Testing for Bubbles in LME Non-Ferrous Metals Prices

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This paper applies the mildly explosive/multiple bubbles testing methodology developed by Phillips, Shi and Yu (2013, Cowles Foundation Discussion Paper Nos. 1914 and 1915) to examine the recent time series behaviour of London Metal Exchange (LME) non-ferrous metals futures prices. The methodology allows for consistent date-stamping of the origination and collapse of bubbles. We find bubbles in all non-ferrous metals apart from aluminium. The bubble dates differ across the other five metals but fall into four distinct groups that can be related to the behaviour of fundamentals that economic theory suggests underlie them. This offers support for the efficacy of the PSY test if metals prices are primarily driven by fundamentals.

Keywords: mildly explosive process; recursive regression; sup ADF test; economic bubbles; non-ferrous metals

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1. INTRODUCTION

This paper tests for types of non-stationary behaviour in London Metal Exchange (LME) metals futures prices with a view towards characterizing their time series properties. Through a baseline test of a null of unit-root non-stationarity, it relates to a subject area of a significant part of John Nankervis’s work. Perhaps John’s best known contribution in this area was through the *Econometrica* paper by DeJong, Nankervis, Savin and Whiteman (1992), alongside a significant number of other articles in principal econometrics and statistics journals. Here, our concern is directed towards mildly explosive alternatives, which constitutes a similar approach but with interest centred on an autoregressive parameter possibly taking a value to the right of unity.

As such, the paper has connections to work being undertaken by John’s former colleagues in the Finance Group at the Essex Business School and, in particular, to the paper by Coakley, Kellard and Tsvetanov (2013) which uses the same test as here in an investigation of WTI crude oil prices. This work uses the multiple bubbles dating technology recently proposed by Phillips, Shi and Yu (2013a, 2013b, PSY). In a companion paper, Figuerola-Ferretti, Gilbert and McCrorie (2013b, FGM) have used a version of the test proposed by Phillips, Wu and Yu (2011, PWY) and Phillips and Yu (2011, PY) to examine the behaviour of many key commodity futures price series, focusing on bubble-like behaviour in terms of rapid run-ups in prices followed by abrupt collapse. Their main conclusion was that some commodity futures markets are prone to bubble-like phenomena and that the majority of the bubbles detected occurred in the twelve months August 2007 to July 2008, around the time of the 2007-08 financial crisis. Their secondary conclusion, looking more closely at the identified bubble periods, was that the time series behaviour was commodity-specific, with bubble incidence able to be related to known supply and demand features of past fundamentals behaviour that differed across the commodities. The test results therefore provided informal corroboratory evidence on the side of fundamentals-driven commodities rather than on the side of the rival financialization\(^1\) explanation based on speculation on the so-called commodities asset class which entails different commodities moving together.

The purpose of this paper is to use the PSY bubbles detection algorithm to examine LME non-ferrous metals futures prices. In FGM, we were primarily
interested in the existence (or not) of one bubble in the energy and agricultural sectors around the 2007-08 financial crisis, we used the PWY-PY testing methodology. The finance literature sees an asset class as a group of assets sharing a unique and common risk factor. On this criterion, non-ferrous metals form a distinct commodity class in that they are governed by largely common demand factors (Scherer and He, 2008).

In earlier work using the PWY-PY methodology, Gilbert (2010) reported evidence of multiple bubbles in a number of non-ferrous metals markets. This points towards use of a technology robust to possible multiple bubbles. Homm and Breitung (2012) have shown that the PWY-PY test has some efficacy against a multiple bubbles alternative, PSY have shown their test supersedes the PWY-PY test in such an environment.

Our aims are twofold. Firstly, we implement the PSY test using LME metals futures prices, with a view towards adding to the evidence base that describes the time series properties of such series. Secondly, we use the results obtained to assess the efficacy of the PSY test as a bubble detection algorithm against known fundamental behaviour, recognizing the competing explanations of commodity futures price behaviour discussed above. Results that identify multiple bubbles in metals prices at points in the sample period specific to each metal, that can be related to metal-specific fundamentals behaviour, offer corroboratory evidence to support the efficacy of the PSY test in a world where commodities are driven at least in part by fundamentals.

Section 2 introduces the PWY-PY-PSY bubbles-testing methodology. Section 3 discusses the LME non-ferrous metals and describes their raw time series properties over the chosen sample period. The PSY multiple bubbles test is applied to each metal price series in Section 4. Some interpretation of our results is offered in Section 5. Section 6 concludes.

2. THE PSY BUBBLE-TESTING METHODOLOGY

Phillips, Wu and Yu (2011, PWY), Phillips and Yu (2011, PY) and Phillips, Shi and Yu (2012, 2013a, 2013b, PSY) have recently developed a statistical testing methodology based upon a reduced-form autoregressive (AR) model and an invariance principle for mildly explosive processes by Phillips and Magdalinos (2007, PM). The test looks for evidence of temporary regime shifts, from unit-root non-
stationarity under the null, towards possibly multiple periods of mild explosivity followed by collapse under the alternative. PSY (2013a, pp. 6-7) explain how the test procedure can be interpreted as a test for economic bubbles in the context of a standard asset pricing equation.3

As used in this literature, the term “bubble” implies a mildly explosive departure from a unit root data generating process (DGP). It does not entail any causal explanation for this departure. If the unit root DGP is the consequence of an asset pricing process in which the price is linearly related to an underlying market fundamental (the dividend process in the case of an equity index) then, in the absence a structural change in the fundamental process, the period of explosive prices must have a non-fundamental explanation. By contrast, in the economics literature bubbles are often seen as caused by excess speculative activity that take prices away from fundamental values. In this paper, we use the term “bubble” in the narrower sense defined at the start of this paragraph. There is a strong theoretical basis for supposing that the prices of non-ferrous metals are related in a nonlinear manner to market fundamentals. In section 5 we consider to what extent bubbles in their prices may be explained by response nonlinearities.

The PSY test is prototypically based on a null specification involving a random walk process and a local-to-zero intercept term:4

\[ x_t = d T^{-\eta} + \theta x_{t-1} + \varepsilon_t, \quad \varepsilon_t \sim \text{i.i.d.}(0, \sigma^2), \quad \theta = 1, \quad t = 1, \ldots, T; \quad (1) \]

where \( x_0 = O_p(1) \), \( d \) and \( \sigma \ (0 < \sigma < \infty) \) are unknown parameters and \( \eta \) is a localizing coefficient that controls the magnitude of the drift as \( T \to \infty \) which, in principle, could be estimated on the basis of data (see PSY (2013b)). When \( \eta > \frac{1}{2} \), the drift is small compared with the random walk component and, under the given conditions, the partial sums of \( \varepsilon_t \) satisfy the functional central limit theorem

\[ T^{-1/2} \sum_{t=1}^{T} \varepsilon_t \Rightarrow B(\cdot) = \sigma W(\cdot), \quad (2) \]

where \( W \) is standard Brownian motion. The prototypical null set-up can be generalized in various directions, e.g. allowing for martingale difference sequence errors or some weak dependence.

PWY (2011) and PY (2011) consider the following single-bubble data generation process (DGP) under the alternative hypothesis5
\[ x_t = x_{t-1} 1(t < \tau_e \text{ or } t > \tau_f ) + \delta_T x_{t-1} 1(\tau_e \leq t \leq \tau_f ) + \left( \sum_{k=1}^{T} e_k + X^*_{\tau_f} \right) 1\{t > \tau_f \} + \varepsilon_t 1\{t \leq \tau_f \} \quad (t = 1, \ldots, T) \]  

(3)

\[ \delta_T = 1 + c/T^\alpha, \quad c > 0, \quad \alpha \in (0, 1). \]  

(4)

Under the conditions on \( c \) and \( \alpha \), the autoregressive parameter \( \delta_T \) is greater than unity and the model has what Phillips and Magdalinos (2007, PM) called a mildly-integrated root (on the explosive side of unity). The system (3) and (4) exhibits unit root (trend) behaviour until \( \tau_e \), mildly explosive (bubble) behaviour between \( \tau_e \) and \( \tau_f \), and then, upon re-initialization at \( X^*_{\tau_f} \) following bubble collapse, resumes its trend behaviour at \( \tau_f \). PSY (2013) considered alternative hypotheses based on (natural) extensions of (3) and (4) to include two or more bubble periods.

The regression model will, as in conventional (left-sided) augmented Dickey-Fuller (ADF) tests, usually involve transient dynamics: the PSY approach is recursive\(^6\) and implements an ADF-type regression in a rolling window. Starting from a fraction \( r_1 \) and ending at a fraction \( r_2 \) of the total sample, with the window size therefore being the fraction \( r_w = r_2 - r_1 \), we fit the regression model

\[ \Delta x_t = \alpha_{r_1, r_2} + \beta_{r_1, r_2} y_{t-1} + \sum_{i=1}^{k} \psi_{r_1, r_2} \Delta y_{t-i} + \varepsilon_t, \]  

where \( k \) is the lag order and \( \varepsilon_t \sim \text{i.i.d.} (0, \sigma^2_{\varepsilon_{r_1, r_2}}) \). The number of observations in the regression is \( T_w = [T r_w] \), where \([\cdot]\) denotes the integer part of its argument. The ADF-statistic (t-ratio) based on this regression is denoted by \( ADF^r_{i} \).

The earlier PWY (2011) and PY (2011) methodology involved forming a recursive sequence of right-tailed ADF-type tests based on a forward-expanding sample and used the supremum of these as the basis of a test. Homm and Breitung (2012) showed the test had some efficacy as a bubble detection algorithm for one and two bubble alternatives. However, PSY (2013) show that when the sample period contains two bubbles, although the PWY procedure consistently estimates the start date of the first bubble in any sample, it can fail to identify or consistently date-stamp the second bubble. This motivated PSY to formulate a test that covers more subsamples of the data and has more flexibility to choose a subsample that contains a bubble episode. Specifically, they formulate a \textit{backward sup ADF} test, where the
endpoint of the sample is fixed at a fraction \( r_2 \) of the total sample and the window size is allowed to expand from an initial fraction \( r_0 \) of the total sample to \( r_2 \). The backward sup ADF statistic is defined as

\[
BSADF_{r_2} (r_0) = \sup_{\eta \in (0, r_2 - r_0)} \{ADF_{r_2}^\eta \}.
\]

A test is then based on the generalized sup ADF (GSADF) statistic which is constructed through repeated implementation of the BSADF test procedure for each \( r_2 \in [r_0, 1] \). The GSADF test statistic is defined as the supremum of the BSADF test statistics:

\[
GSADF (r_0) = \sup_{r_2 \in (r_0, 1]} \{BSADF_{r_2} (r_0)\},
\]

PSY (2013a, Theorem 1, p. 11) gives the limiting distribution of (7) under (1) with asymptotically negligible drift \( (\eta > \frac{1}{2}) \). Critical values obtained by numerical simulation are set against an explosive alternative (for specific values of \( r_0 \)). Subject to a minimum bubble-duration condition (see below) date-stamping is achieved through the BSADF statistic: the origination and termination points of a bubble, \( r_e \) and \( r_f \), are estimated by

\[
\hat{r}_e = \inf_{r_2 \in (r_0, 1]} \{r_2 : BSADF_{r_2} (r_0) > scv_{r_2}^{\beta_r} \},
\]

\[
\hat{r}_f = \inf_{r_2 \in [\hat{r}_e + \log(T)/T, 1]} \{r_2 : BSADF_{r_2} (r_0) > scv_{r_2}^{\beta_r} \},
\]

where \( scv_{r_2}^{\beta_r} \) is the \( 100(1 - \beta_r) \) right-sided critical value of the BSADF statistic based on \([T r_2]\) observations.\(^7\)

It can happen that the BSADF statistic exceeds its critical value for a single observation or a small number of observations. The PY, PWY and PSY procedures all require that for an excess to count as signifying a bubble, it should last over a number of observations. On this basis, they define a “minimum bubble condition” of \([\log(T)]\) observations. On this basis, the termination of a bubble is the first observation after \([T \hat{r}_e] + \big[ \log(T) \big]\) for which the BSADF statistic falls below its critical value.

Conditional on a first bubble having been found and terminated at \( \hat{r}_f \), the procedure is then repeated upon re-initialization at this point in search of a second and possibly additional bubbles. PSY (2013) show that subject to rate conditions both this procedure and sequential application of the PWY procedure with re-initialization can provide consistent estimates of the origination and termination dates of two bubbles.
Simulation evidence in PSY (2013b) based on finite samples suggests that, while both have efficacy, the PSY moving window procedure is more reliable than sequential application of the PWY procedure.

3. **NON-FERROUS METALS PRICES SINCE 2000**

Metals fall into four broad groups: precious metals, ferrous metals, non-ferrous metals and minor metals. In this paper we look at the six major non-ferrous metals (aluminium, copper, nickel, lead, zinc and tin), all of which are traded on the London Metal Exchange (LME). We do not consider ferrous metals (iron ore and steel). Until around 2010, iron ore and steel prices were based on annual contracts negotiated between producers and consumers. The LME steel billet contract was only launched in 2008 and trades very low volumes while futures trading in iron ore began only in 2010. Precious metals, which are closely linked to monetary sector assets, form a separate asset class which we plan to address this class in a future paper. Prices for minor metals are quoted by a number of price reporting services but these markets are thin and prices change at infrequent intervals when supply is abundant.

Aluminium and copper are the two most important non-ferrous metals by value, both being widely used across the entire range of industry and in construction. Nickel is the third most important, its main use being as an input to the production of stainless and special steels. Lead and zinc, which have lower value-to-weight ratios, exhibit more complicated price behaviour arising out of the fact that they are joint products, typically cropping together in the same ores. Many lead-zinc mines therefore produce the two metals in proportions determined by geological factors with the consequence is that either one or the other metal is often in excess supply. Tin is the least important member of the group by value. The tin industry was in decline for much of the period since 2000 owing to the virtual disappearance of tin-plating, formally its principal end use. However, it has recently found new application in the microelectronics sector resulting in a resurgence of demand.

In this paper, we analyze weekly official prices for cash and three month LME aluminium, copper, lead, nickel, tin and zinc from January 2000 to December 2011 (612 observations). Table 1 reports descriptive statistics. The LME has been the dominant world non-ferrous futures market throughout the period under consideration. There are two other important copper futures markets: Comex
NYMEX, now part of the CME Group) in New York and the Shanghai Futures Exchange (SHFE). Comex trades aluminium and copper contracts. Comex copper prices are closely arbitrated with LME prices but there is very little volume in the aluminium contract. The SHFE trades aluminium, copper, lead and zinc for which it sets reference prices for intra-Chinese commerce. However the financial and administrative costs associated with importing metal into China, particularly those associated with access to hard currency, can result in substantial differences between world prices and internal Chinese prices. The sample is chosen to include the first twelve years of the new century which involve the end of the low demand period and the emergence of the metals price boom culminating in the 2008 financial crisis. It therefore covers a number of different demand and supply regimes.

The three month price series are charted for the period 2000-2012 in Figure 1. Non-ferrous metals prices had been subdued in the final decade of the twentieth century resulting in low sectoral profitability and low levels of investment in new mine capacity. Prices continued to be flat at low levels in the initial years of the new century. The closing months of 2003 saw the start of a period of renewed GDP growth in the OECD in conjunction with rapid industrial growth in Asia, particularly China, for which the mining industry was unprepared. On a measure of “world” industrial production formed by weighting the IMF “advanced economies” industrial production index with those for the four BRICs (Brazil, China, India and the Russian Federation), world industrial production grew at an average rate of 5.1% over the five year period 2003-07 as against 2.4% over the previous five years (1998-2002).

Metals intensities, which measure the metal content of industrial production, are generally seen as determined by engineering considerations and are little affected by prices over the short to medium term (Radetzki and Tilton, 1990). The result is that increases in industrial production typically translate into increases in the consumption of non-ferrous metals but this potential growth can be constrained by high prices if production is unable to respond. Consumption growth in aluminium increased from 3.0% per annum over 1998-2002 to 7.8% in 2003-07 while in copper and nickel, which both experienced slow growth in production, consumption grew at 3.4% and 2.7% respectively in 2003-07 against 2.9% and 3.6% in 1998-2002. This combination of inelastic supply with the rightward shift of the metals demand curve
pushed copper and nickel prices sharply higher from the end of 2003. Lead showed a similar pattern. Zinc prices were also stronger after 2003, but most notably so in those periods in which lead prices were relatively low. The prices of all six metals slumped to low levels in the post-Lehman period but then recovered in 2010-11 with copper, lead and tin reaching new peaks.

Differently from the other five non-ferrous metals, the aluminium market saw prices languishing at relatively low levels throughout the twelve year period. The demand side drivers for aluminium are the same as those for copper and nickel. Although aluminium price movements were correlated with those in copper and nickel, they were much more modest. The difference in these price patterns relates to the supply side. The decade saw a very substantial expansion of Chinese aluminium smelting capacity with the consequence that Chinese trade in aluminium remained in approximate balance over our sample, recording an average deficit of 0.9% over 1998-2002, an average surplus of 1.7% over 2003-07 and an almost exact balance over 2008-11. This contrasts with the situation in copper, where Chinese consumption of refined copper exceeded production by an average of 3.9% of world refined consumption over 1998-2002 rising to 6.6% in 2003-07 and 15.0% in 2008-11. Similarly in nickel, Chinese nickel consumption exceed domestic production by 1.0% in 1998-2002, 6.8% for 2003-07 and 14.8% for 2008-11.

The stock-consumption ratio is one of the two most widely used measures of market fundamental in the metals industry (the other is metal balance, the excess of consumption over production). Over the five years 1998-2002, LME warehouse stocks of copper averaged 2.1 weeks 2002 consumption but this fell to average of 0.8 weeks 2002 consumption over the five years 2003-07. Conversely, in aluminium the stock-consumption ratio rose from 1.5 weeks to 1.8 weeks 2002 consumption over the same period.

Cash and three month prices are linked by a cost of carry relationship and so typically move closely together. According to this relationship, the three month price will exceed the cash price by the warehousing and interest cost less the convenience yield which reflects the option value of immediate access to the metal. When stocks are plentiful, convenience yield is near zero and the market is said to be at “full carry”. Since warehousing costs and interest rates vary only slowly, the two prices generally move in step. However, when stocks are limited, convenience yield can be both high and variable resulting in a backwardation market in which the cash price is
at a premium to the three month price. A temporary shortage will have greater impact on the cash price than on the three month price making it more likely that an apparently explosive sequence may emerge in cash prices. This is reflected in the higher cash price volatilities – see Table 1, column 1. We therefore consider both sets of prices.

In order to maintain compatibility with the rational bubbles model, we follow PY, PWY and PSY in analyzing price changes and not price returns. In analyzing these data, we modify the basic PSY procedure in two respects. The first relates to possible serial dependence in the disturbances. A standard result in finance theory is that the price of a futures contract should follow a martingale process – see Samuelson (1965). Differently from other commodity futures markets which have single contract maturing each month (or each two or three months in agricultural markets), a different LME contract “becomes prompt” each trading day. Prices on successive days therefore relate to different contracts with the consequence that there is no longer any reason to expect a strict martingale property to hold. The lag lengths, selected using the Akaike Information Criterion (AIC), adopted in the ADF tests reported in the final column of Table 1 indicate the presence of significant serial dependence in the price changes for all the metals with the exception of copper. The possibility of serially correlated disturbances is anticipated in the code developed by PSY. We exploit this feature.

The second complication, which is standard in financial markets, is excess kurtosis. The second and third columns of Table 1 show that while there is little evidence of asymmetry in the price change distributions, these do exhibit fat tails. Normality is decisively rejected in all twelve cases. The i.i.d. assumption in PSY (2013b) can be extended to more general martingale difference sequence errors. We interpret the PSY choice of normality as being made for convenience and compute revised critical values using a Student distribution.

4. TEST RESULTS

Table 2 presents a summary of the periods for which bubbles are identified using the PSY procedure modified to account for serial dependence and excess kurtosis. The explosive episodes differ across the six metals. Except in one case in which an explosive episode is identified (tin in 2007), these periods are common to both cash
and three month prices, albeit sometimes with slightly different start and end dates. Aluminium is the only metal for which there is no evidence of explosive behaviour. The dates of the periods of explosive growth differ across the other five metals but fall into four groups:

a) 2003-04: Copper, lead and tin.
b) 2005-06: Copper and zinc.
c) 2007: Lead, nickel, tin and zinc.
d) 2008: Tin.

With this background, we consider the results in greater detail. The GSADF statistic tests for the presence of at least one episode of explosive behaviour in the sample. Table 3 compares the twelve GSADF statistics (two contracts for each of the six metals) with the PSY 95% critical value (column 2) and the critical values modified for sample excess kurtosis and serial dependence (column 3). Critical values are obtained by Monte Carlo simulation using the program made available by Shu-Ping Shi. In each case, they are based on 5,000 replications over the subsample 

\[ T_0 : T \]

where \( T_0 = r_0 T \) and \( T \) is the full span. We follow PSY in setting \( r_0 = 0.02 \). In the base case, the critical values depend only on the sample size and so are the same in all twelve cases. The GSDAF statistics exceed the base critical values in all twelve cases although the excess is only small in the case of aluminium.

In the modified procedure (columns 3 and 4), which allows for serially correlated and kurtotic disturbances, the critical values are calibrated against the metal in question. Autoregressive parameters and kurtosis were obtained from an AR(3) estimated over the entire sample for each metal. The Monte Carlo replications, which use a slightly modified version of the Shi program, are based on an AR(3) data generating process with the same parameter values and a Student-distributed innovation process with degrees of freedom selected to match the sample kurtosis. The critical values therefore differ across the six metals. In practice, kurtosis is very similar for the cash and three month prices in all cases except that of nickel (see Table 1) and the same is true of the autoregressive parameters. This allows us to economise by using common critical values for the cash and three month prices for each of these five metals giving a total of seven critical values and critical value functions. Allowance for autoregression changes the resulting GSADF statistics, although not in
a consistent direction. The qualitative conclusions are unchanged from the base case with explosive behaviour signalled in all twelve cases.

These significant values for the GSADF statistic indicate the sample contains at least one explosive episode but are not informative about how many such episodes there were or the dates between which these occurred. It is also possible that an explosive episode identified by the GSADF statistic may fail to meet minimum bubble length criterion. Bubble identification results from comparison of the sequence of SADF statistics generated by sequential backward application of the ADF procedure on the sample \([rT_0: T]\) as \(r\) moves backwards through the range \([r_0 : 1]\). The SADF\((r)\) statistic is the supremum of the resulting ADF statistics over the range \([rT_0: rT]\). If the SADF\((r)\) statistic exceeds its critical value over the range of observations \([rT: rT + \log T]\), we have identified a series of explosive prices commencing at date \(rT\) which satisfy the minimum bubble length criterion and which may therefore potentially be explicable as a bubble. Periods of explosive prices terminate on the first subsequent date on which the SADF\((r)\) statistic falls beneath the relevant critical value.

The critical value sequences for the SADF statistics are calculated using a slightly modified version of Shi’s Backward SADF program with the same parameter settings as for the GSADF statistic. This generates a sequence of critical values over the range \([r_0: 1]\) against which the SADF\((r)\) sequence can be compared. These sequences are identical across metals in the base case but are metal-specific (and in the case of nickel, also contract specific) once we allow for serial dependence and kurtosis.

[Figures 2-7 near here]

Figures 2-7 plot the backward SADF sequence and the associated critical value sequences for the three month prices. These are computed allowing for serial dependence and excess kurtosis. They also show (feint) the time path of the three month price series. During periods of rapidly rising prices, the SADF statistic and the price tend to rise together but this comovement stops if the price falls back even if only temporarily. Charts for the cash prices are similar but are omitted to conserve space. These charts define the explosive periods previously listed in Table 1. Detailed results are as follows:

**Aluminium:** The SADF statistic exceeds the 95% critical value sequence on ten separate occasions with the supremum occurring in May 2006. However, these
excesses are either isolated incidents, as in 2004, or fail to overcome the minimum bubble length of seven successive weeks, as in 2006. The apparent excess of the SADF over the critical value in 2008-09 also fails the same length criterion but was, in any case, be associated with a downward price movement, incompatible with rational bubble theory.

**Copper:** Two bubble periods are identified. The first started in mid-January 2004 (end December for cash prices) and extend to mid-April 2004, a total of 14 weeks for the three month prices and 17 weeks for cash prices. The second period is longer and more pronounced. It started in November 2005 and extended through to June 2006 amounting to 28 weeks in the case of the three months price and 30 weeks in the cash case. Using the basic PSY procedure, the periods of explosive behaviour are even longer and extend through to November 2006, making almost a full year, for the three month case. There are also nine isolated excesses which fall at the minimum length fence.

**Nickel:** This is the most straightforward of the six cases. A single bubble is identified extending from mid-January 2007 to mid-June 2007 and amounting to 21 weeks for both cash and three months contracts. There are five isolated excesses of the SADF statistic over its critical value, all in the closing months of 2006 immediately prior to the 2007 explosive period.

**Lead:** Two bubbles are identified. The first, which mirrors the developments in the copper market, runs from mid-December 2003 to early March 2004 and amounts to 13 weeks for the three month price and 11 for the cash price. The second period extends from early June 2007 through to mid-November 2007, 25 weeks for the three months price and 24 for the cash price. There are also 16 isolated excesses or short periods of excess which fail to meet the minimum bubble length criterion.

**Tin:** This is the most complicated of the six cases considered. The procedure identifies three periods in which the three month price was explosive – mid-February to mid-April 2004 (ten weeks), early October to early December 2007 (nine weeks) and early February to early June 2008 (18 weeks). The first and third of these periods are also identified for the cash price. In the first case, the explosive period extends to May 2004 (14 weeks) while in the 2008 case it extends to early August (26 weeks). In 2007, the supposedly explosive period is less-well defined in the cash data and fails the seven week minimum length criterion. In the three month case there are also 21 periods in which the SADF statistic exceeds its critical value on isolated weeks or
when the periods of excess are insufficiently long to meet the minimum length criterion. Several of these periods occur in the second half of 2001 when the tin price was falling. The basic PSY procedure signals a downwardly explosive process over that period which would be incompatible with rational bubble theory.

**Zinc:** Two bubble periods are identified, the first from end-November 2005 to mid-June 2006 (29 weeks for both cash and three month prices) and the second from July 2006 (September 2006 in the case of cash prices) to early January 2007 (15 weeks for the cash price, 28 for the three months price). Figure 7 suggests that these two periods of explosive behaviour can probably be regarded as a single bubble episode but interrupted by a deceleration but not a collapse of prices in the late summer of 2006. There is a relatively small number (eight) of isolated excesses.

Two conclusions emerge from these results. First, in terms of the bubble identification methodology, the PSY procedure appears successful in discriminating between cases in which there are no periods of explosive behaviour (aluminium), a single period (nickel and perhaps zinc) and multiple periods (copper, lead and tin). A minor problem which arose in our aluminium example but which we have not attempted to rectify, is that the GSADF can reject the absence of explosive behaviour while the backward sequence SADF tests can nevertheless fail to identify an explosive period because of imposition of a the minimum length requirement. Secondly and substantively, the results confirm Gilbert’s (2010b) conclusion, derived from use of the PWY methodology, the LME copper market did exhibit two bubbles over the period we have examined. This tendency to multiple periods of explosivity generalizes here across other non-ferrous metals markets. The PSY procedure appears well-adapted for analysis of these markets.

**5. INTERPRETATION**

A number of commentators, both academic and in the wider world, have asserted that the initial decade of this century was characterized by bubbles. In particular, Caballero, Fahri and Gourinchas (2008a, 2008b) argued that the US subprime crisis in 2007 was just the most important of a sequence of bubble creations and collapses that started with the NASDAQ tech boom in 1997-2000 and which subsequently migrated from one sector of the economy to another. The crude oil price peaked in the summer of 2008 and many food prices also reached very high levels earlier in the same year.
PY, PWY and PSY are explicit that bubble identification provides their motivation in testing for periods of weakly explosive price behaviour.

Accounts of bubble creation typically link the emergence of bubbles with excessive or misguided speculation which divorces the market price form the underlying fundamental value. The Commodity Futures Trading Commission (CFTC), which is the U.S. futures market regulator, defines a speculative bubble as a rapid run-up in prices caused by excessive buying that is unrelated to any of the basic, underlying factors affecting the supply or demand for a commodity or other asset. Speculative bubbles are usually associated with a “bandwagon” effect in which speculators rush to buy the commodity (in the case of futures, “to take positions”) before the price trend ends, and an even greater rush to sell the commodity (unwind positions) when prices reverse.

The CFTC’s bubble characterization links bubbles explicitly with excess speculation. Financialization is the process through which large numbers of financial actors, specifically investment banks, hedge funds and index investors, have become involved in commodity futures markets. There is a widespread public perception that financialization may have contributed to the commodity price spikes in 2007-08. This perception has been stimulated by pronouncements by prominent politicians and leading market commentators. To give three examples, hedge fund manager Michael Masters (2008) argued in evidence before a US Senate subcommittee that index-based investment both raised the levels of commodity prices; a 2009 U.S. Senate Subcommittee report examined “excessive speculation” in the wheat market (United States Senate Permanent Subcommittee on Investigations, 2009); and French President Sarkozy asked in 2011, “Speculation, panic and lack of transparency have seen prices soaring. Is that the world we want?” The financialization literature is surveyed by Mayer (2010, 2011), Irwin and Sanders (2012b) and Gilbert and Pfuderer (2014).

Several authors have explored the routes by which speculation might result in price bubbles. Diba and Grossman (1988) set out the theory of rational bubbles in which the expected return on the bubble path is sufficient to compensate for the risk that the bubble may burst. Branch and Evans (2011) provide a comprehensive bibliography of this literature. Behavioural economists emphasize benchmarking and herding behaviour (Froot, Scharfstein and Stein, 1992). There is also a large literature, initiated by Smith, Suchanek and Williams (1988), based on laboratory
market experiments – see Porter and Smith (2003) for a survey. This literature confirms that, at least in a laboratory environment, speculation can result in bubbles.

It is also possible that an explosive price sequence might result either from explosive developments in the underlying fundamentals, as explored in PWY (2011), or from nonlinearities in the relationship between fundamentals and the commodity prices. The reverse inference, that a period of explosive price growth was a speculative bubble, is therefore not warranted.

The market fundamental in equities markets may be represented by the dividends on the basket of shares that make up the market index. Share prices should be linearly related to this fundamental – see Gordon (1959) and end note 3. PWY (2011) relate Nasdaq prices to the Nasdaq dividend yield and show there is evidence of explosivity in the former but not the latter. The implication is that the evidence of price explosion over the late 1990s cannot be explained by fundamental developments but points to a financial market epiphenomenon. From a formal point of view, convenience yield\(^{16}\) is the counterpart in the commodities markets to dividend yield in the equities market – see Pindyck (1993). However, while dividend yield may be taken as an exogenous process, convenience yield is jointly determined with the futures price structure with the consequence that an explosive price process might generate an explosive convenience yield.\(^{17}\) Commodity market analysts therefore typically refer to the stock-consumption ratio and the market balance as the market fundamentals.

The commodities literature emphasizes the non-negativity constraint on stocks as a cause of nonlinear responses. An increase in production or shortfall in production can be met by a run-down in stock levels so long as a positive inventory is held. However, once stock-out has been reached, this option is no longer available and a larger price response is required to achieve market balance – see Williams and Wright (1991) and Deaton and Laroque (1992).

A second and perhaps more important source of nonlinearity in metals market price behaviour arises out of a combination of geological and engineering constraints on increasing mine production beyond full capacity levels. For example, the initial post-mining stage of copper production involves beneficiation or concentration of the ore. The capacity of a mine’s concentrators impose a physical limit on the tonnage of ore that a mine can process. Installation of an additional concentrator is a major investment which will take two to three years before it can be brought into use. In
principle, a mine can increase copper output by processing higher grade ores but generally mines have little flexibility in choosing the grade or ore obtainable from any given rock face. Consequently, when a mining industry is working near full capacity, market balance can only be achieved by reduced consumption, increased production from scrap\(^{18}\) or a run-down in stocks. Demand elasticities are typically low so once stocks come close to exhaustion, market clearing prices can become very high.

The simple annual agriculturally-orientated models of Williams and Wright (1991) and Deaton and Laroque (1992) possess a single state variable, availability, equal to the current harvest plus the carryover from the previous year, which determines both the price and the end-year stock level. Availability is therefore the ideal measure of market fundamentals in these markets. In metals markets, where consumption is also highly inelastic, it may be preferable to regard metal requirements, equal to consumption less primary production and lagged carryover, as the measure of fundamentals. This combines the metals balance measure and stock-consumption measures used by metals industry practitioners.

Even if the absence of any element of explosivity in the metals market fundamental, it may still be the case that the response nonlinearity resulting from the stock-out and capacity constraints might generate periods of explosive prices. The price flexibility of a market measures the proportional change in price resulting from a proportional unit shock to demand or supply. Metals market price flexibilities (i.e. the reciprocals of net demand elasticities) will rise as full capacity and stock-out are approached. A sequence of shocks which tend to reduce stocks towards their minimum attainable levels or to increase production towards full capacity will have a progressively larger impact. The price rises which result in these circumstances may be explosive.\(^{19}\) In such cases, explosive price behaviour arises out of nonlinearities in the reaction of prices to market fundamentals and not in breaks or discontinuities in the fundamental process itself.

**[Table 4 around here]**

We investigate this possibility is an informal way by constructing an annual measure of market fundamentals, the consumption-supply ratio (CSR). This is defined as the ratio of consumption of the metal in question to production in the same year plus the stock level at the end of the previous year. The ratio satisfies \(CSR < 1\) by construction with a value close to unity indicating a tight market. This is a scale-free measure of the requirements state variable discussed above. Values of the CSR are
shown in Table 4. We relate this variable to periods of explosive price behaviour by defining a dummy variable, explode, which takes the value unity for any year-metal pair in which the PSY procedure registers an explosive price sequence for both the cash and three month prices and is zero otherwise. This gives us 72 observations (six metals for twelve years) of which twelve yield non-zero values for explode. A simple logit relationship relating CSR to the dummy variable explode and the lagged value of the bubble dummy gives a t-statistic of 1.30 on CSR (AIC = 63.53). The discussion earlier in this section emphasized nonlinearity in metals market responses. Replacing CSR in the logit by the transformed variable \( \frac{1}{1-CSR} \), which tends to infinity as CSR tends to unity, raises this t-statistic to 2.24 (AIC = 59.22). This informal test therefore confirms that periods of explosive behaviour are associated at least in part with market fundamentals as accentuated by response nonlinearities.

The fitted probabilities from the logit exceed 0.5 for copper in 2005 and for lead in 2004. In addition, they exceed 0.33 for copper in 2004, 2006 and 2007, for lead in 2005, 2007 and 2008 and for zinc in 2006 and 2007. The fitted equation correctly predicts the absence of bubbles in aluminium where the highest fitted bubble probability is 0.1.

The claim that periods of explosive behaviour in the non-ferrous metals markets were associated with market fundamentals does not rule out a role for speculation which may have exacerbated or otherwise amplified explosive movements arising out of market tightness. Furthermore, the fitted probabilities fail to account for the explosive episodes in nickel and tin. Nevertheless, these results are sufficient to reject the view that the bubbles in non-ferrous metals markets during the first decade of the century were purely financial epiphenomena.

6. CONCLUSION

The PSY test that we have applied to LME metals prices entails using a reduced form approach that first and foremost tries to capture the essential time series properties in the data. A significant part of John’s approach to Econometrics was in contributing methodologically to the problem of testing between stationary and non-stationary time series, in a way that helped us more powerfully characterize and distinguish between the essentially different properties of a time series in different regions of the
parameter space. In this context, our paper has used the PSY recursive test to look at the extent to which it identifies bubbles (short-term parameter non-constancy in the form of mild explosivity) that are thought to characterize certain economic time series. Economic theory suggests that relationship of non-ferrous metals prices to market fundamentals is likely to be strongly nonlinear and these nonlinearities may result in periods of explosive price behaviour.

Under the PSY procedure, we detected explosive episodes in the price histories five of the six non-ferrous metals prices, with start and end dates that were metal-specific. In four of the six cases, we found multiple bubbles. In a previous paper (Gilbert, 2010), the second author previously reported multiple bubbles for copper but this was based on repeated use of the PWY (2010) procedure which is not robust to the presence of multiple bubbles. The results in this paper confirm that the PSY procedure performs well and in a straightforward manner. At this point in time, the technology to undertake this when the series exhibit possible bubble-type behaviour is not available but the multivariate theory derived by Phillips and Magdalinos (2007) might offer a basis for such a formal test. We leave this to future work.

The search for periods of explosive prices is in large part motivated by the concern that markets are prone to speculative bubbles which divorce prices from market fundamentals. Univariate methods are necessarily uninformative about market fundamentals so the inferences about bubble causation cannot be direct. In the non-ferrous metals markets, in which short term price responses are highly inelastic, fundamentals can be characterized in terms of a market balance variable which, however, can only be constructed at a relatively slow data frequency. An informal exercise suggests that the explosive behaviour in three of the metals analyzed, copper, lead and zinc, were related to market fundamentals. The presence of a fundamental foundation for these three metals does not rule out the possibility that speculators may have exacerbated these movements but it does caution that we should not rush to characterize all explosive price periods in commodity markets as being speculative bubbles. Attempts to assimilate commodity market bubbles into an account of the build-up to the 2008 financial crisis appear misconceived since the sequence of non-ferrous metals bubbles commenced at the end of 2003 and were largely over by 2008.

The third author remembers an occasion when John was an integral member of the group that participated in a memorial conference for another New Zealander econometrician, Rex Bergstrom, who had also been in Essex for a major part of his
career. No-one at that conference could have imagined the same circumstances would arise from John’s unexpected death only seven years later. John will be remembered as someone who took an optimistic view over what Econometrics as a discipline could achieve, especially in terms of its statistical role in helping to explain the world around us. Yet his insights were derived, exactly in this context, through his sensitivity to its limitations. He was always willing to see the best in people and, in conversation, he was distinguished by the fairness and integrity with which he judged the contribution of others.

ACKNOWLEDGEMENTS

We have benefited from the GAUSS code for the PSY procedure made available on Shu-Ping Shi’s website (https://sites.google.com/site/shupingshi/home/). Figuerola-Ferretti thanks participants at the 21st Annual Symposium of the Society for Nonlinear Dynamics and Econometrics, University of Milano-Bicocca, March 2013, for comments on the paper in its preparatory stages, and the Spanish Ministry of Education and Science for support under grants SEJ2010-0047-001 and SEJ2011-0031-001.
NOTES

1. Financialization is the process through which large numbers of financial actors, specifically investment banks, hedge funds and index investors have entered commodity futures markets, notably since 2003. They amount to non-commercial actors who have begun treating commodities as a distinct asset class, alongside equities, bonds and real estate. Evidence that such actors have affected futures prices through speculation has been offered by Masters (2008, 2010), Soros (2008) and the U.S. Senate Permanent Subcommittee on Investigations (2009). Because they affect the market mainly through investment strategies that replicate the returns on commodity indices such as the S&P GSCI and DJ UBS indices, their impact is common across commodities. Evidence from the PSY test of commodity-specific time series behaviour can therefore be interpreted as evidence against financialization being the major driver of commodity futures prices.

2. The central difficulty in formulating tests for bubbles relates to the explosive nature of the non-stationarity underlying them. In the AR framework, the investigator has to confront the lack of any invariance principle under explosivity (Anderson, 1959), with any limiting distribution depending on the actual (unknown) distribution of the model’s disturbances. PM established an invariance principle for a class of processes they called mildly explosive processes, which PWY, PY and PSY have shown forms a sufficiently wide context to have empirical relevance. See Gürkaynak (2008) and PWY (2011) for up-to-date bibliographies on empirical tests for bubbles. Branch and Evans (2011) provide a reasonably complete bibliography for the rational bubble literature.

3. The standard asset pricing model is given by

$$ P_t = \sum_{i=0}^{\infty} \left( \frac{1}{1+r_f} \right)^i E_t(D_{t+i} + U_{t+i}) + B_t, $$

where $P_t$ is the (present-value) price of an asset, $D_t$ is the payoff received from the asset (which in the case of stocks is the dividend and for commodities like LME metals is the convenience yield), $U_t$ represents unobservable fundamentals, with $r_f$ the risk-free interest rate. $P_t - B_t$ is the market fundamental, where $B_t$ defines the bubble component which is assumed to satisfy

$$ E_t(B_{t+i}) = (1 + r_f) B_t. $$

When $B_t = 0$, the degree of non-stationarity in $P_t$ is controlled by the nature of the series $D_t$ and $U_t$: if $D_t$ is an I(1) process and $U_t$ is either an I(1) or I(0) process, then $P_t$ would be at most an I(1) process. Under (2), $P_t$ will be explosive in the presence of bubbles and, under the assumed conditions on $D_t$ and $U_t$, the observation of mildly explosive behaviour in $P_t$ (i.e. non-stationarity of an order greater than unit root non-stationarity) will offer, under (1), evidence of
bubble behaviour. As Evans (1991) showed, the econometric problem is complicated because it possible for bubbles to satisfy property (2) yet be periodically collapsing, making it difficult to distinguish between such series and stationary series. Such problems of power are ameliorated by PWY, PY and PSY through formulating the test on a recursive basis.

4. In unit root testing, whether on the left or right side of unity, the null specification is important because the distribution theory of test statistics can depend on it (see Hamilton, 1994, Ch.17, for a discussion in the context of conventional unit root tests directed towards stationarity). Here, the null specification is a hybrid model in which the relative size of the deterministic drift and random walk component depends on the localizing coefficient.

5. The process is formulated without an intercept since a non-zero intercept would produce a dominating deterministic component of an empirically unrealistic explosive form. See pp. 5-6 of Phillips and Yu (2009) for further discussion.

6. PY incorporate an additional tuning parameter in the statement of their minimum bubble-duration condition that they note can be chosen, in principle, on the basis of sampling frequency. They did not, however, use it in their subsequent empirical work and it does not feature in the PSY approach, and for these reasons we have chosen to suppress it. See PY, pp. 467-468, for further discussion. That said, autoregressive models are not time-invariant models, and indeed FGM have shown that test results for a given commodity series can be affected by the choice of data span and sampling frequency. The aspect of temporal aggregation and the choice of time span are not in the scope of the PSY paper which we leave to a future paper. PSY do demonstrate the efficacy of their tests in simulations, viewing sample size conventionally as a one-dimensional concept.

7. The significance level $\beta_T$ depends on the sample size and drops towards zero as the sample size tends to infinity.

8. We use settlement price and three-month mid prices. Prices relate to Wednesdays or, in the case that a Wednesday was a holiday, the immediately prior trading day. The LME operates a system in which a contract expires on each trading day. The consequence is that prices for successive days always relate to different contracts. The martingale property may therefore not apply in a strict manner. It is nevertheless generally regarded as an acceptable approximation. Data sources for prices - LME and the World Bureau of Metals Statistics, \textit{World Metal Statistics}.


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11. Throughout the period under consideration, China has been a major importer of both bauxite and alumina, the raw materials from which aluminium is obtained. However, the most important value added component in aluminium is the energy input in smelting. China has been able to use stranded electricity, where generating capacity has been installed distant from other industrial users, to fuel its aluminium smelting industry.

12. In a logarithmic model, an explosive process would be reflected in a non-zero intercept in the return equation, not in the autoregressive coefficient.

13. Skewness and kurtosis are both reduced by moving to price returns (changes in log prices) instead of price changes. However excess kurtosis remains substantial and normality is still rejected in each case.

14. Using 5,000 simulations, the critical value sequences generated by the Shi program exhibit a slightly jagged pattern which violates monotonicity and smoothness requirements. We modify the program by fitting a fifth order polynomial approximation to the raw critical values. The resulting critical value sequences are smooth by construction and are monotonically increasing in $r$ throughout the range $[r_0 : 1]$.


16. Convenience yield is the percentage premium of the cash price over a deferred price less the interest rate, storage cost and the rate of deterioration. It may be interpreted as the premium stockholders will pay for immediate access to inventory of known specification and location.

17. Suppose there is a period of explosive prices in the cash price of a commodity, perhaps resulting from speculation but that deferred prices are either less affected or remain unaffected. This will result in an explosion in convenience yield. Pindyck (1993) shows that convenience yield theory imposes restrictions across the cash and futures price series. Copper is the single non-ferrous metal that he investigates. These restrictions are rejected.

18. Secondary production, i.e. production from scrap, is important in the copper and lead industries but less so in the other four industries. In aluminium, scrap is generally re-melted for can production without passing through the intermediate stage of ingot.
19. Consider the following simple model. Write the market clearing price $p_t$ in period $t$ as a function of the stock level $s_t$, $p_t = f(s_t)$ where $f' < 0$ and $f'' > 0$. The stock level satisfies $s_t = s_{t-1} + x_t$, where $x_t$ is net supply and which we take to be independent of current and recent prices. Suppose $x_t$ follows a stationary AR(1) so that $x_t = \alpha x_{t-1} + \epsilon_t$, where $1 > \alpha > 0$ and the disturbance $\epsilon_t$ is serially independent. The market clearing price inherits the autoregressive property so that

$$\Delta p_t = \alpha \frac{f'(s_t)}{f'(s_{t-1})} \Delta p_{t-1} + f'(s_t) \epsilon_t = \theta(s_{t-1}, \epsilon_t) \Delta p_{t-1} + \nu_t.$$ 

A negative supply shock (positive demand shock) will result in an increase in the autoregressive coefficient $\theta$. If the coefficient $\alpha$ is sufficiently high, it is easy to obtain $\theta > 1$. A simple simulation exercise is sufficient to show that the recursive Dickey-Fuller regression of $p_t$ on $p_{t-1}$ can generate a greater than unit root over low stock periods.

20. It would be desirable to report this measure at a quarterly or monthly frequency. This faces two difficulties. First, the World Bureau of Metal Statistics (WBMS) publishes preliminary estimates of metals production and consumption at the monthly and quarterly frequency but only gives final, revised, figures at the annual frequency. Second, both production and consumption are subject to strong seasonality which has also varied over time as consumption has shifted from Europe and North America to Asia. Higher frequency measures would therefore be subject to significant uncertainty.


22. Both production and consumption may be influenced by the price in the current year although short run elasticities are generally believed to be small. Any simultaneity bias will be negative and will therefore tend to decrease the significance of the estimated logit coefficient.
REFERENCES


___________(2013b) Understanding commodity futures prices: fundamentals, financialization and bubble characteristics, preprint. [FGM]


——— (2011) Dating the timeline of financial bubbles during the subprime crisis. *Quantitative Economics* 2, 455-491. [PY]


Figure 1. LME Three Month Prices, 2000-11
Table 1
Descriptive Statistics

<table>
<thead>
<tr>
<th></th>
<th>Volatility</th>
<th>Skewness</th>
<th>Excess kurtosis</th>
<th>Normality $\chi^2$</th>
<th>Stationarity</th>
</tr>
</thead>
<tbody>
<tr>
<td>Aluminium</td>
<td>Cash</td>
<td>23.7%</td>
<td>0.20</td>
<td>2.40</td>
<td>88.6</td>
</tr>
<tr>
<td></td>
<td>3 months</td>
<td>22.9%</td>
<td>0.20</td>
<td>2.55</td>
<td>97.3</td>
</tr>
<tr>
<td>Copper</td>
<td>Cash</td>
<td>33.8%</td>
<td>-0.20</td>
<td>2.99</td>
<td>124.5</td>
</tr>
<tr>
<td></td>
<td>3 months</td>
<td>33.5%</td>
<td>-0.22</td>
<td>2.98</td>
<td>122.5</td>
</tr>
<tr>
<td>Nickel</td>
<td>Cash</td>
<td>46.8%</td>
<td>0.11</td>
<td>8.18</td>
<td>516.8</td>
</tr>
<tr>
<td></td>
<td>3 months</td>
<td>43.0%</td>
<td>-0.03</td>
<td>4.75</td>
<td>252.6</td>
</tr>
<tr>
<td>Lead</td>
<td>Cash</td>
<td>49.3%</td>
<td>-0.70</td>
<td>6.00</td>
<td>247.2</td>
</tr>
<tr>
<td></td>
<td>3 months</td>
<td>48.2%</td>
<td>-0.62</td>
<td>5.59</td>
<td>238.3</td>
</tr>
<tr>
<td>Tin</td>
<td>Cash</td>
<td>38.0%</td>
<td>-0.68</td>
<td>7.04</td>
<td>321.4</td>
</tr>
<tr>
<td></td>
<td>3 months</td>
<td>37.7%</td>
<td>-0.70</td>
<td>6.92</td>
<td>308.8</td>
</tr>
<tr>
<td>Zinc</td>
<td>Cash</td>
<td>41.9%</td>
<td>-0.59</td>
<td>5.09</td>
<td>210.7</td>
</tr>
<tr>
<td></td>
<td>3 months</td>
<td>40.2%</td>
<td>-0.65</td>
<td>4.90</td>
<td>186.5</td>
</tr>
</tbody>
</table>

Volatilities are the standard deviations of the weekly first differences of the prices divided by the means of the price levels reported on an annual basis by multiplication by $\sqrt{52}$. The skewness, excess kurtosis, normality and stationarity statistics all relate to the price differences. The normality test is that given by Doornik and Hansen (1994) and implemented in PcGive. The lag length $p$ in the ADF($p$) test is selected to minimize the Akaike Information Criterion (AIC). When $p = 0$, the test reduces to the simple DF test.
Sample: 12 January 2000 to 28 December 2011.
<table>
<thead>
<tr>
<th>Number</th>
<th>Contract</th>
<th>Start month</th>
<th>End month</th>
</tr>
</thead>
<tbody>
<tr>
<td>Aluminium</td>
<td>0</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Copper</td>
<td>2</td>
<td>Cash 3 months</td>
<td>December 2003</td>
</tr>
<tr>
<td></td>
<td></td>
<td>Cash + 3 months</td>
<td>November 2005</td>
</tr>
<tr>
<td>Nickel</td>
<td>1</td>
<td>Cash + 3 months</td>
<td>January 2007</td>
</tr>
<tr>
<td>Lead</td>
<td>2</td>
<td>Cash 3 months</td>
<td>December 2003</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td>June 2007</td>
</tr>
<tr>
<td>Tin</td>
<td>2 (cash) 3 (3 months)</td>
<td>Cash 3 months</td>
<td>February 2004</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td>April 2004</td>
</tr>
<tr>
<td></td>
<td></td>
<td>Only 3 months</td>
<td>October 2007</td>
</tr>
<tr>
<td></td>
<td></td>
<td>Cash 3 months</td>
<td>February 2008</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td>June 2008</td>
</tr>
<tr>
<td>Zinc</td>
<td>2</td>
<td>Cash + 3 months</td>
<td>November 2005</td>
</tr>
<tr>
<td></td>
<td></td>
<td>Cash 3 months</td>
<td>September 2006</td>
</tr>
</tbody>
</table>

The table reports the periods for which weakly explosive behaviour is identified using the PSY (2012) procedure modified to allow for serial dependence and excess kurtosis and in conjunction with a 5% test size.
<table>
<thead>
<tr>
<th></th>
<th>Base case GSADF</th>
<th>GSADF c.v.</th>
<th>ADF(3) + kurtosis GSADF</th>
<th>c.v.</th>
</tr>
</thead>
<tbody>
<tr>
<td>Aluminium</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Cash 3 months</td>
<td>3.31</td>
<td>3.57</td>
<td>4.14</td>
<td>3.56</td>
</tr>
<tr>
<td>Copper</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Cash 3 months</td>
<td>8.68</td>
<td>9.37</td>
<td>5.60</td>
<td>4.19</td>
</tr>
<tr>
<td>Nickel</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Cash 3 months</td>
<td>3.81</td>
<td>4.36</td>
<td>6.93</td>
<td>4.02</td>
</tr>
<tr>
<td>Lead</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Cash 3 months</td>
<td>6.19</td>
<td>6.74</td>
<td>5.07</td>
<td>3.62</td>
</tr>
<tr>
<td>Tin</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Cash 3 months</td>
<td>6.01</td>
<td>5.81</td>
<td>6.02</td>
<td>3.61</td>
</tr>
<tr>
<td>Zinc</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Cash 3 months</td>
<td>10.02</td>
<td>10.04</td>
<td>10.02</td>
<td>3.61</td>
</tr>
</tbody>
</table>

The table reports GSADF statistics for cash and 3 month prices (column 1) and the base PSY critical values (column 2). Columns 3 and 4 repeat the tests allowing for serial correlation and kurtosis. In the column 3 tests, critical values are simulated using autoregressive coefficients and kurtosis levels estimated from a prior AR(3). Common parameter values are used for both cash and three month prices except in the case of nickel where separate critical values are simulated for the cash and 3 month processes.
<table>
<thead>
<tr>
<th>Year</th>
<th>Aluminium</th>
<th>Copper</th>
<th>Nickel</th>
<th>Lead</th>
<th>Tin</th>
<th>Zinc</th>
</tr>
</thead>
<tbody>
<tr>
<td>2000</td>
<td>0.922</td>
<td>0.939</td>
<td>0.951</td>
<td>0.616</td>
<td>0.747</td>
<td>0.891</td>
</tr>
<tr>
<td>2001</td>
<td>0.885</td>
<td>0.882</td>
<td>0.910</td>
<td>0.907</td>
<td>0.750</td>
<td>0.885</td>
</tr>
<tr>
<td>2002</td>
<td>0.878</td>
<td>0.886</td>
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<td>0.952</td>
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<td>0.866</td>
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</table>

The table gives the ratio of consumption of each metal to total supply in the same year, where supply is production plus the closing stock in the previous year.

Figure 2: Aluminium BSADF sequence  
(critical values calibrated for Student distribution with 6.5 degrees of freedom)

Figure 3: Copper BSADF sequence  
(critical values calibrated for Student distribution with 6.5 degrees of freedom)
Figure 4: Nickel BSADF sequence
(critical values calibrated for Student distribution with 5.3 degrees of freedom)

Figure 5: Lead BSADF sequence
(critical values calibrated for Student distribution with 5.25 degrees of freedom)
Figure 6: Tin BSADF sequence
(critical values calibrated for Student distribution with 4.85 degrees of freedom)

Figure 7: Zinc BSADF sequence
(critical values calibrated for Student distribution with 5.25 degrees of freedom)