

NONLINEAR BUBBLES IN CHINESE STOCK MARKETS IN THE 1990s

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Abstract:

A time series of the Shanghai stock index in China for the 1990s is studied for the possible existence of nonlinear speculative bubbles. Three alternative specifications of fundamentals are estimated using VAR models of domestic and international variables. These are subjected to regime switching tests and rescaled range analysis tests. Nulls of no persistence were mostly rejected, suggesting the strong possibility of bubbles. Nonlinearities beyond ARCH effects using the BDS test could not be rejected. The paper also discusses the special circumstances of the stock market in an emerging transition economy.

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INTRODUCTION

The economy of the Peoples' Republic of China (PRC) was the most rapidly growing in the world during the decade of the 1990s. This has resulted in a substantial improvement of the material standard of living for most of the Chinese people. It also continued a process of systemic transition that began with the reforms initiated by the late Deng Xiaoping in 1978, possibly the most successful such transition from a largely planned socialist economy to a more market-oriented economy in the world [Rosser and Rosser, 2004, Ch. 15]. Furthermore, as the world's largest nation in population with by some measures its second largest economy, the economic development of China is of immense global significance quite aside from the special characteristics that it exhibits.

As a part of the process of developing its markets and growing and opening to the world economy, China established a stock market in its traditional pre-socialist financial center of Shanghai in December, 1990, as well as another in the Special Enterprise Zone (SEZ) of Shenzhen, adjacent to Hong Kong, in April, 1991 [Mookerjee and Yu, 1995]. Thus, it has joined the company of nations with "emerging stock markets," albeit one with the special character of emerging from a socialist system as well as from a condition of low per capita output. These markets have continued to be hindered by various institutional constraints, including serious limitations of direct participation by foreigners¹ as well as the substantial presence of state-owned enterprises (SOEs) in the markets [Wong, 1993] with the accompanying suspicion by many observers of interventions and manipulations by the government of the markets through its widespread

ownership of stock in the SOEs listed on the markets. Nevertheless, they have seen rapid expansion and growth by all measures during the decade, an unsurprising outcome given the rapid growth of the economy and the interest of the government in subjecting the SOEs to the discipline of market forces. This growth for the Shanghai and Shenzhen markets is shown in Table 1 for 1991-99. Figure 1 shows a time series of the main index on the Shanghai market for this period. Macroeconomic and more general data for the PRC is shown in Table 2 for the same period.² Figure 2 shows a time series plot of Industrial Production Index and Retail Price Index. Both of these variables are used in estimating Models II and III.

Like many other emerging markets during this period the Chinese markets appear to have exhibited considerable volatility [Bailey, 1994]. Mookerjee and Yu [1999] have studied the efficiency of the Chinese markets during the early 1990s and found them to fail the efficient market hypothesis on various grounds, including the presence of both seasonal and weekend anomalies, although no January effect. They suggest that concern over the volatility of the markets has led the government to hold back from establishing another one in Tianjin, although there may be other reasons for this reluctance. Although China avoided the more serious oscillations and repercussions from the East Asian financial crisis of 1997-98, these events have understandably heightened concerns about the functioning of these markets and their volatility, even with the limited participation by foreign investors.

This paper extends these earlier studies both by extending the data set to the end of the 1990s as well as by investigating whether or not the volatility of these markets represents the emergence and collapse of speculative bubbles and the nature of those bubbles. Even prior to the dramatic events of 1997, the Southeast Asian markets showed evidence of mean reversion [Malliaropulos and Priestley, 1999]. Likewise, for most of the Pacific-Rim markets considerable evidence has been adduced in favor of the hypothesis that speculative bubbles of a nonlinear character beyond ARCH effects existed in the period prior to the dramatic events of 1997 [Ahmed, Rosser, and Uppal, 1999] with similar evidence present for Pakistan as well [Ahmed and Rosser, 1995; Ahmed, Rosser, and Uppal, 1996]. This paper will generally follow the approach used in these latter papers to study the behavior of the Chinese stock markets.

The basic method is to estimate hypothetical fundamentals models for the stock markets via vector autoregressive (VAR) models that link domestic and global variables with stock market indices, following the approach of Canova and Ito [1991]. Residuals are then estimated for these models from the relevant series, which are then further analyzed. Two tests are used to test the null of hypothesis of no bubbles. If there are no bubbles, these residuals should oscillate approximately randomly as white noise without exhibiting any sustained up or down trends.

One test is a regime switching test due to Hamilton [1989], an approach also used by van Norden and Schaller [1993]. The no bubbles hypothesis is viewed as a null of no trends which can be tested for using a Wald test, as bubbles

would presumably show up as trends in the residuals away from the fundamental in one direction or another (bubbles can be both positive and negative in direction). We also use Hurst coefficient methodology of rescaled range analysis (RSS), which was initially developed for detecting persistence of annual flooding patterns on the Nile River [Hurst, 1951], with short-term persistence of residuals away from fundamentals again being viewed as prima facie evidence of bubbles. Finally, the residual series has autoregressive conditional heteroskedasticity (ARCH) effects removed and the remaining series is tested for the existence of continuing nonlinearities using the BDS statistic (Brock, Dechert, LeBaron, and Scheinkman, 1997). Our empirical results are consistent with rejecting all nulls of no trends or persistence and also of rejecting the null of no nonlinearities beyond ARCH effects. Results using the rescaled range analysis are more mixed, but with most showing significant Hurst coefficients indicating positive persistence effects. Thus, our empirical results are mostly consistent with the possible existence of nonlinear speculative bubbles in Chinese stock markets.

We note one major caveat at this point regarding these results. This is the famous "misspecified fundamentals" problem, first noted by Flood and Garber [1980]. They argued that it is impossible to properly test for bubbles because what the econometrician observes may not reflect what was in the minds of investors in the market. More precisely, the econometrician may be observing part of a rationally expected skewed distribution whose tail does not appear in the sample that is observed. It is precisely because of this problem that we estimate three different fundamentals models. We reach the same conclusions

regarding for all three models for the regime switching and BDS tests, but get more mixed results on the other models for the rescaled range analysis, with the model using only domestic Chinese variables fully supporting the nonlinear bubbles hypothesis. But we recognize that for this data set there is no ultimate way to get around this problem.³ Thus, our results must be viewed as provisional.

THEORY OF BUBBLES

A speculative bubble exists for a series if $B_t > 0$ for an observable period for

$$P_t = F_t + B_t + \varepsilon_t \quad (1)$$

where P is the price of the asset, F is the fundamental value of the asset, B is the bubble component of the price of the asset, ε is determined by a stochastic process, and all variables are indexed for time t . We have deliberately not specified a functional form for the bubble process here and note that it could involve a variety of possible forms, including some that might involve interactions with the value of the fundamental, although in our empirical testing it will be assumed to exhibit some kind of time trend. We also do not specify here the nature of the residual stochastic process, although later we shall use the null of an assumed i.i.d. pattern for testing for nonlinearity of the process.

Clearly the difficult issue here is to identify the fundamental, F , the point made from an econometric perspective by Flood and Garber [1980]. One definition of a fundamental is a competitive general equilibrium value for the asset. A

complication can arise here if there are multiple such equilibria. In such a case, what appears to be a bubble dynamic might really be just a movement from one equilibrium to another. More commonly for an asset earning a real return, the fundamental is identified with the present value of a rationally expected future stream of real returns, that is the sum of those returns properly discounted.

Of course we can easily see how the misspecified fundamental problem can arise as there is no way to know for sure what the rationally expected stream of future returns is for most assets is. There is also the more serious issue of how to measure the fundamental for a nonfinancial asset, although in some cases it may be possible to use financial data as a proxy as proposed by Cardell, Kling, and Petry [1995] in studying the market for postage stamps.

Within the context of models with fully informed and rational agents, the existence of bubbles is strictly limited because of the "backward induction" problem that it will be irrational to hold the asset just prior to the bursting of the bubble and therefore it will be irrational to hold it any earlier time, thereby ruling out the bubble ever beginning. Tirole [1982] provided a formal development of this argument with the conclusion that bubbles cannot occur in models with discrete time, a finite number of fully-informed, risk-averse, infinitely-lived agents, trading a finite number of real assets with real returns. In some cases, the kinds of bubbles that can arise as these conditions are variously relaxed may be of the essentially unobservable stationary variety, as for example in overlapping generations models which allow for an infinite number of finitely-lived agents who keep passing the bubble forward to the next generation [Tirole, 1985]. The

bubble must be stationary because otherwise it would explode to infinity thus violating transversality conditions. Indeed, Samuelson [1958] used the overlapping generations model to show the possible existence of fiat money, which can be viewed as having a fundamental of zero and thus whose stationary value represents a kind of bubble.

Allowing for an infinite number of commodities allows for certain kinds of stationary "charges" to exist above and beyond identifiable fundamentals. These charges can be interpreted as being bubbles, albeit again of the stationary variety [Gilles and LeRoy, 1992; Magill and Quinzii, 1996].

Another relaxation is to allow for continuous time trading. Faust [1989] shows that this can also allow for bubbles essentially by undoing the backward induction logic. There is no "period before the bubble crashes" as there are no discrete periods and thus no backward induction on such a sequence of discrete periods.

Another approach that has been widely used for empirical testing is that of stochastically crashing rational bubbles, in which agents are imperfectly informed regarding the date of the crash of the bubble, although they know the probability of a crash at any point in time [Blanchard and Watson, 1982]. Such models lead to bubbles that rise more rapidly than the rate of interest so as to provide a risk premium to the rational investors for the probability of the crash. Among those using such a methodology to search for bubbles have been Frankel [1985] who found a bubble in the U.S. dollar but one that failed to rise sufficiently rapidly unless one assumed a declining probability of a crash over time, Dwyer and Hafer [1990] who came to similar conclusions as Frankel for a set of stock

markets in seven countries, Elwood, Ahmed, and Rosser [1999] who found such a possible bubble in the yen-Deutschemark exchange rate in 1989-90, and Johansen, Sornette, and Ledoit [1999] who found bubbles using a log-periodic method based on this model for eight stock and foreign exchange markets.⁴

Discussion of the possible emergence of bubbles has increasingly focused on models with heterogeneous agents in which they may vary according to their degrees of information or according to their strategies, with some assumption of limited information on all or some of their parts. Black [1986] initiated serious discussion of such models by introducing the concept of noise traders, although this idea has a long history in the discussion of speculative bubbles, going all the way back to such historical episodes as the Mississippi and South Sea bubbles of the early 1700s [Kindleberger, 2001; Rosser, 2000]. Observers of those antique events long ago noted the contrast between smarter insiders who got in and out early and made money and less savvy outsiders who got in late and failed to get out before the crash, thereby losing money, in some cases lots of it, the proverbial suckers who ended up with the "hot potato" but who are ruled out of existence in models of fully informed rational agents.

A wide variety of specifications have been made for such models of heterogeneity that can lead to bubbles emerging. Allen and Gorton [1993] posit asymmetric information with fully rational agents. Bhattacharya and Lipman [1995] posit differences in initial wealth where wealth is private information. Shiller [1984] posits outright irrationality in the form of fad behavior. Herding behavior as a response to limited information is shown by Lux [1995]. More

generally, Thaler [1991] labels models with heterogeneous agents of varying degrees of information and rationality as *quasi-rational*.

Models with many categories of interacting agents who learn to use evolving strategies have been studied that show bubbles emerging [Arthur, Holland, LeBaron, Palmer, and Tayler, 1997; Lux, 1998; Bullard and Duffey, 2001]. DeLong, Shleifer, Summers, and Waldmann [1990] allow for some risk-loving agents who nevertheless make money and survive. In a later model, they [1991] show such a formulation allowing for a period of declining prices after a peak before there is a crash, a pattern that Kindleberger [2001] argues is typical of most actual historical bubble candidates and which is studied further by Rosser [1997].

As we shall not estimate a particular theoretical model for generating bubbles, other than that there should be an identifiable and statistically significant motion away from the estimated fundamental, we shall not present any specific theoretical model in detail. However, we note the model of Day and Huang [1990], which allowed for three types of agents, rational fundamentalists, trend-chasing "sheep," and market-making specialists. Their model allows for a variety of dynamics including positive and negative bubbles and even chaotic dynamics under certain circumstances (the trend-chasing sheep must be very active). Gu [1993] has empirically fit this model to asset market data. Another model that reaches similar conclusions is that of De Grauwe, Dewachter, and Embrechts [1993] for foreign exchange rate dynamics, with just two categories of agents.

Unsurprisingly, chaotic dynamics can arise in a variety of the more complex models with many categories of learning and interacting agents [Lux, 1998].

Estimating Fundamentals

Ideally, the fundamental of a financial asset is the sum of discounted expected real returns. However, one cannot assume that ex post observed returns equal ex ante expected ones, nor can one assume that ex post observed market interest rates equal ex ante subjective discount rates used by rational investors. Thus, Shiller [1981] initiated the study of excess volatility in stock markets by comparing stock price variability with stock dividend variability, finding the former to be significantly greater than the latter. However, others such as Marsh and Merton [1986] argued that there were explanations for this such as a tendency for corporate managers to smooth dividends. Much debate has surrounded this topic as reviewed by LeRoy [1989]. However, as we lack dividend data we are unable to use them to attempt to estimate fundamentals.

Furthermore, although we possess data on market interest rates, in China these are strictly controlled by the central bank and vary very infrequently compared to most other economies. However, some have argued that real discount rates may vary because of time-varying risk premia, with Domowitz and Hakkio [1985] suggesting the use of ARCH effects models for getting at these in the case of foreign exchange rate models. However, Froot and Frankel [1989] used survey data to show that most investors in foreign exchange markets do not have risk premia that vary substantially with time, although this result does not

necessarily carry over to investors in Chinese stock markets. Nevertheless, we do not use interest rates or risk premia estimates to model fundamentals.

Following Canova and Ito [1991] and Ahmed, Rosser, and Uppal [1996, 1999], we seek to model the fundamental by estimating a VAR on certain other variables that should relate to the profitability of the stock market. In doing so we seek a daily series of the fundamental that is presumed to vary with new information. We estimate three alternative fundamentals models, one based on global variables, one based on domestic variables, and one combining the two. These are shown below. The method for estimating the VARs follows Ahmed, Rosser, and Sheehan [1988, 1989] and Rosser and Sheehan [1995] and is of the first differences of the natural logarithms of the respective variables.

Model I: (Shanghai stock index, Fareast stock index, World stock index, U.S. stock index)

Model II: (Shanghai stock index, Chinese industrial production index, Chinese retail price index)

Model III: (Shanghai stock index, Fareast stock index, World stock index, U.S. stock index, Chinese industrial production index, Chinese retail price index)

The latter two variables in Model II were interpolated to provide a daily series. We considered using the exchange rate as well, but as with interest rates, this is fixed against the dollar most of the time and rarely varies. For all three models a residual series is constructed that is studied in the tests below. Footnote 4

provides a brief description of the data and data sources⁵. We fully recognize that there are other approaches and methods for making these estimates.

REGIME SWITCHING TESTS

Hamilton [1989] introduced an approach to regime switching tests that can be used to test for trends in time series and switches in trends, as used in Engel and Hamilton [1990] and van Norden and Schaller [1993]. We use this approach as our main test for the null of no bubbles on the residual series derived above which is given by

$$\varepsilon_t = n_t + Z_t \quad (2)$$

where

$$n_t = \mu_1 + \mu_2 s_t \quad (3)$$

and

$$Z_t - Z_{t-1} = N_1(Z_{t-1} - Z_{t-2}) + \dots + N_r(Z_{t-r} - Z_{t-r-1}) + \gamma \quad (4)$$

with $s = 1$ being a positive trend, $s = 0$ being a negative trend, and $\mu_1 \neq 0$ indicating the possible existence of a trend element beyond the VAR process.

Furthermore, let

$$\text{Prob}[s_t = 1 \mid s_{t-1} = 1] = p, \text{Prob}[s_t = 0 \mid s_{t-1} = 1] = 1 - p \quad (5)$$

$$\text{Prob}[s_t = 0 \mid s_{t-1} = 0] = q, \text{Prob}[s_t = 1 \mid s_{t-1} = 0] = 1 - q. \quad (6)$$

Following Engel and Hamilton [1990] a "no bubbles" test proposes a null hypothesis of no trends given by $p = 1 - q$. This is tested by with a Wald test statistic given by

$$[p - (1 - q)]/[var(p) + var(1 - q) + covar(p, 1 - q)]. \quad (7)$$

Results of this test are shown in Table 3 for Model I, in Table 4 for Model II, and in Table 5 for Model III. The critical value for rejecting the null of no trends is $\chi^2 = 3.8$. Clearly the null is strongly rejected in all cases, with the respective values being 47.6 for Model I, 40.98 for Model II, and 62.77 for Model III.

HURST PERSISTENCE TESTS

Hurst [1951] developed a test to study persistence of Nile River annual flows which was first applied to economic data by Mandelbrot [1972]. For a series x_t with n observations, mean of x^*m and a max and a min value, the range $R(n)$ is

$$R(n) = [\max_{1 \leq k \leq n} \sum_{j=1}^k (x_j - x^*) - \min_{1 \leq k \leq n} \sum_{j=1}^k (x_j - x^*)]. \quad (8)$$

The scale factor, $S(n, q)$ is the square root of a consistent estimator for spectral density at frequency zero, with $q < n$,

$$S(n, q)^2 = g_0 + 2 \sum_{j=1}^q w_j(q) g_j, \quad w_j(q) = 1 - [j/(q-1)], \quad (9)$$

with g 's autocovariances and w 's weights based on the truncation parameter, q , which is a period of short-term dependence.⁶ The classical Hurst case has $q = 0$, which reduces the scaling factor to a simple standard deviation.

Feller [1951] showed that if x_t is a Gaussian i.i.d. series then

$$R(n)/S(n) \propto n^H, \quad (10)$$

with $H = 1/2$, which implies integer integrodifferentiation and thus standard Brownian motion, the "random walk." H is the Hurst coefficient which can vary from zero to one with a value of $1/2$ implying no persistence in a process, a value significantly less than $1/2$ implying "anti-persistence" and a value significantly greater than $1/2$ implying positive persistence. The significance test involves breaking the sample in two parts and then estimating a Chow test on the null that the subperiods possess identical slopes. This technique is also called *rescaled range analysis*.

Table 6 presents the results of this test for Model I. H is estimated to equal 0.58. Chow tests of the significance of this coefficient provide mixed results. For a test of a model with both slope and intercept the computed F -value is 61.90, substantially above the 3.15 cutoff for a significant rejection of the null hypothesis that the coefficient is equal to 0.50 (thus indicating no persistence). However, for a test of a model with the intercept suppressed, the F -value is only 1.94, falling below the level of significance for rejection.

Table 7 presents the results of this test for Model II. H is estimated to equal 0.57. Chow tests of this find a significant rejection of the null hypothesis for both

the case of slope and intercept, with an *F-value* of 59.7, and for the case with the intercept suppressed, with an *F-value* of 4.17.

Table 8 presents the results of this test for Model III, which qualitatively resemble the outcome for Model I. The estimated value of *H* is 0.56. The Chow test with both slope and intercept rejects the null with a computed *F-value* of 11.51, while that suppressing the intercept produces an *F* of 0.5968 that fails to reject the null hypothesis.

The model (II) with only the domestic variables more strongly exhibits persistence than do either of the models with the international stock market index variables. Curiously, the regime switching tests show a stronger rejection of the null hypothesis of no trends for the complete model with both domestic and international variables (III), while for the Hurst rescaled range analysis this model shows the weakest evidence of persistence. We have no ready explanation for this apparent discrepancy between these two tests.

NONLINEARITY TESTS

We test for nonlinearity of the VAR residual series in two stages. The first is to remove ARCH effects. Engle [1982] the nonlinear variance dependence measure of autoregressive conditional heteroskedasticity (ARCH) as

$$x_t = \delta_t \mu_t \quad (11)$$

$$\delta_t^2 = \alpha_0 + \sum_{i=0}^n \alpha_i x_{t-i}^2 \quad (12)$$

with μ_t i.i.d. and the α_i 's different lags. We use a three period lag and, as expected, found significant ARCH effects in all series, available on request from the authors.⁷

The second stage involves removing variability attributable to the estimated ARCH effects from the VAR residual series for both models. The remaining residual series is run through the BDS test due to Brock, Dechert, LeBaron, and Scheinkman [1996], with useful guidance on certain aspects in Brock, Hsieh, and LeBaron [1991]. This statistic tests for generalized nonlinear structure but does not test for any specific form such as alternative ARCH forms or chaos.

The correlation integral for a data series $x_t, t = 1, \dots, T$ results from forming m -histories such that $x = [x_t, x_{t+1}, \dots, x_{t+m+1}]$ for any embedding dimension m . It is

$$c_m T(\varepsilon) = \sum_{t < s} I_\varepsilon(x_t^m, x_s^m) [2/T_m(T_m - 1)] \quad (13)$$

with a tolerance distance of ε , conventionally measured by the standard deviation divided by the spread of the data, $I_\varepsilon(x_t^m, x_s^m)$ is an indicator function equaling 1 if $\|x_t^m - x_s^m\| < \varepsilon$ and equaling zero otherwise, and $T_m = T - (m - 1)$.

The BDS statistic comes from the correlation integral as

$$BDS(m, \varepsilon) = T^{1/2} \{c_m(\varepsilon) - [c_1(\varepsilon)]^m\} / b_m \quad (14)$$

where b_m is the standard deviation of the BDS statistic dependent on the embedding dimension m . The null hypothesis is that the series is i.i.d., meaning that for a given ε and an $m > 1$, $c_m(\varepsilon) - [c_1(\varepsilon)]^m$ equals zero. Thus, sufficiently large values of the BDS statistic indicate nonlinear structure in the remaining series. This test is subject to severe small sample bias with a cutoff of 500

observations sufficient to overcome this, a minimum both of our daily series easily achieve.

Tables 9 to 11 present the results of this test for Models I to II for embedding dimensions, $m = 2$ to 4 ($m = 3$ is conventional). The critical value for rejecting the null of i.i.d. is approximately 6. With estimated BDS statistics ranging from 11.26 to 25.21 for the various models and m 's the null is rejected for all cases. Thus, there appears to be remaining nonlinearity beyond basic ARCH in the VAR residual series.

Of course, just as our earlier tests are subject to the validity of our original VAR specifications, likewise so is this test. We also emphasize that the nature of the remaining nonlinearity remains unknown.

CONCLUSIONS

This paper has examined the behavior of the Shanghai stock index during the 1990s since this stock market was established, the largest stock market in China. Three alternative VAR models were estimated, one with global variables, one with domestic variables, and one combining both, in first differenced natural logarithmic form, as possible proxies for the fundamental series of the stock market, recognizing that these are all seriously subject to the misspecified fundamentals problem. These residuals were then subjected to tests for the existence of trends using regime switching and rescaled range methods. The null of no trends were rejected for all cases using the regime switching test. They were also rejected for tests of persistence using the rescaled range

analysis (Hurst coefficient) except for the models with international variables when the Chow tests suppress the slope (the null is rejected in all cases when the intercept is included). That the model combining domestic and international variables (III) more strongly shows trends for the regime switching test but shows less evidence of persistence for the rescaled range analysis reminds us that we still face the misspecified fundamentals problem.

Furthermore, the series were examined for nonlinearity beyond ARCH effects by removing variability attributable to ARCH and then testing the remaining series for remaining nonlinearity using the BDS statistic. The null hypothesis that the remaining series is i.i.d. was rejected for all models for various embedding dimensions. These tests are mostly consistent with the possible existence of speculative bubbles with complex nonlinear components in the Shanghai stock market, a conclusion that is consistent with the perceptions of many market participants.

At this point we must ask what the significance of this finding is. At one level it is unsurprising, given the evidence that many emerging stock markets and also those in the Pacific basin apparently also exhibit similar evidence. Even if what is being observed are not actually speculative bubbles, it is clear that the Shanghai market exhibits considerable volatility, so considerable that as noted above the Chinese government has held back from establishing another such market in Tianjin. Thus, this volatility and possible presence of bubbles is slowing the movement towards a more normal financial markets system as exists in most countries and more broadly the market reform movement.

However, the Chinese markets have special features reflecting the nature of the Chinese economic system in its intermediate form between socialism and capitalism.⁸ Not only have foreigners been forbidden until recently from participating in the markets, but most of the firms on the markets are state owned. Indeed, one motive for establishing the markets was to possibly establish a mechanism for encouraging the restructuring of the SOEs as they face market valuation. But this becomes difficult in an environment of great volatility. Furthermore, anecdotal evidence suggests that it is widely believed that the Chinese government actively intervenes in the markets in various ways. It is widely thought that these interventions are aimed at curbing speculation and stabilizing the markets. But it is very unclear whether or not they succeed, and it may be that these interventions actually exacerbate the tendencies to high volatility.

In short, how to deal with its nascent stock markets is a serious policy issue for China that is intimately and deeply connected with the broader policy issues related to its general reform efforts. The government faces many questions regarding whether to expand the markets, to intervene or not to intervene, as well as the question of whether and how to allow foreign participation.

One thing that the Chinese markets almost certainly share with other emerging markets is involvement in a learning process. It may be that with time and development the Chinese stock markets will not exhibit so much volatility. But, given that even markets in the advanced economies appear to be susceptible to speculative bubbles and crashes, no one should expect that these

phenomena will disappear from the Chinese markets either.

NOTES

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1. The markets consist of A and B stocks. Foreigners are only allowed to purchase B stocks, which are a minority of those issued. Kim and Singal [2000] argue that opening stock markets to foreigners does not necessarily lead to increased volatility if accompanied by a well-structured plan of general market liberalization.
2. For a discussion of the evolution of Chinese monetary policy and its relation to the macroeconomy, see Li [1996]. We focus on the Shanghai stock market as the larger of the two, although their behavior has been fairly similar.
3. There may be some data sets for which this problem may be resolved. Ahmed, Koppl, Rosser, and White [1997] propose that the rapid emergence and then disappearance of premia on closed-end funds may be such a case, as the net asset value of such a fund can be viewed as a reasonable approximation of the true fundamental, taking into account certain transactions costs. These authors present evidence for bubbles in many closed-end funds in 1989-90.

4. Sornette [2003] presents the argument that stochastically crashing rational bubble will hit an infinite rate of growth in a finite time, which will be the expected time for the crash of the bubble. The methodology used by him and his coauthors draws on *econophysics* approaches that emphasize models that can explain the presence of power law distributions in stock return data series. Others pursuing related approaches include Lux and Marchesi [1999] and Farmer and Joshi [2002].
5. Daily data for the S&P 500 Stock Index, World Stock Index, and Fareast Stock Index came from Datastream International and cover the period from December 12, 1990 to October 10, 1999. Data for the Shanghai Stock Index came from Zhongyin Network and covers the same period. The data for Chinese Industrial Production and the Retail Price Index came from the Almanac of China's Economy [1999] and covers the time period from January 1992 to October 1999. The original series was monthly, which was used to generate daily data. We performed Augmented Dickey-Fuller tests on the series, which found stationarity for logarithmic first differences of the stock market series, results available from authors.
6. Lo [1991] has criticized the use of the classical Hurst coefficient for studying long-term persistence in stock markets precisely because of this presence of short-term dependence for which he proposes a method of avoiding. However, this is not a problem for us because it is precisely short-term dependence that we are interested in detecting.

7. In addition to basic ARCH models we also estimated simple generalized autoregressive conditional heteroskedastic (GARCH) models [Bollerslev, 1986] for these residuals series. The results were essentially very similar to those for the simple ARCH models and are available from the authors on request.
8. The question arises to what extent the Chinese stock market may behave like those in other emerging transitional stock markets. Certainly the experience of the Russian market and its well-publicized crash in 1998 suggests some possible similarities [Sornette, 2003, Ch. 8], as well as considerable volatility being exhibited in other similar markets. However, it would appear that during the phase of actual transition the differences among these markets may be as important as their similarities, with very different rules and institutional arrangements holding, differences that may become reduced as time goes by and their market participants become more experienced. Evidence that stock market dynamics in some of these countries may be quite different from those in other markets has been argued by Gnaciński and Makowiec [2003] whose modeling of the Polish stock market suggests that it has behaved in a quite unique manner.

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Table 1

Chinese Stock Market

Year	Listed companies	Shares	Value	cash flow	Investors
		(billion)	(billion Yuan)	(billion Yuan)	(million)
1991	14	0.63	10.9	5.2	0.37
1992	53	7.32	104.8	68.1	2.16
1993	183	32.87	353.1	366.7	7.78
1994	291	63.97	369.1	812.8	10.59
1995	323	76.60	347.4	439.6	12.42
1996	530	111.04	984.2	2133.2	23.07
1997	745	177.12	1752.9	3072.0	33.32
1998	851	252.68	1950.6	2354.4	39.11
1999	947	290.89	2647.0	3132.0	44.8

1US\$=8.3RMB Yuan

Figure 1

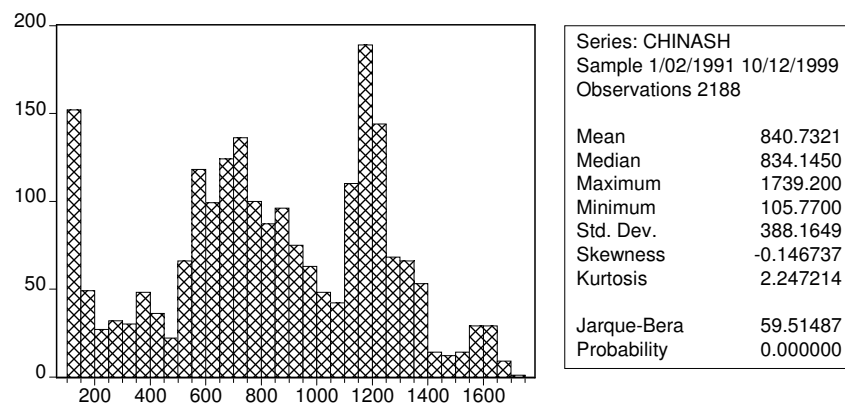
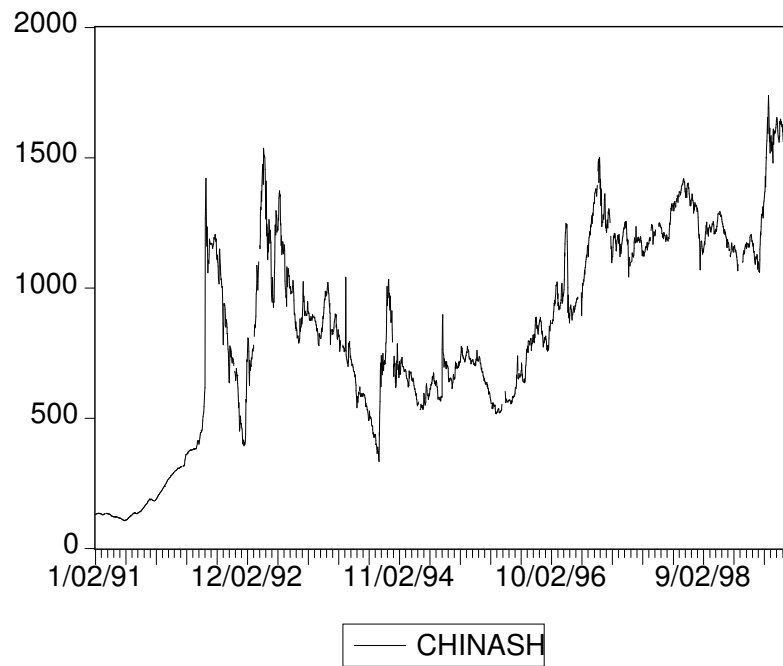


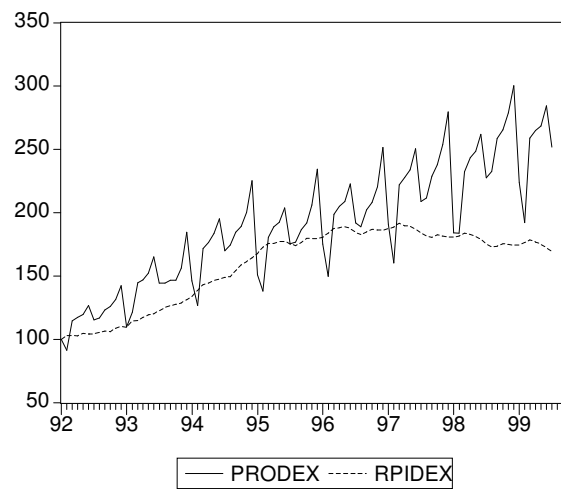
Table 2

Chinese Macroeconomic Data

Year	GDP (billion RMB)	Growth rate (%)	Inflation (%)	interest rate (annual,%)	Exchange rate (RMB/US\$)
1990	1855	3.8	2.1	8.64	4.78
1991	2162	9.2	2.9	7.56	5.32
1992	2664	14.2	5.4	7.56	5.51
1993	3463	13.5	13.2	10.98	5.76
1994	4676	12.6	21.7	10.98	8.62
1995	5848	10.5	14.8	10.98	8.35
1996	6789	9.6	6.1	7.47	8.31
1997	7477	8.8	0.8	5.67	8.29
1998	7955	7.8	-2.6	3.78	8.28
1999	8205	7.1	-3.0	2.25	8.28

Resources: Almanac of China's Economy, China Statistical Yearbook.

Figure 2



Period: 1992,1 To 1999,10
PRODEX: Industrial production index
RPIDEX: Retail price index
Source: Almanac of China's Economy

Table 3

Wald Test Results on Residuals from Four-Variable VAR
Model of Shanghai Stock Index, Fareast Stock Index,
World Stock Index and U.S. Stock Index (S&P 500)

Sample Period	$H_0: P_1=1-P_2 \chi^2(1)$
December 12, 1990-October 10, 1999	47.6

Critical Value $\chi^2(1)=3.8$

Table 4

Wald Test Results on Residuals from Three-Variable VAR
Model of Shanghai Stock Index, Industrial Production Index and
Retail Price Index

Sample Period	$H_0: P_1=1-P_2 \chi^2(1)$
January 2, 1992-June 30, 1999	40.98

Critical Value $\chi^2(1)=3.8$

Table 5

Wald Test Results on Residuals from Six-Variable VAR
Model of Shanghai Stock Index, Fareast Stock Index,
World Stock Index , U.S. Stock Index (S&P 500), Industrial Production Index and
Retail Price Index

Sample Period	$H_0: P_1=1-P_2\chi^2(1)$
January 2, 1992-June 30, 1999	62.77

Critical Value $\chi^2(1)=3.8$

Table 6

Hurst Coefficients and Related Chow Tests

<p>Hurst Coefficients and Chow Test Results on Residuals from Four-Variable VAR Model of Shanghai Stock Index, Fareast Stock Index, World Stock Index and U.S. Stock Index (S&P 500)</p>	<p>Hurst Coefficient =0.58 Chow Test for Slope and Intercept: Critical F, (.05)_{2, 73}= 3.07-3.15 Computed F= 61.90 We reject the null hypothesis that Pre and Post bubble periods are the same. Chow Test for Slope Alone: Critical F, 0.05(1, 73) =3.92-4.00 Computed F= 1.94 We do not reject the null hypothesis that both slopes are the same.</p>
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Table 7

Hurst Coefficients and Related Chow Tests

<p>Hurst Coefficients and Chow Test Results on Residuals from Three-Variable VAR Model of Shanghai Stock Index, Industrial Production Index and Retail Price Index</p>	<p>Hurst Coefficient= 0.57 Chow Test for Slope and Intercept: Critical F, 0.05 (2, 67)= 3.07-3.15 Computed F= 59.7 We reject the null hypothesis that pre and post bubbles periods are the same. Chow Test for Slope Alone: Critical F, 0.05 (1, 67)= 3.92-4.00 Computed F= 4.17 Reject the null hypothesis that slopes are the same.</p>
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Table 8

Hurst Coefficients and Related Chow Tests

<p>Hurst Coefficients and Chow Test Results on Residuals From Six-Variable VAR Model of Shanghai Stock Index, Fareast Stock Index, World Stock Index, U.S. Stock Index (S & P 500), Industrial Production and Retail Price Index</p>	<p>Hurst Coefficient = 0.56 Chow Test for Slope and Intercept: Critical F, 0.05 (1, 69)= 3.07-3.15 Computed F= 11.51 We reject the null hypothesis that pre and post bubble periods are the same. Chow Test for Slope Alone: Critical F, 0.05(2. 69)= 3.07-3.15 Computed F= 0.5968 Do not reject the null hypothesis that slopes are the same.</p>
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Table 9

BDS/SD Statistics Based on Residual From Four-Variable VAR Model
 Shanghai Stock Index, Fareast Stock Index,
 World Stock Index and U. S. Stock Index (S&P 500)

Embedding Dimensions (M)	T= No. Of Observations	BDS/SD Statistics
2	2184	18.37
3	2184	22.45
4	2184	25.51

Data use for BDS test are residuals form ARCH process conducted on residuals from VAR. For beginning and ending dates, please see table 2.

Table 10

BDS/SD Statistics Based on Residuals From Three-Variable VAR Model
Shanghai Stock Index, Industrial Production Index and Retail Price Index

Embedding Dimensions (M)	T= No. of Observations	BDS/SD Statistics
2	1866	11.26
3	1866	18.25
4	1866	22.17

Data used for BDS test are residuals from ARCH process conducted on residuals from VAR. For beginning and ending dates, please see tables 2 and 4.

Table 11

BDS/SD Statistics Based on Residuals From Six –Variable VAR Model
 Shanghai Stock Index, Fareast Stock Index,
 World Stock Index, U. S. Stock Index (S&P 500), Industrial Production Index and
 Retail Price Index

Embedding Dimensions (M)	T=No. of Observations	BDS/SD Statistics
2	1848	11.54
3	1848	11.89
4	1848	21.32

Data used for BDS test are residuals from ARCH process conducted on residuals from a Six-Variable VAR model. Sample set is the same as Table 5