate the univariate BN decomposition for the Metal Bulletin price using equation (9). For the remaining three periods, we use the cointegrated version of the decomposition following equation (21). We also estimate the cointegrated representation for the entire period 1979-2000 for which we have both LME and *MB* prices. Results are presented in Table 8

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| Table 8a: Beveridge Nelson Decomposition and the transitory variance (1970-1986) | | | | | |
|---|---|---|---|---|---|
| | ***1970-1978*** | ***1979-2000*** | | ***1979-1985*** | |
| | **MB** | **MB** | **LME** | **MB** | **LME** |
| $\sigma_e$ | 28.48% | 11.09% | 9,54% | 7.29% | 7.26% |
| $\sigma_p$ | 16.99% | 5.60% | 3,00% | 5.87% | 5.86% |
| **sd** | 4.85% | 5.70% | 4,94% | 4.41% | 4.29% |

The table gives the results of the Beveridge-Nelson (1981) decomposition of the variances of the monthly averages of the *Metal Bulletin* (MB) and LME price series. The variance $sd^2$ is decomposed into a trend component $\sigma_p^2$ and a transitory component $\sigma_e^2$.

| Table 8b: Beveridge Nelson decomposition and the transitory variance(1986-2000) | | | | | | | |
|---|---|---|---|---|---|---|---|
| ***1986-2000*** | | ***1986-1990*** | | ***1991-2000*** | | ***1994-2000*** | |
| **MB** | **LME** | **MB** | **LME** | **MB** | **LME** | **MB** | **LME** |
| 6.50% | 5.95% | 12.05% | 10.30% | 0.97% | 1.10% | 1.80% | 1.48% |
| 5.49% | 4.55% | 5.78% | 5.08% | 5.66% | 4.37% | 6.30% | 5.30% |
| 6.20% | 5.14% | 8.32% | 7.33% | 4.85% | 3.75% | 4.85% | 4.09% |

The table gives the results of the Beveridge-Nelson (1981) decomposition of the variances of the monthly averages of the *Metal Bulletin* (MB) and LME price series. The variance $sd^2$ is decomposed into a trend component $\sigma_p^2$ and a transitory component $\sigma_e^2$.

Tables 8a-8b show that, if we ignore the 1986-1990 period were metal markets were in tight supply, we can see that the transitory variances for both LME and MB price series tend to diminish over time. The transitory variance for both series reaches its minimum point in the 1991-2000 period. This implies that both series become increasingly more efficient in reflecting the underlying trend.

Comparison of LME and MB transitory standard deviations $\sigma_e$ across samples shows that the LME has lower transitory variance in each sample. The only exception to this is the 1991-2000 period were we see MB transitory variance slightly lower value than the LME transitory variance. We believe that this small difference is due to sample bias, since when we select different sub-sample such as the 1994-2000 period we find the LME transitory variance is lower than the MB transitory variance ( 1.48% and 1.8% respectively). This suggests that our results are not robust to the sample specification and therefore we should consider a variable parameterisation formulation of the latent variable hypothesis.

## 5. Moving Window Latent Variable Analysis

In this section we attempt a formal analysis of the variability of the latent aluminium transactions price based on the recorded LME and *Metal Bulletin* prices. As was shown in the previous section, it is unrealistic to suppose that the relationship between the LME and *Metal Bulletin* prices has been constant over time - the reality is that the LME price has become increasingly important over time. This argues in favour of adoption of a moving window analysis.

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 - see also Griliches (1977) and Zellner (1977). An advantage of this approach is that it leaves open whether the exchange of the trade journal price is the more accurate measure of actual transactions prices, and, in our implementation, allows the relative importance of these two measures to evolve over time.

Write the unobserved latent price as $Z_t$. We model the changes in the two observed prices as third order autoregressions, augmented by error correction terms.[19]

[19] By the Granger Representation Theorem (Granger, 1986; Engle and Granger, 1987), cointegration of ln*LME* and ln*MB* implies that one or both of these equations must contain an error correction term.

$$\Delta \ln LME_t = \alpha_0 + \Delta Z_t + \sum_{j=1}^{3} \alpha_{1j} \Delta \ln LME_{t-j} + \alpha_2 \ln\left(\frac{LME}{MB}\right)_{t-1} + \varepsilon_{1t} \qquad (26)$$

$$\Delta \ln MB_t = \beta_0 + \Delta Z_t + \sum_{j=1}^{3} \beta_{1j} \Delta \ln MB_{t-j} + \beta_2 \ln\left(\frac{LME}{MB}\right)_{t-1} + \varepsilon_{2t} \qquad (27)$$

We may write this system as

$$\Delta \ln LME_t = \alpha_0 + \sum_{j=1}^{3} \alpha_{1j} \Delta \ln LME_{t-j} + \alpha_2 \ln\left(\frac{LME}{MB}\right)_{t-1} + e_{1t} \qquad (28)$$

$$\Delta \ln MB_t = \beta_0 + \sum_{j=1}^{3} \beta_{1j} \Delta \ln MB_{t-j} + \beta_2 \ln\left(\frac{LME}{MB}\right)_{t-1} + e_{2t} \qquad (29)$$

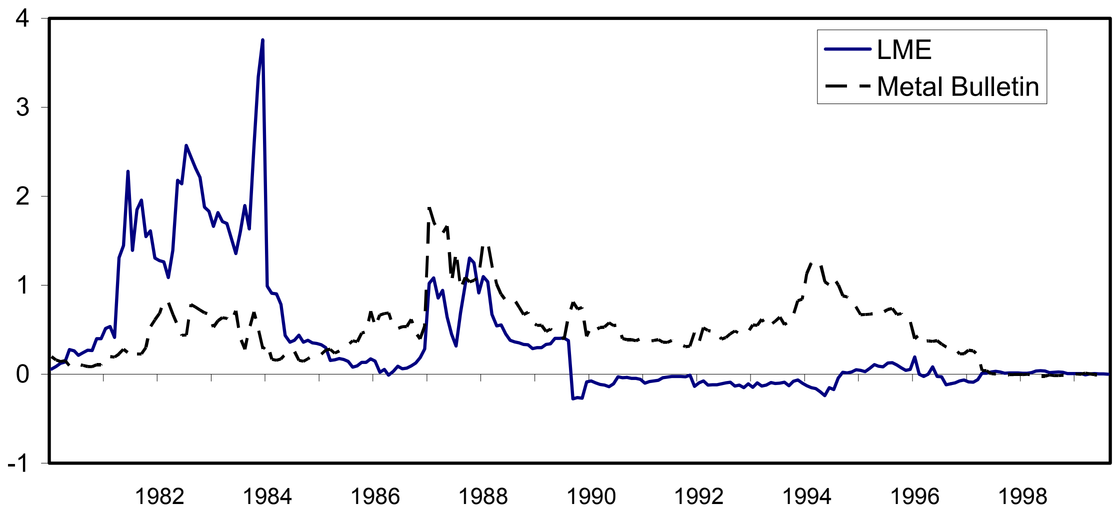
where $e_{jt} = \Delta Z_t + \varepsilon_{jt}$ ($j$=1,2). Assuming the two measurement errors are uncorrelated, ie $E(\varepsilon_t \varepsilon_t') = \begin{pmatrix} \omega_1^2 & 0 \\ 0 & \omega_2^2 \end{pmatrix}$, and $E(Z_t^2) = \zeta^2$ with $E(Z_s Z_t) = 0$ for $s \neq t$, it follows that

$$E(e_t e_t') = \begin{pmatrix} \sigma_1^2 & \rho\sigma_1\sigma_2 \\ \rho\sigma_1\sigma_2 & \sigma_2^2 \end{pmatrix} = \begin{pmatrix} \omega_1^2 + \zeta^2 & \zeta^2 \\ \zeta^2 & \omega_2^2 + \zeta^2 \end{pmatrix}. \qquad (30)$$

Define the noise-to-signal ratios $\lambda_j = \frac{\omega_j^2}{\zeta_j^2} = \frac{\sigma_j^2 - \rho\sigma_1\sigma_2}{\rho\sigma_1\sigma_2}$ ($j$=1,2). Using the entire sample April 1979 to June 2000 we estimate $\lambda_1 = 0.481$ and $\lambda_2 = 0.485$, so the two prices appear to be equally affected by measurement error.[20]

[20] Observations for January – March 1979 are lost through lag creation. Full estimates are available from the first author on request.

**Figure 4**
**Rolling Regression Estimates of the Noise-Signal Ratio**




This model is exactly identified – there is the same number of parameters in the model as moments defined by the data implying that there are no testable overidentifying restrictions. We estimate the system using a rolling twenty-four month window. The estimated values of $\lambda_1$ and $\lambda_2$ are plotted in Figure 4. The plot shows that, if we accept this model as a valid representation of the relationship between the two prices,

- The LME price was very noisy over the period 1980-84 while producers continued to set list prices.
- In the period 1985-90, the LME price became less noisy, but both this and the *Metal Bulletin* price continued to exhibit high noise-to-signal ratios. This was a period in which the prices of all non-ferrous metals were very volatile.
- During the period 1990-97, both price series appear to have become more accurate, with the LME price series, in particular, showing little evidence of measurement error.
- Finally, in the most recent period (1997-2000), the *Metal Bulletin* price series also shows little evidence of measurement error, consistent with the general move to pricing metals deliveries against the LME price.

# 6. Conclusions

In this paper we have used the *Metal Bulletin* transaction price series for aluminium to examine whether the move from producer list pricing to exchange pricing was associated with an increase in price variability (volatility). A formal statistical test failed to reject the hypothesis that the volatility of aluminium transactions prices differed in the nineteen nineties from its level prior to the introduction of futures trading in the nineteen seventies. This implies that the move from producer to exchange prices was not associated with a change in the volatility process of aluminium transaction prices.

We also view LME price and prices reported in the trade press as measures of the same underlying value. This view is justified by the cointegration of the two price series and their high correlation.

a) We extended the Beveridge-Nelson decomposition to the cointegrated bivariate case and used this decomposition to show that the LME price is and has always been associated with a lower transitory variance than the *Metal Bulletin* price. This allows us to interpret the LME price as a more reliable guide to actual transactions prices.
b) In a moving window framework in which both prices may be regarded as measuring an unobserved latent transactions price but with error, the LME price becomes an increasingly more accurate measure of that underlying value.

The Beveridge-Nelson decomposition suggests that the LME price is more informative than the *Metal Bulletin* prices, and was so right from the start of exchange trading of aluminium. This result is line with the view that exchange trading improves price discovery.

The moving window latent variable analysis suggests a more complicated story. In the first half of the nineteen eighties, prior to the general acceptance of the LME as a pricing basis, the LME price was indeed a more accurate measure of the latent transactions prices than was the *Metal Bulletin* price. Subsequently, the *Metal Bulletin* price became more accurate while, the increased price volatility which affected all non-ferrous metals markets in the late nineteen eighties resulted in LME prices exhibiting rather more noise than the prices reported in the *Metal Bulletin*.

Subsequently, the two prices appear to have converged and therefore contain the same information content.

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