SPECULATIVE BUBBLES IN MATURE STOCK MARKETS: DO THEY EXIST AND ARE THEY RELATED?

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Abstract

Economists have long conjectured that movements in stock prices may involve speculative bubbles. A speculative bubble is usually defined as the difference between the market value of a security and its fundamental value. Although there are several important theoretical issues surrounding the topic of asset bubbles, their existence is inherently an empirical issue that has not been settled. This paper proposes a new methodology for testing for the existence of rational bubbles. Unlike previous authors, we treat both the dividend and the bubble process as part of the state vector. The new methodology is applied to the four mature markets of the U.S., Japan, England and Germany to test whether a bubble was present during the period of January 1951 to December 1998. This paper also examines whether there are linkages between these national bubbles. We find evidence that U.S. bubbles cause bubbles in the other three markets but we find no evidence for reverse causality.

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Abstract

Economists have long conjectured that movements in stock prices may involve speculative bubbles. A speculative bubble is usually defined as the difference between the market value of a security and its fundamental value. Although there are several important theoretical issues surrounding the topic of asset bubbles, their existence is inherently an empirical issue that has not been settled. This paper proposes a new methodology for testing for the existence of rational bubbles. Unlike previous authors, we treat both the dividend and the bubble process as part of the state vector. The new methodology is applied to the four mature markets of the U.S., Japan, England and Germany to test whether a bubble was present during the period of January 1951 to December 1998. This paper also examines whether there are linkages between these national bubbles. We find evidence that U.S. bubbles cause bubbles in the other three markets but we find no evidence for reverse causality.

1. INTRODUCTION

Many economists accept market efficiency, surveyed recently in Campbell (2000), as the well-established paradigm of financial economics but also acknowledge that asset prices are too volatile. For example, NASDAQ, from its peak on March 10, 2000 when it stood at 5048.62 to its low of 1454.04 on September 21, 2001, declined by 71.25%. Is this significant decline caused only because of substantial revisions of the expected payoffs and/or changes in the discount factor?

This paper modifies certain methodologies employed to test for bubbles and explores the existence of relationships among bubbles in mature stock markets. Section 2 reviews the general ideas of asset bubbles and section 3 reviews global stock market integration. In section 4, we motivate our methodological contribution by reviewing the important contribution of Wu (1995, 1997) and in section 5 we highlight our new methodology for testing for the existence of asset bubbles and how it is applied to the stock markets of the U.S., Japan, England and Germany. We find evidence of asset bubbles in all four stock markets. We then proceed to examine whether these bubbles travel across mature economies. In sections 6,7 and 8 we elaborate in detail our methodological procedures for testing for linkages between bubbles in mature stock markets. Our main findings and conclusions are given in the last section.

2. ASSET BUBBLES

Economists have long conjectured that movements in stock prices may involve speculative bubbles, as trading is often said to generate over-priced or under-priced markets. Some financial economists, however, believe that stock price fluctuations reflect changes in the values of the underlying market fundamentals. The standard definition of fundamental value is the summed discounted value of all future cash flows. The difference, if any, between the market value of the security and its fundamental value is termed a speculative bubble. Yet confusion persists about what factors generate bubbles. Fads and irrationality have always figured prominently, and the hypothesis that these factors are important has gained some empirical support from the literature on asset price volatility. Another bubble-producing factor is the structure of information in the market. In a partial-equilibrium setting, Allen and Gorton (1991) showed that rational bubbles could exist with a finite number of agents who had asymmetric information.

The existence of bubbles is inherently an empirical issue that has not been settled. A number of studies such as Blanchard and Watson (1982) and West (1988) have argued that dividend and stock price data are not consistent with the "market fundamentals" hypothesis, in which prices are given by the present discounted values of expected dividends. These results have often been construed as evidence for the existence of bubbles or fads.

According to Shiller (1981), and LeRoy and Porter (1981) the variability of stock price movements is too large to be explained by the discounted present value of future dividends. Over the past century U.S. stock prices are five to thirteen times more volatile than can be justified by new information about future dividends. Campbell and Shiller (1988a, b) and West (1987, 1988) remove the assumption of a constant discount rate. However, a variable discount rate provides only marginal support in explaining stock price volatility. These authors reject the null hypothesis of no bubbles. See also Rappoport and White (1993, 1994).

A major problem with such arguments is that evidence for bubbles can be reinterpreted in terms of market fundamentals that are unobserved by the econometricians as argued by Flood and Garber (1980), Hamilton and Whiteman (1985) and Hamilton (1986). Diba and Grossman (1984, 1988a, b) have recommended an alternative strategy for testing for rational bubbles by investigating the stationarity properties of asset prices and observable fundamentals. In essence, the argument for equities is that if stock prices are not more explosive than dividends, then it can be concluded that rational bubbles are not present, since they would generate an explosive component in stock prices. Using unit-root tests, autocorrelation patterns, and cointegration tests to implement this procedure, Dezhbakhsh and Demirguc-Kunt (1990) reach the conclusion that stock prices do not contain explosive rational bubbles. Evans (1991) criticizes tests for bubbles based on an investigation of the stationarity properties of stock prices and dividends. He demonstrates by Monte-Carlo simulations that an important class of rational bubbles cannot be detected by these tests even though the bubbles are explosive.

Wu (1997) examines a rational bubble, able to burst and restart continuously. The specification is parsimonious and allows easy estimation. The model fits the data reasonably well, especially during several bull and bear markets in this century. Such rational bubbles can explain much of the deviation of U.S. stock prices from the simple present-value model. Wu's work is reviewed in more detail in the next section.

Most of the references above address issues of asset bubbles either theoretically or in the context of a mature economy. To complete this rapid bibliographical review we need to mention three additional trends in this literature. First several studies have tested for the existence of bubbles in emerging markets. For example, Richards (1996) claims that emerging markets have not consistently been subject to fads or bubbles. Chan and coauthors (1998) test for rational bubbles in Asian stock markets and Chen (1999) specializes his search for bubbles in the Hong Kong market. Sarno and Taylor (1999) find evidence of bubbles in all East Asian economies.

Second, significant increases in cross-markets linkages, after a shock, have become a topic of current research under the term "contagion". Several important papers collected in Claessens and Forbes (2001) discuss both methodological issues and case studies of contagion.

Third, beyond the existence or not of bubbles, economists have also studied in detail the implications of a stock market bubble to the economy at large. Biswanger (1999) offers a comprehensive review of these issues and Chirinko and Schaller (1996) argue that bubbles have existed over certain periods in the US stock market but real investment decisions have been determined by fundamentals.

3. GLOBAL STOCK MARKET INTEGRATION

Once bubbles are confirmed empirically in the four mature stock markets, we proceed to test linkages between these markets in terms of both the fundamental price and the bubble price. In this context we adopt a sub-set VAR methodology presented in Lutkepohl (1993, p. 179). The approach builds into it the causal relations between the series and this gives us the opportunity to analyze the potential global interaction among these national equity markets through the speculative component of the prices. The potential existence of global linkages among equity markets has attracted great interest among scholars because of its impact on global diversification.

During the past thirty years, world stock markets have become more integrated, primarily because of financial deregulation and advances in computer technology. Financial researchers have examined various aspects of the evolution of this particular aspect of world integration. For example, the early studies by Grubel (1968), Levy and Sarnat (1970), Grubel and Fadner (1971), Agmon (1972, 1973), Ripley (1973), Solnik (1974a), and Lessard (1973, 1974, 1976) have investigated the benefits from international portfolio diversification. While some studies, such as Solnik (1974b), were exclusively theoretical in extending the capital asset pricing model to a world economy, others such as Levy and Sarnat (1970) used both theory and empirical testing to confirm the existence of financial benefits from international diversification.

Similar benefits were also confirmed by Grubel (1968), Grubel and Fadner (1971), Ripley (1973), Lessard (1973, 1974, 1976), Agmon (1972, 1973), Makridakis and Wheelwright (1974), and others, who studied the relations among equity markets in various countries. Specifically, Agmon (1972, 1973) investigated the relationships among the equity markets of the U.S., United Kingdom, Germany and Japan, while Lessard (1973) considered a group of Latin American countries.

By 1976, eight years after the pioneering work of Grubel (1968), enough knowledge had been accumulated on this subject to induce Panton, Lessing and Joy (1976) to offer a taxonomy. It seems reasonable to argue that although these studies had used different methodologies and diverse data from a variety of countries, their main conclusions confirmed that correlations among national stock market returns were low and that national speculative markets were largely responding to domestic economic fundamentals.

Theoretical developments on continuous time stochastic processes and arbitrage theory were quickly incorporated into international finance. Stulz (1981) developed a continuous time model of international asset pricing while Solnik (1983) extended arbitrage theory to an international setting. Adler and Dumas (1983) integrated international portfolio choice and corporate finance. Empirical research also continued to flow such as Hilliard (1979), Moldonado and Saunders (1981), Christofi and Philippatos (1987), Philippatos, Christofi and Christofi (1983) and also Grauer and Hakansson (1987), Schollhammer and Sand (1987), Wheatley (1988), Eun and Shim (1989), von Furstenberg and Jeon

(1989), Becker, Finnerty and Gupta (1990), Fisher and Palasvirta (1990), French and Poterba (1991) and Harvey (1991).

These numerous studies employ various recent methodologies and larger databases than the earlier studies to test for interdependencies between the time series of national stock market returns. The underlying issue remains the empirical assessment of how much integration exists among national stock markets. In contrast to earlier results, and despite some reservations, several of these new studies find high and statistically significant level of interdependence between national markets supporting the hypothesis that global stock markets are becoming more integrated.

In comparing the results of the earlier studies with those of the more recent ones, one could deduce that greater global integration implies lesser benefits from international portfolio diversification. If this is true, how can one explain the ever-increasing flow of big sums of money invested in international markets? To put differently, while Tesar and Werner (1992) confirm the home bias in the globalization of stock markets, why are increasing amounts of funds invested in non-home equity markets?

The analysis of the October 19, 1987 stock market crash may offer some insight in answering this question. Roll (1988, 1989), King and Wadhwani (1990), Hamao, Musulis and Ng (1990) and Malliaris and Urrutia (1992) confirm that almost all stock markets fell together during the October 1987 crash despite the existing differences of the national economies while no significant interrelationships seem to exist for periods prior and post the crash. Malliaris and Urrutia (1997) also confirm the simultaneous fall of national stock market returns because of the Iraqi invasion of Kuwait in July 1990. This evidence supports the hypothesis that certain global events, such as the crash of October 1987 or the Invasion of Kuwait in July 1990, tend to move world equity markets in the same direction, thus reducing the effectiveness of international diversification. On the other hand, in the absence of global events, national markets are dominated by domestic fundamentals, and international investing increases the benefits of diversification. Exceptions exist, as in the case of regional markets, such as the European stock markets reported in Malliaris and Urrutia (1996). Longin and Solnik (2001) distinguish between bear and bull markets in international equity markets and find that correlation increases in bear markets, but not in bull markets.

A review of the literature on linkages among international stock markets can be found in McCarthy and Najand (1995). These authors adopt the state space methodology to infer the linkage relationships between the stock markets in Canada, Germany, Japan, U.K and the U.S. The authors claim that this approach not only determines the causal relationship, in the Granger sense, but it delivers the result with minimum number of parameters necessary. They report that the U.S. market exerts the most influence on other markets. Since these authors use daily data there is some overlap in the market trading time and they attempt to take care of that in the interpretation of their results. The main finding is consistent with similar findings by other researchers, such as, Eun and Shim (1989), who examine nine stock markets in the North America and Europe over period 1980-1985 in a VAR framework.

From this rapid review of global stock market integration, it becomes apparent that the topic of linkages between bubbles has not been addressed. Our methodology for testing the existence of bubbles in national markets has the additional advantage that it renders itself for also testing for possible linkages between bubbles in these mature stock markets. We augment our contribution to the literature by exploring this issue also.

4. REVIEW OF KEY EMPIRICAL PAPERS

To motivate our methodological contribution to testing for asset bubbles, we review two influential papers in this area.

4.1. Wu (1997)

The paper estimates a rational stochastic bubble using the Kalman filtering technique. The bubble grows at the discount rate in expectation and it can collapse and restart continuously, allowing for the possibility of a negative bubble. The log dividends follows a general ARIMA (p, 1, q) process. The model for stock prices with the bubble component, the dividend process and the bubble process are expressed in the state-space form with the bubble being treated as an unobserved state vector. The model parameters are estimated by the method of maximum likelihood and obtain optimal estimates of stochastic bubbles through the Kalman filter.

Consider the standard linear rational expectations model of stock price determination:

$$\left[E_t(P_{t+1} + D_t) - P_t\right] / P_t = r\,, \tag{1}$$

where p=real stock price at time t, D=the real dividend at time t, E=the mathematical expectation conditional on information available at time t and r=the required real rate of return, r>0. The log-linear approximation of (1) can be written as follows:

$$q = k + \psi E_t p_{t+1} + (1-\psi)d_t - p_t\,, \tag{2}$$

where, q=required log gross return rate, Ψ=average ratio of the stock price to the sum of the stock price and the dividend, k=-ln(Ψ)-(1-Ψ)ln(1/Ψ-1), p=ln(P) and d=ln(D).

The general solution to (2) is given by:

$$p_t = (k-q)/(1-\psi) + (1-\psi)\sum_{1=0}^{\infty}\psi^i E_t(d_{t+i}) + b_t = p_t^f + b_t \tag{3}$$

where b_t satisfies the following homogeneous difference equation:

$$E_t(b_{t+i}) = (1/\psi)^i b_t\,. \tag{4}$$

In equation (2), the no-bubble solution p is exclusively determined by dividends, while b can be driven by events extraneous to the market and is referred to as a rational speculative bubble. After defining the stock price equation, the parametric bubble process and the dividend process in a state-space form, the bubble is treated as an unobserved state vector, which can be estimated by the Kalman filtering technique.

Wu finds statistically significant estimate of the innovation variance for the bubble process. During the bull market of the 1960s the bubble accounts for between 40% to 50% of the actual stock prices. Negative bubbles are found during the 1919-1921 bear market, in which the bubble explains between 20% to 30% of the decline in stock prices.

4.2 Wu (1995)

The model just discussed has been also used by the same author to estimate the unobserved bubbly component of the exchange rate and to test whether it is significantly different from zero. Using the monetary model of exchange rate determination, the solution for the exchange rate is the sum of two components. The first component, called the fundamental solution, is a function of the observed market fundamental variables. The second component is an unobserved process, which satisfies the monetary model and is called the stochastic bubble. The monetary model, the market fundamental process and the bubble process are expressed in the state-space form, with the bubble being treated as a state variable. The Kalman filter can then be used to estimate the state variable.

The author finds no significant estimate of a bubble component during the period 1974-1988. Similar results were obtained for the sub-sample, 1981 through 1985, during which the US dollar appreciated most drastically and a bubble might have occurred.

5. OUR METHODOLOGICAL CONTRIBUTION

The purpose of our study is to search empirically for bubbles in national stock markets using a state-of-the-art methodology with emphasis on the U.S., Japan, Germany and the United Kingdom. We focus on the post-war period in these four countries as opposed to the authors reviewed in the previous section who concentrate on only the U.S. All data are monthly returns of the S&P 500, Nikkei 225, Dax-30 and FT-100 indexes ranging from January 1951 to December 1998, that is, 576 observations. All data are converted to real values using the corresponding CPI measures and Global Financial Data provided the data. In order to establish the soundness of our methodology we have reproduced the results from Wu (1997) using annual U.S. data (also obtained from Global Financial Data) covering the period 1871 - 1998.

Although we employ the unobserved component modelling approach that is similar to Wu (1997), our implementation is quite different. We follow Shumway and Stoffer (2000, p. 306) to develop a Dynamic Linear Model, DLM, to treat both the dividend process and the bubble process as part of the unobserved components, that is, the state vector. The state equations also include their own system errors, which are assumed uncorrelated. The measurement vector in this case contains the price and the realised dividend without any measurement errors. The advantage of this way of modeling is that the comparison with the no bubble solution becomes much more straightforward. Wu (1997) had to resort to alternative way (GMM) of estimating the no bubble solution and the model adequacy tests are not performed there. Besides, the precise moment conditions used in the GMM estimation are not reported there. On the other hand, in our approach we are able to subject both the bubble and the no bubble solutions to a battery of diagnostics test applicable to state space systems. In sections 6,7 and 8 we describe in detail the mathematical structures of our models and the estimation strategies.

6. DYNAMIC LINEAR MODELS FOR BUBBLE SOLUTIONS

Our starting point are equations (3) and (4) described earlier. As our preliminary investigations reveal that both the log real price and log real dividend series are non-stationary, we choose to work with the first differenced series. Thus, equation (3) becomes,

$$\Delta p_t = \Delta p_t^f + \Delta b_t \tag{5}$$

where, $\Delta p_t^f \equiv (1-\psi)\sum_{i=0}^{\infty}\psi^i E_t\left(d_{t+i}\right) - (1-\psi)\sum_{i=0}^{\infty}\psi^i E_{t-1}\left(d_{t-1+i}\right)$. Assume the parametric representation of equation (4) is

$$b_{t+1} = \frac{1}{\psi} b_t + \varepsilon_\eta,\ \varepsilon_\eta \sim N\left(0, \sigma_\eta^2\right), \tag{6}$$

$$\Delta b_t = \frac{1}{\psi}\left(b_t - b_{t-1}\right). \tag{7}$$

In order to express the fundamental component of the price, Δp_t^f, in terms of the dividend process, we fit an appropriate AR model of sufficient order so that the Akaike information criterion, AIC, is minimized. We find that for the Japanese data a AR(1) model is sufficient whereas for the other three countries we need AR(3) models. The infinite sums in the expression for Δp_t^f may be expressed in terms of the parameters of the dividend process once we note the following conditions:

- The differenced log real dividend series is stationary, therefore the infinite sum converges,
- Any finite order AR process can be expressed in companion form (VAR of order 1) by using extended state variables, i.e. suitable lags of the original variables, (Campbell, Lo and MacKinlay (1997), p. 280),
- Using demeaned variables the VAR(1) process can be easily used for multiperiod ahead forecast (Campbell, Lo and MacKinlay (1997), p. 280).

Assuming the demeaned log real dividend process has the following AR(3) representation,

$$\Delta d_t = \phi_1 \Delta d_{t-1} + \phi_2 \Delta d_{t-2} + \phi_3 \Delta d_{t-3} + \varepsilon_\delta,\ \varepsilon_\delta \sim N\left(0, \sigma_\delta^2\right), \tag{8}$$

the companion form may be written as,

$$\begin{bmatrix} \Delta d_t \\ \Delta d_{t-1} \\ \Delta d_{t-2} \end{bmatrix} = \begin{bmatrix} \phi_1 & \phi_2 & \phi_3 \\ 1 & 0 & 0 \\ 0 & 1 & 0 \end{bmatrix} \begin{bmatrix} \Delta d_{t-1} \\ \Delta d_{t-2} \\ \Delta d_{t-3} \end{bmatrix} + \begin{bmatrix} \varepsilon_\delta \\ 0 \\ 0 \end{bmatrix}, \text{ or} \tag{9}$$

$$X_t = \Phi X_{t-1} + \Xi_t, \tag{10}$$

where, the definitions of X_t, Φ, and Ξ_t are obvious from comparison of equations (9) and (10). Following Campbell, Lo and MacKinlay (1997, p. 280), Δp_t^f may be expressed as, (with I being the identity matrix of the same dimension as Φ)

$$\Delta p_t^f = \Delta d_t + \psi\Phi\left(I - \psi\Phi\right)^{-1}\Delta X_t . \tag{11}$$

We can now express equation (5) in terms of the fundamental component and the bubble component,

$$\Delta p_t = \Delta d_t + e'\psi\Phi\left(I - \psi\Phi\right)^{-1}\Delta X_t + \Delta b_t , \tag{12}$$

where $e' \equiv \begin{bmatrix} 1 & 0 & 0 \end{bmatrix}$

Equation (12) represents the measurement equation of the DLM and we need to suitably define the state equation for the model. An examination of equations (7) and (9) suggests that the following state equation represent the dynamics of the dividend and the bubble process:

$$\begin{bmatrix} \Delta d_t \\ \Delta d_{t-1} \\ \Delta d_{t-2} \\ \Delta d_{t-3} \\ b_t \\ b_{t-1} \end{bmatrix} = \begin{bmatrix} \phi_1 & \phi_2 & \phi_3 & 0 & 0 & 0 \\ 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & \frac{1}{\psi} & 0 \\ 0 & 0 & 0 & 0 & 1 & 0 \end{bmatrix} \begin{bmatrix} \Delta d_{t-1} \\ \Delta d_{t-2} \\ \Delta d_{t-3} \\ \Delta d_{t-4} \\ b_{t-1} \\ b_{t-2} \end{bmatrix} + \begin{bmatrix} \varepsilon_\delta & 0 \\ 0 & 0 \\ 0 & 0 \\ 0 & 0 \\ 0 & \varepsilon_\eta \\ 0 & 0 \end{bmatrix}, \tag{13}$$

$$\begin{pmatrix} \varepsilon_\delta \\ \varepsilon_\eta \end{pmatrix} \sim N\left(\begin{bmatrix} 0 \\ 0 \end{bmatrix}, \begin{bmatrix} \sigma_\delta^2 & 0 \\ 0 & \sigma_\eta^2 \end{bmatrix} \right). \tag{13'}$$

We are in a position now to define the measurement equation of the DLM in terms of the state vector in equation (13). This is achieved by examining equation (12) and defining a row vector, $M \equiv e'\psi\Phi\left(I - \psi\Phi\right)^{-1} = \left[m_1, m_2, m_3\right]$, as follows:

$$\Delta p_t = \Delta d_t + \left[m_1, m_2, m_3\right] \begin{bmatrix} \Delta d_t - \Delta d_{t-1} \\ \Delta d_{t-1} - \Delta d_{t-2} \\ \Delta d_{t-2} - \Delta d_{t-3} \end{bmatrix} + \Delta b_t \text{, or}$$

$$\begin{bmatrix}\Delta p_t\\ \Delta d_t\end{bmatrix}=\begin{bmatrix}(1+m_1) & (m_2-m_1) & (m_3-m_2) & -m_3 & 1 & -1\\ 1 & 0 & 0 & 0 & 0 & 0\end{bmatrix}\begin{bmatrix}\Delta d_t\\ \Delta d_{t-1}\\ \Delta d_{t-2}\\ \Delta d_{t-3}\\ b_t\\ b_{t-1}\end{bmatrix}. \tag{14}$$

Equation (14) determines the measurement equation of the DLM without any measurement error. In other words, the evolution of the state vector in equation (13) results in the measurement of the measurement vector through equation (14). Equations (13) and (14) represent the DLM for the bubble solution when the dividend process is described by the AR(3) system in equation (9). In our sample this is the case for Germany, U.K. and the U.S. Since the data for Japan required only a AR(1) process for the dividend in equation (9), the DLM, in this case, may be written directly as:

$$\begin{bmatrix}\Delta d_t\\ \Delta d_{t-1}\\ b_t\\ b_{t-1}\end{bmatrix}=\begin{bmatrix}\phi_1 & 0 & 0 & 0\\ 1 & 0 & 0 & 0\\ 0 & 0 & \frac{1}{\psi} & 0\\ 0 & 0 & 1 & 0\end{bmatrix}\begin{bmatrix}\Delta d_{t-1}\\ \Delta d_{t-2}\\ b_{t-1}\\ b_{t-2}\end{bmatrix}+\begin{bmatrix}\varepsilon_\delta & 0\\ 0 & 0\\ 0 & \varepsilon_\eta\\ 0 & 0\end{bmatrix}, \tag{15}$$

$$\begin{pmatrix}\varepsilon_\delta\\ \varepsilon_\eta\end{pmatrix}\sim N\left(\begin{bmatrix}0\\0\end{bmatrix},\begin{bmatrix}\sigma_\delta^2 & 0\\ 0 & \sigma_\eta^2\end{bmatrix}\right). \tag{15'}$$

Similarly, the measurement equation for the DLM of the bubble solution for the Japanese data becomes,

$$\begin{bmatrix}\Delta p_t\\ \Delta d_t\end{bmatrix}=\begin{bmatrix}(1+m_1) & -m_1 & 1 & -1\\ 1 & 0 & 0 & 0\end{bmatrix}\begin{bmatrix}\Delta d_t\\ \Delta d_{t-1}\\ b_t\\ b_{t-1}\end{bmatrix}, \tag{16}$$

where, $M \equiv e'\psi\Phi\left(I-\psi\Phi\right)^{-1}=\left[m_1\right]$, since $e'=[1]$, $\Phi=\left[\phi_1\right]$.

We have now set up the DLM for the bubble solution for Germany, U.K., and the U.S. given by equations (13) and (14). For the Japanese data, on the other hand, these are given by equations (15) and (16). The parameters of the models embedded in these equations and both the filtered and the smoothed estimates of the bubble series are to be estimated from the observed price and the dividend series. The details of the estimation procedure are described in appendix A. In the next section we proceed to set up the DLMs for the no-bubble solutions.

7. DYNAMIC LINEAR MODELS FOR NO-BUBBLE SOLUTIONS

In order to compare the performance of the bubble solution discussed in the previous section we develop the DLM for a no-bubble solution. We maintain the same framework so that a comparison can be more meaningful. This is in contrast to the approach taken by Wu (1997) where the no-bubble solution was estimated in the GMM framework. We also note that the model should account for the correlations in the variance of the stock return series. This is done by incorporating the GARCH(1,1) effect in the price equation (10) without the bubble component. In this context we adopt the methodology of Harvey, Ruiz and Sentana (1992) and follow Kim and Nelson (1999, page 144) to suitably augment the state vector of the DLM.

For Germany, U.K. and the U.S.A date set the state equation (13) becomes,

$$\begin{bmatrix} \Delta d_t \\ \Delta d_{t-1} \\ \Delta d_{t-2} \\ \Delta d_{t-3} \\ \varepsilon_{p,t} \end{bmatrix} = \begin{bmatrix} \phi_1 & \phi_2 & \phi_3 & 0 & 0 \\ 1 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 \end{bmatrix} \begin{bmatrix} \Delta d_{t-1} \\ \Delta d_{t-2} \\ \Delta d_{t-3} \\ \Delta d_{t-4} \\ \varepsilon_{p,t-1} \end{bmatrix} + \begin{bmatrix} \varepsilon_\delta & 0 \\ 0 & 0 \\ 0 & 0 \\ 0 & 0 \\ 0 & \varepsilon_{p,t} \end{bmatrix}, \tag{17}$$

$$\begin{pmatrix} \varepsilon_\delta \\ \varepsilon_{p,t} | \omega_{t-1} \end{pmatrix} \sim N\left(\begin{bmatrix} 0 \\ 0 \end{bmatrix}, \begin{bmatrix} \sigma_\delta^2 & 0 \\ 0 & h_t \end{bmatrix} \right), \; h_t = \alpha_0 + \alpha_1 \varepsilon_{p,t-1}^2 + \beta_1 h_{t-1}, \tag{17'}$$

and ω_{t-1} is the information set at time t-1. The corresponding measurement equation becomes,

$$\begin{bmatrix} \Delta p_t \\ \Delta d_t \end{bmatrix} = \begin{bmatrix} (1+m_1) & (m_2 - m_1) & (m_3 - m_2) & -m_3 & 1 \\ 1 & 0 & 0 & 0 & 0 \end{bmatrix} \begin{bmatrix} \Delta d_t \\ \Delta d_{t-1} \\ \Delta d_{t-2} \\ \Delta d_{t-3} \\ \varepsilon_{p,t} \end{bmatrix}. \tag{18}$$

For the Japanese data with an AR(1) dividend process, the no-bubble DLM may be written following the approach above. The state equation (15) becomes,

$$\begin{bmatrix} \Delta d_t \\ \Delta d_{t-1} \\ \varepsilon_{p,t} \end{bmatrix} = \begin{bmatrix} \phi_1 & 0 & 0 \\ 1 & 0 & 0 \\ 0 & 0 & 0 \end{bmatrix} \begin{bmatrix} \Delta d_t \\ \Delta d_{t-1} \\ \varepsilon_{p,t-1} \end{bmatrix} + \begin{bmatrix} \varepsilon_\delta & 0 \\ 0 & 0 \\ 0 & \varepsilon_{p,t} \end{bmatrix}, \tag{19}$$

$$\begin{pmatrix} \varepsilon_\delta \\ \varepsilon_{p,t} \,|\omega_{t-1} \end{pmatrix} \sim N\left(\begin{bmatrix} 0 \\ 0 \end{bmatrix}, \begin{bmatrix} \sigma_\delta^2 & 0 \\ 0 & h_t \end{bmatrix} \right),\; h_t = \alpha_0 + \alpha_1 \varepsilon_{p,t-1}^2 + \beta_1 h_{t-1}. \tag{19'}$$

The corresponding measurement (21) becomes,

$$\begin{bmatrix} \Delta p_t \\ \Delta d_t \end{bmatrix} = \begin{bmatrix} (1+m_1) & -m_1 & 1 \\ 1 & 0 & 0 \end{bmatrix} \begin{bmatrix} \Delta d_t \\ \Delta d_{t-1} \\ \varepsilon_{p,t} \end{bmatrix}. \tag{20}$$

In the no-bubble solutions, the parameters to be estimated are those of the dividend process and the GARCH(1,1) coefficients. The procedure for this is the same as that for the bubble solutions and is described in detail in appendix A. The next section takes up the issues in modeling the linkages between the markets in the subset VAR framework.

8. SUBSET VAR FRAMEWORK FOR ESTABLISHING LINKAGES BETWEEN MARKETS

The methodology developed in this paper allows us to decompose the stock prices in their fundamental and the bubble components. We, analyze the linkage relationship both through the fundamental as well as through the speculative component. This helps us understand whether the market linkages are through the fundamental or through the speculative components of the price series. Also, since we are dealing with monthly data, the time overlap problem between markets is largely non-existent.

The econometric procedure we adopt is referred to as the subset VAR. Use of standard VAR approach to study causal relations between variables is frequently employed. A typical VAR model involves a large number of coefficients to be estimated and thus estimation uncertainty remains. Some of the coefficients may in fact be zero. When we impose zero constraints on the coefficients in full VAR estimation problem what results is the subset VAR. But, since most often no a priori knowledge is available that will guide us to constrain certain coefficients, we base the modeling strategy on information provided by the AIC (Akaike Information Criterion) and the HQ (Hannan-Quinn) model selection criteria. Actual mathematical definitions and the details of this approach can be found in Lutkepohl (1993, chapter 5). Below we describe this procedure very briefly.

We first obtain the order of the VAR process for the four variables using the information criterion mentioned above. The top-down strategy starts from this full VAR model and the coefficients are deleted one at a time (from the highest lag term) from the four equations separately. Each time a coefficient is deleted the model is estimated using least-square algorithm and the information criterion is compared with the previous minimum one. If the current value of the criterion is greater than the previous minimum value, the coefficient is maintained otherwise it is deleted. The process is repeated for each of the four equations in the system. Once all the zero restrictions are determined the final set of equations are estimated again which gives the most parsimonious model. We also check for the adequacy of this model by examining the multivariate version of the portmanteau test for whiteness of the residuals as suggested by Lutkepohl (1993, p. 188). Once the subset VAR model is estimated there is no further need for testing causal relations and/or linkages between the variables. The causality testing is built into the model development process. Therefore, we examine linkages between the four markets in our study using this subset VAR model.

As mentioned earlier we explore linkages between these markets in two stages. In the first stage, the fundamental price series are all found to be stationary, and hence in this case the modeling is done using the levels of the variables. We find evidence of one unit root in the speculative components of the price series for all the four markets. As we suspect existence of a cointegrating relation between these speculative components, we explore this using Johansen's cointegration test and find evidence of one cointegrating vector. It is, therefore, natural to estimate a vector error correction model, which is essentially a restricted VAR model with the cointegrating relation designed into it. As suggested in Lutkepohl (1993, p. 378) we examine the causal relation between these variables in the same way as for a stable system. In other words, we explore the linkages as for the fundamental price component but in this case we use first differenced form and use the lagged values of the cointegrating vector as well.

9. DISCUSSION OF RESULTS

First we discuss the estimations results of the dynamic linear model with the bubble solution for the annual U.S. data series. In Table 1 we find all the parameter estimates are statistically significant. The significance of the parameter, σ_η, implies highly variable bubble component of the price throughout the period 1871 to 1998. The parameters describing the real dividend process are very close to the univariate estimation (not included) results of the dividend series. Besides, the discount parameter, ψ, is close to its sample value.

In Table 2 we present the estimates of the no-bubble solution with a GARCH (1,1) error structure for the price equation for the same U.S. annual data. Here also, most of the parameters are statistically significant. The significance of the GARCH parameter, β_1, implies persistence in the residual volatility. This model is used to compare the results of the bubble solution. We would like to stress the fact we implemented the GARCH (1,1) model also in the state space framework so that the comparison with the bubble solution would be more realistic. This is, however, not the case with Wu (1997), who uses the GMM methodology. This approach also allows us to check the performance of both models by analyzing the residual diagnostics. We present these test results in Table 3. The portmanteau tests support the whiteness of the residuals and the ARCH tests indicate no remaining heteroscedasticity in the residuals. Besides, the Kolmogorov-Smirnov tests support the normality of the residuals. These three tests overwhelmingly support the modeling approach adopted here and, therefore, the conclusions drawn are statistically meaningful.

In addition to the three tests just outlined above, we also include two additional tests particularly designed for recursive residuals produced by the dynamic linear systems developed in this study. The first is the modified von-Neumann ratio test against serial correlations in the residuals and the second is the recursive t-test to check for correct model specification. As the entries in Table 3 suggest the dynamic linear models of the bubble and the no-bubble solution perform extremely well with respect to these two tests. There is strong support for the adequacy of the models in describing the price process.

In view of these tests, we can now proceed to analyze the rest of the results. As discussed in Wu (1997) the rational stochastic bubble can alternate between positive and negative values. Wu argues that stocks may be overvalued when the participants are bullish and may be undervalued when the participants are bearish. Figure 1 shows negative bubble in the very early part of the sample as well as during the early 1920s. It is obvious though that the stochastic bubbles account for a substantial percentage of the stock price in the sample. It is also interesting to note that in spite of the drop in the bubble percentage during the oil shock of the1970's and the stock market crash of 1987 there has been a upward trend of the bubble percentage throughout the latter part of the sample period considered. This particular feature is most clearly visible in the lower panel of the Figure 1, which separates the stock price in the fundamental and the bubble components.

Next, we compare the performance of the bubble and the no-bubble solutions by examining the in sample fitting of the stock prices. In Table 4 we display the criteria used and these are defined as,

$$\text{RMSE} = \frac{1}{T}\sum_{t=1}^{T}(\hat{p}_t - p_t)^2, \ \text{MAE} = \frac{1}{T}\sum_{t=1}^{T}|\hat{p}_t - p_t|,$$

where $\hat{p}_t$ is the fitted price and T is the number of observations. The entries in Table 4 clearly demonstrate the superiority of the bubble solution to capture the price process over the sample period.

We next analyze the monthly data, covering the post war period, for the four mature stock markets of Germany, Japan, U.K. and the U.S. In Table 5 we present the estimation results of the bubble solutions. It is clear that most of the parameters are statistically significant. The discount parameter, ψ, as before is close to the respective sample values while the significant σ_η for all the four countries imply highly variable bubble components. Needless to say that the estimated parameters of the dividend processes are close to their respective uni-variate estimation (not reported here) results. As is evident from Table 6, the significant ARCH and the GARCH parameters indicate appropriateness of the error specification for the log price difference series. There is substantial persistence in the variance process.

We now move to analyze the residual diagnostics in order to ascertain the appropriateness of the model for the monthly data series for all four countries. As with the annual data (for the U.S.) we find evidence of whiteness on residuals from the portmanteau test and the lack of ARCH effect in the residuals from ARCH test results. The U.S. data also support the normality of the residuals. More importantly, however, the tests for model adequacy are captured by the von-Neumann ratio and the recursive t-test. As pointed out in Harvey (1990, page 157), the von-Neumann test provides the most appropriate basis for a general test of misspecification with recursive residuals. In this context the dynamic linear models for the bubble and the no-bubble solutions both perform extremely well.

Figure 2 plots the bubble price ratio for the sample period and the substantial variation of the bubble component is visible for all the countries. Except for the U.S., there is evidence of negative bubble for the other three countries in the initial part of the sample period. Each country was affected differently by the oil price shock of the 1970s. The most severe impact appears to have occurred in the U.K. The fall in the bubble percentage during the October 1987 stock market crash is evident for all countries. It is also worth observing that there is a general upward trend for the bubble price ratio toward the later part of sample period for Germany, U.K. and the U.S. but not for Japan. Figure 3 depicts the contrasting views of the fundamental prices as computed by the model against the observed prices. This provides the visual evidence of the collapsing and self-starting nature of the stochastic bubble we have attempted to capture in this study.

In order to quantify the performance improvement of the bubble solution compared to the no-bubble case with GARCH (1,1) errors, we present in Table 8 the in sample fitting statistics, RMSE and MAE. The entries in Table 8 confirm that the bubble solution does a credible job in terms of both metrics. For example, the bubble solution for the U.S. monthly data reduces the metric RMSE to 7% and the metric MAE to 52% of the no-bubble solution respectively.

We indicated earlier the importance and the extent of investigation into the study of market linkages by various researchers. In this paper we are able to focus on this aspect in two different levels. The study of stochastic bubbles through the dynamic linear models enables us to decompose the price into a fundamental and a bubble component. It is, therefore, natural to examine whether the market linkages exist via both these components. McCarthy and Najand (1995) demonstrated the influence of the U.S. market on several other OECD countries using daily data which might have unintended consequences of trading time overlap in these markets. Using monthly data over a period of 48 years we are in a better position to analyze the market interrelationships.

VAR methodology is often employed to study causal relationships. If some variables are not Granger-causal for the others, then zero coefficients are obtained. Besides, the information in the data may not be sufficient to provide precise estimates of the coefficients. In this context the top-down strategy of the subset VAR approach described in the earlier section is most suitable. For the fundamental price series we adopt this approach in the levels of the variables since these are all found to be stationary. Using the Hannan-Quinn criterion we start our VAR model with a lag of one and follow the subset analysis process described before. This gives us the model presented in Table 9. As with McCarthy and Najand (1995) we find strong evidence of the U.S. dominance on all the other three countries, but no reverse causality. This is a particularly important finding in the sense that this causality exists in the fundamental components of the prices. Intuitively, this evidence suggests that the US economy, as represented by the stock market data, acts as the engine of global growth. For Germany and Japan the causality from the U.S. is significant at the 5% level whereas for the U.K it is significant at the 1% level only. The overall significance of this modeling approach is also established by testing the multivariate version of the portmanteau test to detect whiteness of the residuals.

We also apply the top-down strategy for the subset VAR approach to the bubble components to examine the causality between the four markets. Since the bubble components are found to be non-stationary (results for the unit root tests not included) we model this using the first difference of the log prices. With the non-stationary bubble price series it is natural to expect some long-term equilibrium relationship between these variables. We detected one cointegrating vector using Johansen's procedure and this has been described in Table 10. We follow the same procedure (as for the fundamental prices) to obtain the subset VAR model, including the cointegrating vector that describes the causal relationship between these markets. Table 10 shows that causality exists from the U.S. to the other three markets. Also, these linkages are significant at the 5% level for Germany and Japan and only at the 1% level for the U.K. Similar to the fundamental prices there is no reverse causality in the bubble price components as well. It is also observed that the strength of this causality from the U.S. to Japan is slightly stronger for the bubble price process, 0.1915 as opposed to 0.1878 for the fundamental prices.

It is also noted from Table 10 that the coefficients of the error correction term i.e. 'Coint (-1)' are statistically significant. This implies that the modeled variables i.e. the changes in log prices, adjust to departures from the equilibrium relationship. The magnitude of the coefficient 'Coint (-1)' for the Japanese log price difference is much higher than the others, capturing, first the upward and later, the downward trend in the Japanese market. Although, the existence of an error correction model implies some form of forecasting ability, we do not pursue this in this paper. Finally, we note the multivariate portmanteau test for whiteness of residuals in Table 10. This again supports the model adequacy and hence the inferences drawn are statistically meaningful.

10. CONCLUSION

Economists have long conjectured that movements in stock prices may involve speculative bubbles because trading often generates over-priced or under-priced markets. A speculative bubble is usually defined as the difference between the market value of a security and its fundamental value. Although there are several important theoretical issues surrounding the topic of asset bubbles, the existence of bubbles is inherently an empirical issue that has not been settled.

This paper reviews several important tests and offers a new methodology that improves upon the existing ones. In particular, we use the unobserved component methodology also used by other authors but our implementation of the state space form is different. We treat both the dividend process and the bubble process as part of the state vector in a dynamic linear model that allows for a straightforward comparison with the no bubble solution. The new methodology is applied to the four mature markets of the U.S., Japan, England and Germany to test whether a bubble was present during the period of January 1951 to December 1998. To establish the soundness of our methodology, we have also reproduced the empirical results of other authors using annual U.S. data covering the long period 1871-1998.

Once we find evidence of bubbles in these four mature stock markets, we next ask the question whether these bubbles are interrelated. We avoid using the technical term of contagion because it has a very specific meaning. Several authors use contagion to mean a significant increase in cross-market linkages, usually after a major shock. For example, when the Thai economy experienced a major devaluation of its currency during the summer of 1997, the spreading of the crisis across several Asian countries has been viewed as a contagion. Unlike the short-term cross-market linkages that emerge as a result of a major, often regional economic shock, we are here interested in long-run linkages. Bubbles often take long time, that is several years to inflate and one is interested in knowing if such processes travel from one mature economy to another. The bursting of a bubble, as in the case of the Thai market with its impact on the Asian stock markets, can be viewed as a contagion. However, our methodology captures long-term characteristics describing the markets studied over the entire sample period. Our statistical tests of the long-term linkages between the four mature stock markets provide evidence that U.S. bubbles cause bubbles in the other three markets but we find no evidence for reverse causality.

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Appendix A ESTIMATING THE PARAMETERS OF THE DLM

In this appendix we describe briefly how the unknown parameters in the DLM may be estimated. Our aim is to present an overview of the filtering and smoothing algorithm (known as Kalman filter and smoother) and the optimization of the likelihood function. Before proceeding, however, it is advantageous to express the DLM in term of suitable notations. Since the discussion here is applicable to both the bubble solution and the no-bubble solution described earlier we will not make any distinction between the two once the DLM have been defined.

We consider the DLM with reference to the following state and measurement equations:

$$y_t = \Gamma y_{t-1} + w_t, \text{ (State equation)} \tag{A.1}$$

$$z_t = A_t y_t + v_t, \text{ (Measurement equation).} \tag{A.2}$$

In this DLM, y_t is a $p \times 1$ vector of unobserved state variables, Γ is the $p \times p$ state transition matrix governing the evolution of the state vector. w_t is the $p \times 1$ vector of independently and identically distributed, zero-mean normal vector with covariance matrix Q. The state process is assumed to have started with the initial value given by the vector, y_0, taken from normally distributed variables with mean vector μ_0 and the $p \times p$ covariance matrix, Σ_0.

The state vector itself is not observed but some transformation of these is observed but in a linearly added noisy environment. In this sense, the $q \times 1$ vector z_t is observed through the $q \times p$ measurement matrix A_t together with the $q \times 1$ Gaussian white noise v_t, with the covariance matrix, R. We also assume that the two noise sources in the state and the measurement equations are uncorrelated.

The next step is to make use of the Gaussian assumptions and produce estimates of the underlying unobserved state vector given the measurements up to a particular point in time. In other words, we would like to find out, $E\left(y_t \mid \{z_{t-1}, z_{t-2} \cdots z_1\}\right)$ and the covariance matrix,

$P_{t|t-1} = E\left[\left(y_t - y_{t|t-1}\right)\left(y_t - y_{t|t-1}\right)'\right]$. This is achieved by using Kalman filter and the basic system of equations is described below.

Given the initial conditions $y_{0|0} = \mu_0$, and $P_{0|0} = \Sigma_0$, for observations made at time 1,2,3…T,

$$y_{t|t-1} = \Gamma y_{t-1|t-1}, \tag{A.3}$$

$$P_{t|t-1} = \Gamma P_{t-1|t-1}\Gamma' + Q, \tag{A.4}$$

$$y_{t|t} = y_{t|t-1} + K_t\left(z_t - A_t z_{t|t-1}\right), \text{ where the Kalman gain matrix} \tag{A.5}$$

$$K_t = P_{t|t-1}A_t'\left[A_t P_{t|t-1}A' + R\right]^{-1}, \tag{A.6}$$

and the covariance matrix $P_{t|t}$ after the t^{th} measurement has been made is,

$$P_{t|t} = \left[I - K_t A_t\right]P_{t|t-1}. \tag{A.7}$$

Equation (A.3) forecasts the state vector for the next period given the current state vector. Using this one step ahead forecast of the state vector it is possible to define the innovation vector as,

$$\nu_t = z_t - A_t y_{t|t-1} \tag{A.8}$$

and its covariance as,

$$\Sigma_t = A_t P_{t|t-1}A_t' + R. \tag{A.9}$$

Since in finance and economic applications all the observations are available, it is possible to improve the estimates of state vector based upon the whole sample. This is referred to as Kalman smoother and it starts with initial conditions at the last measurement point ie $y_{T|T}$ and $P_{T|T}$. The following set of equations describes the smoother algorithm:

$$y_{t-1|T} = y_{t-1|t-1} + J_{t-1}\left(y_{t|T} - y_{t|t-1}\right), \tag{A.10}$$

$$P_{t-1|T} = P_{t-1|t-1} + J_{t-1}\left(P_{t|T} - P_{t|t-1}\right)J_{t-1}', \text{ where} \tag{A.11}$$

$$J_{t-1} = P_{t-1|t-1}\Gamma'\left[P_{t|t-1}\right]^{-1}. \tag{A.12}$$

It should be clear from the above that to implement the smoothing algorithm the quantities $y_{t|t}$ and $P_{t|t}$ generated during the filter pass must be stored.

With reference to the DLM for the bubble and the no-bubble solutions it is obvious that the parameters of interest are embedded in the matrices Γ and Q, since by construction of our models $R \equiv 0$. The description of the above filtering and the smoothing algorithms assumes that these parameters are known. In fact, we want to determine these parameters and this achieved by maximizing the innovation form of the likelihood function. The one step ahead innovation and its covariance matrix are defined by the equations (A.8 and A.9) and since these are assumed to be independent and conditionally Gaussian, the log likelihood function (without the constant term) is given by,

$$\log(L) = -\sum_{t=1}^{T}\log\left|\Sigma_t(\Theta)\right| - \sum_{t=1}^{T}\nu_t'(\Theta)\Sigma_t^{-1}(\Theta)\nu_t(\Theta). \tag{A.13}$$

In this expression Θ is specifically used to emphasize the dependence of the log likelihood function on the parameters of the model. Once the function is maximized with respect to the parameters of the model, the next step of smoothing can start using those estimated parameters.

Maximization of the function in (A.13) may be achieved using one of two approaches. The first one depends on algorithm like Newton-Raphson and the second one is known as the EM (Expectation Maximization) algorithm. In this paper we employ Newton-Raphson technique to achieve our objective and since the likelihood function is reasonably well behaved, maximization is achieved quite quickly. In some modelling situations it may not be so straightforward. EM algorithm has been reported to be quite stable in the presence of bad starting values, although it may take longer to converge. Some researchers report that when good starting values are hard to obtain, a combination of the two approaches may be useful. In that situation it is preferable to employ EM algorithm first in order to obtain an intermediate estimates and then switch to Newton-Raphson method. Interested readers may refer to Shumway and Stoffer (2000, p. 323).

Table 1
Parameter Estimates from the State Space Model Using Kalman Filter
Bubble Solution: USA Yearly Data

ψ	σ_η	ϕ_1	ϕ_2	σ_δ
0.9830*	0.1857*	0.1929*	-0.2016*	0.1289*
(0.0203)	(0.0110)	(0.0711)	(0.0656)	(0.0049)

Estimates reported here are obtained from maximising the innovation form of the likelihood function. Numerical optimisation procedure in GAUSS is used without any parameter restriction. The standard errors (reported below the parameters in parentheses) are obtained from the Hessian matrix at the point of convergence. These estimates are robust to different starting values including different specification of the prior covariance matrix. Significance at 5% level is indicated by *.

Table 2
Parameter Estimates from the State Space Model Using Kalman Filter
No Bubble Solution with GARCH (1,1) Error for Price Equation: USA Yearly Data

ψ	ϕ_1	ϕ_2	σ_δ	α_0	α_1	β_1
0.6195*	0.2170*	-0.2164*	0.1290*	0.0073	0.1754	0.6054*
(0.1327)	(0.0867)	(0.0856)	(0.0081)	(0.0062)	(0.1279)	(0.2649)

Estimates reported here are obtained from maximising the innovation form of the likelihood function. Numerical optimisation procedure in GAUSS is used without any parameter restriction. The standard errors (reported below the parameters in parentheses) are obtained from the Hessian matrix at the point of convergence. These estimates are robust to different starting values including different specification of the prior covariance matrix. GARCH (1,1) error for state space system implemented following Harvey, Ruiz, Sentana (1992). Significance at 5% level is indicated by *.

Table 3

Residual Diagnostics and Model Adequacy Tests: USA Yearly Data

	Portmanteau	ARCH	KS Test	MNR	Recursive T
Bubble	0.035	0.385	0.138	0.531	0.952
No Bubble	0.033	0.519	0.119	0.422	0.931

Entries are p-values for the respective statistics except for the KS statistic. These diagnostics are computed from the recursive residual of the measurement equation, which corresponds to the real dividend process. The null hypothesis in portmanteau test is that the residuals are serially uncorrelated. The ARCH test checks for no serial correlations in the squared residual up to lag 26. Both these test are applicable to recursive residuals as explained in Wells (1996, page 27). MNR is the modified Von Neumann ratio test using recursive residual for model adequacy (see Harvey (1990, chapter 5). Similarly, if the model is correctly specified then Recursive T has a Student's t-distribution (see Harvey (1990, page 157). KS statistic represents the Kolmogorov-Smirnov test statistic for normality. 95% and 99% significance levels in this test are 0.121 and 0.145 respectively. When KS statistic is less than 0.121 or 0.145 the null hypothesis of normality cannot be rejected at the indicated level of significance.

Table 4

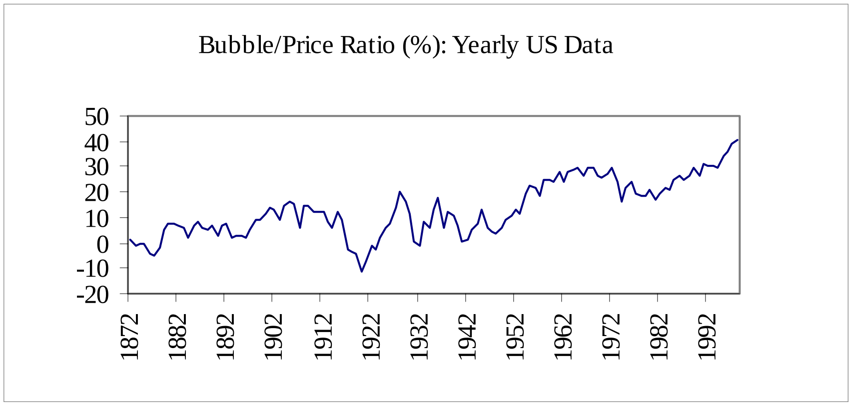
Bubble Solution Versus No Bubble with GARCH Error Compared

USA Yearly Data

	RMSE	MAE
Bubble	0.25	0.34
No Bubble GARCH (1,1)	1.37	1.42

RMSE and MAE stand for 'root mean squared error' and 'mean absolute error' respectively. These are computed from the differences between the actual log prices and the fitted log prices from the corresponding estimated model. Additional details are in the text.

Figure 1
Plots using the Smoothed Estimates of the Bubble from the State Space Model
USA Yearly Data



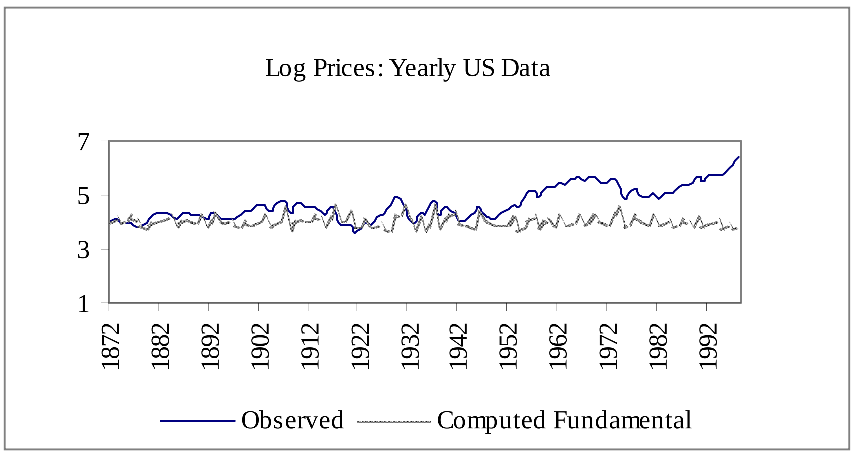


Table 5
Parameter Estimates from the State Space Model Using Kalman Filter
Bubble Solution: Monthly Data

	ψ	σ_η	ϕ_1	ϕ_2	ϕ_3	σ_δ
Germany	0.9980*	0.0470*	-0.0009	0.0611*	0.0947*	0.0475*
	(0.0011)	(0.0010)	(0.0400)	(0.0210)	(0.0271)	(0.0002)
Japan	0.9989*	0.0570*	-0.0879*			0.0511*
	(0.0010)	(0.0013)	(0.0370)			(0.0007)
UK	0.9983*	0.0535*	-0.5210*	-0.3669*	-0.1324*	0.0407*
	(0.0047)	(0.0009)	(0.0144)	(0.0214)	(0.0225)	(0.0003)
USA	0.9964*	0.0416*	-0.7218*	-0.3553*	-0.0969*	0.0287*
	(0.0020)	(0.0009)	(0.0350)	(0.0453)	(0.0387)	(0.0007)

Estimates reported here are obtained from maximising the innovation form of the likelihood function. Numerical optimisation procedure in GAUSS is used without any parameter restriction. The standard errors (reported below the parameters in parentheses) are obtained from the Hessian matrix at the point of convergence. These estimates are robust to different starting values including different specification of the prior covariance matrix. Significance at 5% level is indicated by *.

Table 6
Parameter Estimates from the State Space Model Using Kalman Filter
No Bubble Solution with GARCH (1,1) Error for Price Equation: Monthly Data

	ψ	ϕ_1	ϕ_2	ϕ_3	σ_δ	α_0	α_1	β_1
Germany	0.8526*	0.0047	0.0631	0.0848*	0.0475*	0.0001*	0.1108*	0.8633*
	(0.0391)	(0.0407)	(0.0409)	(0.0415)	(0.0014)	(5.14e-5)	(0.0299)	(0.0341)
Japan	0.5437*	-0.0906*			0.0511*	0.0000	0.0988*	0.8869*
	(0.0372)	(0.0407)			(0.0015)	(0.0000)	(0.0232)	(0.0301)
UK	0.2830*	-0.5331*	-0.3425*	-0.1148*	0.0407*	0.0004*	0.2307*	0.6107*
	(0.0380)	(0.0411)	(0.0440)	(0.0399)	(0.0012)	(0.0001)	(0.0541)	(0.0910)
USA	0.3189*	-0.7213*	-0.3271*	-0.0901*	0.0288*	0.0001*	0.0657*	0.8365*
	(0.0344)	(0.0413)	(0.0484)	(0.0400)	(0.0008)	(4.62e-5)	(0.0274)	(0.0533)

Estimates reported here are obtained from maximising the innovation form of the likelihood function. Numerical optimisation procedure in GAUSS is used without any parameter restriction. The standard errors (reported below the parameters in parentheses) are obtained from the Hessian matrix at the point of convergence. These estimates are robust to different starting values including different specification of the prior covariance matrix. GARCH (1,1) error for state space system implemented following Harvey, Ruiz, Sentana (1992). Significance at 5% level is indicated by *.

Table 7
Residual Diagnostics and Model Adequacy Tests: Monthly Data

	Portmanteau	ARCH	KS Test	MNR	Recursive T
Bubble					
Germany	0.253	0.158	0.176	0.586	0.903
Japan	0.061	0.206	0.093	0.379	0.972
UK	0.366	0.199	0.136	0.467	0.931
USA	0.377	0.327	0.048	0.425	0.894
No Bubble					
Germany	0.254	0.195	0.175	0.466	0.806
Japan	0.017	0.194	0.089	0.186	0.771
UK	0.307	0.179	0.139	0.571	0.907
USA	0.353	0.283	0.047	0.418	0.846

Entries are p-values for the respective statistics except for the KS statistic. These diagnostics are computed from the recursive residual of the measurement equation, which corresponds to the real dividend process. The null hypothesis in portmanteau test is that the residuals are serially uncorrelated. The ARCH test checks for no serial correlations in the squared residual up to lag 26. Both these test are applicable to recursive residuals as explained in Wells (1996, page 27). MNR is the modified Von Neumann ratio test using recursive residual for model adequacy (see Harvey (1990, chapter 5). Similarly, if the model is correctly specified then Recursive T has a Student's t-distribution (see Harvey (1990, page 157). KS statistic represents the Kolmogorov-Smirnov test statistic for normality. 95% and 99% significance levels in this test are 0.057 and 0.068 respectively. When KS statistic is less than 0.057 or 0.068 the null hypothesis of normality cannot be rejected at the indicated level of significance.

Table 8
Bubble Solution Versus No Bubble with GARCH Error Compared
Monthly Data

	RMSE	MAE
Bubble		
Germany	0.796	0.795
Japan	1.730	1.730
UK	0.247	0.366
USA	0.117	0.895
No Bubble GARCH (1,1)		
Germany	2.945	2.945
Japan	4.394	4.395
UK	0.719	0.838
USA	1.734	1.735

RMSE and MAE stand for 'root mean squared error' and 'mean absolute error' respectively. These are computed from the differences between the actual log prices and the fitted log prices from the corresponding estimated model. Additional details are in the text.

Figure 2
Plots using the Smoothed Estimates of the Bubble from the State Space Model
Monthly Data for Germany, Japan, UK, USA

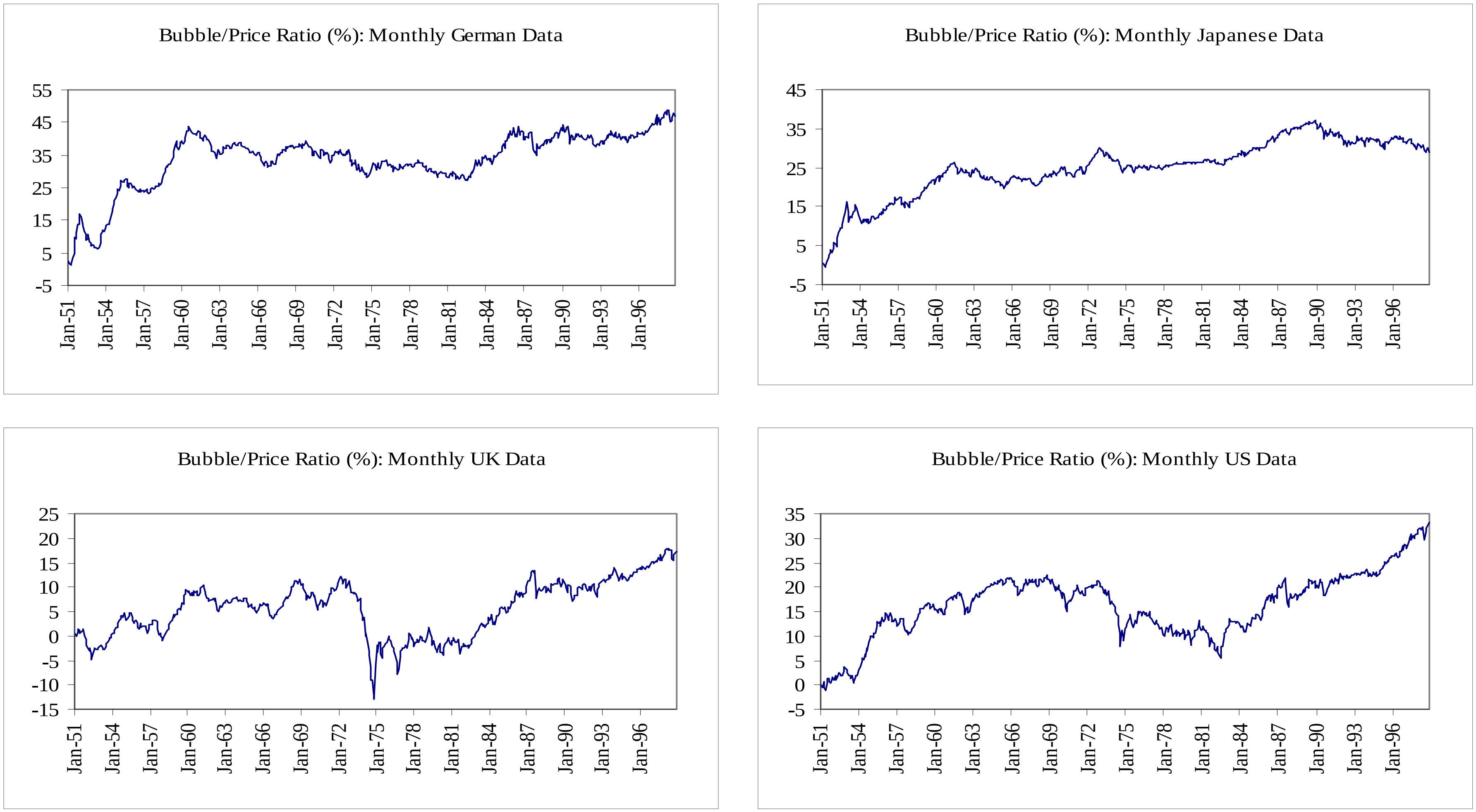
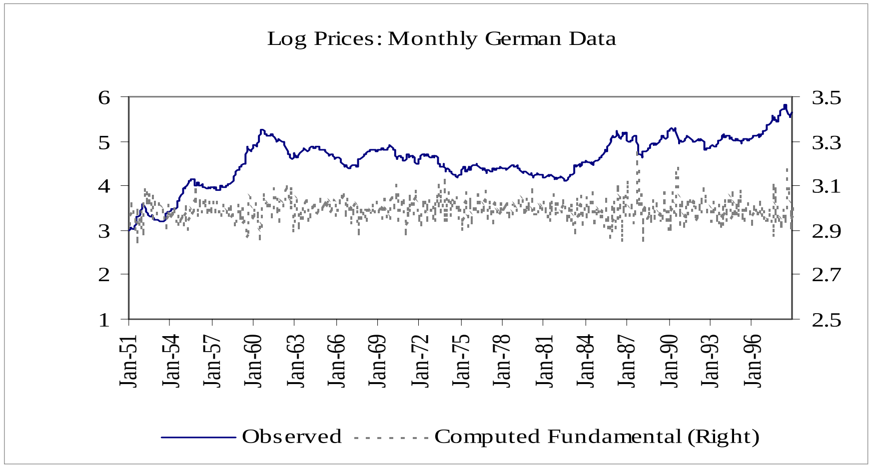
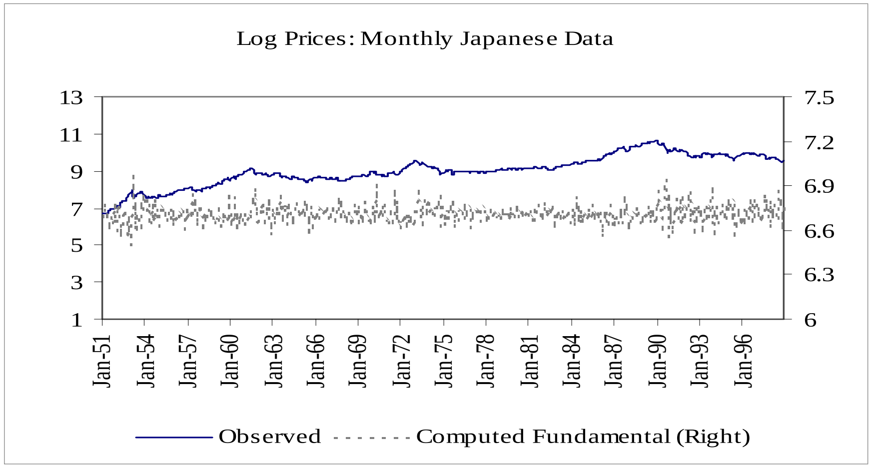
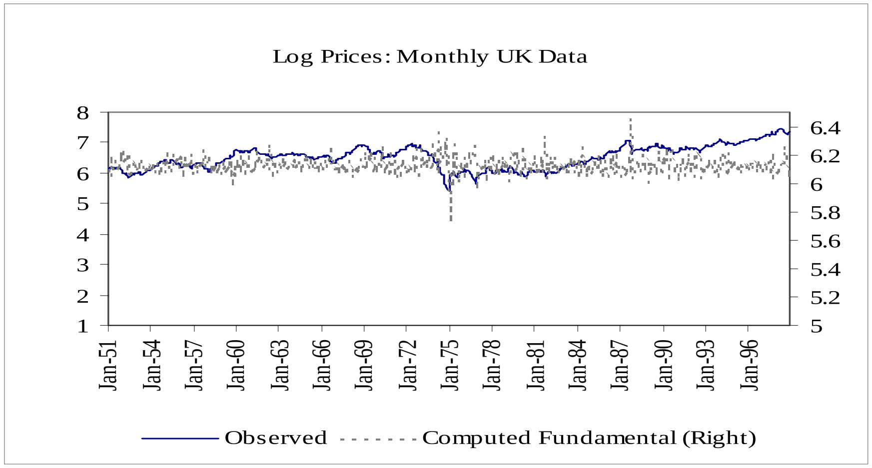


Figure 3
Plots using the Smoothed Estimates of the Bubble from the State Space Model
Monthly Data for Germany, Japan, UK, USA







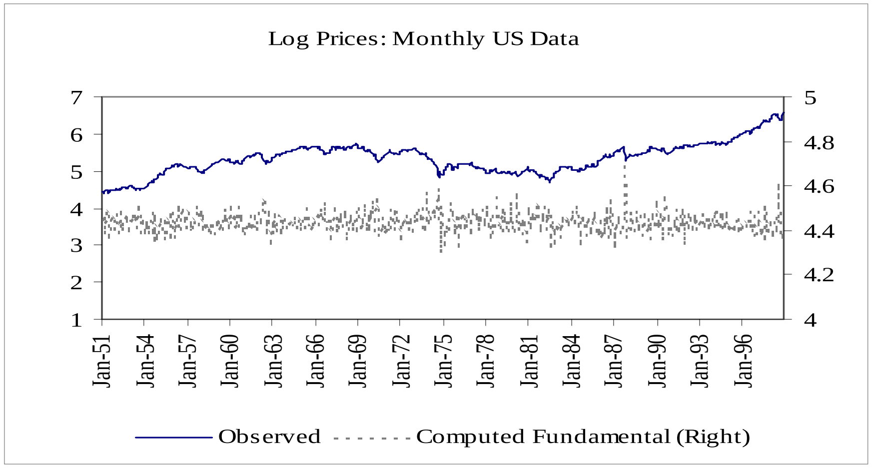


Table 9

Subset VAR Estimation Results for Linkages Between Markets in Fundamental Prices

	GR (-1)	JP (-1)	UK (-1)	US (-1)	Constant
GR	0.2074*			0.1904*	1.7063*
	(3.40)			(3.89)	(8.23)
JP	-0.1029			0.1878*	6.1837*
	(-1.91)			(3.08)	(23.95)
UK			0.0939*	0.1078**	5.0729*
			(1.97)	(1.76)	(18.02)
US					4.4358*
					(25.50)

Details of the methodology for determining the subset VAR relations are given in the text. This has been done in the level variables since the fundamental price series are stationary. The numbers in parentheses are t-statistics for the corresponding coefficient. Significance at 5% and 10% level are indicated by * and ** respectively. The p-value for the multivariate portmanteau statistic for residual white noise is 0.017. This is described in Lutkepohl (1993) page 188. This indicates that the model adequately represents the relationship documented here.

Table 10

Subset VAR Estimation Results for Linkages Between Markets in Bubble Prices

	ΔGR (-1)	ΔJP (-1)	ΔUK (-1)	ΔUS (-1)	Coint (-1)	Constant
ΔGR	0.1289*			0.1904*	0.0071*	0.0033
	(2.94)			(3.91)	(2.47)	(1.74)
ΔJP	-0.1436*			0.1915*	0.0167*	0.0048*
	(-2.67)			(3.20)	(4.76)	(2.09)
ΔUK			0.0956*	0.1064**		0.0016
			(1.99)	(1.73)		(0.74)
ΔUS					0.0009*	0.0038*
					(3.57)	(2.21)

The bubble prices are found non-stationary and Johansen's procedure identified existence of one cointegrating vector. The lagged value of this cointegrating vector (COINT) has been used in estimating the subset VAR relations for the linkages between the markets. The details of the unit root and the cointegration tests are not reported here but can be obtained from the authors. The estimated cointegrating vector (normalized on GR) including TREND and constant terms is given below. The numbers in parentheses are t-statistics for the corresponding coefficient. Significance at 5% and 10% level are indicated by * and ** respectively.

GR (-1) - 1.5826 JP (-1) + 2.7303 UK (-1) - 3.2545 US (-1) + 0.0054 TREND + 2.3772

The p-value for the multivariate portmanteau statistic for residual white noise is 0.068. This is described in Lutkepohl (1993) page 188. This indicates that the model adequately represents the relationship documented here.