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Are there bubbles in Chinese RMB–dollar exchange rate? Evidence from generalized sup ADF tests

Chun Jiang^a, Yi Wang^a, Tsangyao Chang^b and Chi-Wei Su^{c,*}

^a*Department of Finance, Wuhan University, Wuhan 430072, China*

^b*Department of Finance, Feng Chia University, Taichung, Taiwan*

^c*Department of Finance, Ocean University of China, Qingdao 266100, China*

This article uses recently developed generalized sup ADF (GSADF) unit root tests into the analysis of nominal RMB–dollar exchange rates bubbles. Based on the results from the GSADF tests, we find strong evidence of explosive behaviour in the nominal exchange rate and investigate two bubbles there. The first bubble is during 2005–2006 which is determined neither by the relative prices of traded goods nor the relative price of nontraded goods. The second bubble busts in 2008 during subprime crisis period, and which is determined by the relative prices of traded goods but not the relative price of nontraded goods. There is no bubble before 2005 as the exchange rate is under fixed regime. As for this result, some expansionary monetary and fiscal policies are required in China since these are the most efficient and effective under a bubble burst scenario.

Keywords: multiple bubbles; RMB–dollar exchange rate; GSADF test

JEL Classification: C12; C15; F31

I. Introduction

This article examines whether multiple bubbles exist in Chinese renminbi (RMB)–dollar exchange rate during the 1995–2013 period using a new test proposed by Phillips *et al.* (2013). Bubble is an observable economic phenomenon, the common definition of which is an economic cycle characterized by rapid expansion followed by a contraction. Brunnermeier (2009) defines that bubbles are typically associated with dramatic asset price increases followed by a collapse. Bubbles arise if the price exceeds the asset's

fundamental value. This can occur if investors hold the asset because they believe that they can sell it at an even higher price to some other investor even though the asset's price exceeds its fundamental value. Since asset prices affect the real allocation of an economy, it is important to understand the circumstances under which these prices can deviate from their fundamental value. Exchange rate that can be seen as an asset price (Obstfeld and Rogoff, 1996) also has bubbles. The most famous exchange rate bubbles are the Mexican Peso Crisis (1994–1995) when Mexican peso suddenly depreciated for 15%

*Corresponding author. E-mail: cwsu@ouc.edu.cn

and caused a chaos worldwide; and the Asia Crisis (1997–1998) when Thai baht, Korean won, Japanese yen, Hong Kong dollar, Singapore dollar, etc., all could not hold on and rushed down bold and flowing (Wikipedia). Bubbles can harm an economy or they could even benefit the economy (Jiménez, 2011). Every single bubble generates a redistribution of wealth, directly or indirectly, among the various agents in the economy. As bubble collapses will cause profound impact around the world, this intrigued economists and led to several strands of models, empirical tests and experimental studies to work for exchange rate bubbles (Tirole, 1982; Diba and Grossman, 1988b; Evans, 1991; Kindleberger, 2000; Thompson and Hickson, 2006; Brunnermeier, 2009)

Why we choose to study Chinese RMB–dollar exchange rate bubbles? First, the GDP of China has exceeded Japan and thereby became the second largest economy in 2010. Most literatures have pointed out that the relationship between GDP and exchange rate is positive. Balassa (1964) found the relationship between GDP and exchange rate is positive through a reappraisal of the PPP doctrine. Samuelson (1964), Hyder and Mahboob (2005) and Parveen *et al.* (2012) also found the relationship between GDP and exchange rate is positive. Maeso-Fernandez *et al.* (2001) found positive relationship between productivity differential and real exchange rate in Euro by using BEER/PEER approach. Faruqee (1995) and Rajan and Parulkar (2006) also found positive relationship between productivity differential and exchange rate. These literatures all prove that lasting increase in GDP will lead to overvalued exchange rate and thus lead to potential bubbles.

Second, the growing weight of China in the world economy, measured by GDP and quantum of trade, has intensified debate on the potential international role of its currency – the renminbi (RMB). Much progress has been made on RMB settlements for trade involving China and on RMB-denominated bond issuance in Hong Kong, China, but that RMB internationalization is still limited due to capital account controls. A high degree of RMB internationalization requires significant capital account liberalization – supported by financial market liberalization including market-determined interest rates, and by effective financial regulation and supervision – which in turn would call for greater exchange rate flexibility so that the

Chinese government can enjoy monetary policy autonomy. This shift also implies that the country will have to move towards a more market-determined exchange rate. A more flexible exchange rate is an important macroeconomic adjustment facilitator as well as shock absorber for countries with relatively open capital accounts. History has shown that open capital accounts and pegged exchange rates are a toxic mix. Country authorities experiencing substantial capital inflows and outflows need to be able to adjust domestic policy in response by allowing the exchange rate to float. Attempting to peg the exchange rate when the capital account is even moderately open constrains the ability to do so. In addition, exchange rate appreciation in response to balance of payments surpluses would raise the prices of locally produced goods and services relative to those of internationally traded goods – locally produced goods and services being where demand will grow most strongly as China rebalances its economy towards consumption. In this respect then, RMB internationalization should be thought of as integral to the larger ongoing process of rebalancing Chinese exchange rate flexibility.

Third, Johnson (1969) established monetarism exchange rate theory and stated that a country's exchange rate movements depend on three factors, one of which is the change of the expected inflation rate relative to foreign country expected inflation rate. Also from PPP theory, exchange rate is defined by the ratio of the price levels in the two countries. The relative PPP represents a period of exchange rate changes is affected partly because of the inflation rate. Zada (2010) found that higher inflation in local country leads to depreciation of local currency in Pakistan. Parveen *et al.* (2012) examined that high inflation also devalues the currency in Pakistan. Jamal (2005) examined the CPI ratio differential between US versus UK and US versus Japan, both showed negative relationship between exchange rate and CPI ratio, whereby high CPI will devalue the local currency. Ogun (2012) also found negative relationship between inflation rate differential and exchange rate in developing countries. As the CPI ratio in China is high above 4% in periods of year 1995M07–1997M03, 2007M06–2008M10 and 2010M10–2012M01, we are interested in examining whether in these periods there are exchange rate bubbles.

Fourth, several important exchange rate regime reforms had been conducted in China since July 2005 when exchange rate turned into managed floating rate, which implied the Chinese government's great efforts for the progress of exchange rate flexibility. China's new leadership has expressed a strong desire for exchange rate reform. China's November 2013 Third Plenum decision document underlines the goal to 'perfect the market-based renminbi exchange rate formation mechanism'. People's Bank of China Governor Zhou said that China will 'basically exit from normal foreign-exchange market intervention'. China has also made a number of bilateral and multilateral commitments to reform its exchange rate. At the July 2013 Strategic & Economic Dialogue in Washington, DC, China pledged to 'continue exchange rate reform, increase flexibility of the RMB exchange rate, and let the market play a more fundamental role in exchange rate formation'. Moreover, the RMB strengthened more significantly on a trade-weighted basis, but not as fast or by as much as is needed. By peeling back trading restrictions on the RMB, China is working towards its goal of having a currency that trades relatively freely. Overall, as the RMB exchange rate floating space is becoming bigger, whether there exist bubbles in the RMB exchange rate has gradually entered into policymakers' discussions.

This study represents our attempt to investigate whether there exist explosive bubbles in the Chinese RMB-dollar exchange rate using a new test proposed by Phillips *et al.* (2013). Our empirical investigation produces two noteworthy findings. First, the results indicate that there were explosive bubbles in the Chinese RMB-dollar exchange rate. In order to shed light on the causes of the explosiveness, we also test for explosive behaviour in the underlying fundamentals. Betts and Kehoe (2008) pointed out that movements in the nominal exchange rate are mainly driven by the relative prices of traded goods and one-third of nontraded goods. Following Betts and Kehoe (2008), we construct the relative prices of traded and nontraded goods as fundamentals for exchange rates. Results show that the bubble during 2005–2006 is determined neither by the relative prices of traded goods nor by the relative price of nontraded goods. The second bubble busted in 2008 during subprime crisis period is determined by the

relative prices of traded goods but not the relative price of nontraded goods. There is no bubble before 2005 when the exchange rate was under fixed regime. Our findings thus support the claims that the RMB-dollar exchange has been driven by rational bubbles.

This article is organized as follows. Section II summarizes the literature for our study. Section III briefly describes the theoretical background of bubbles. Section IV presents the data used in our study. Section V briefly describes the methodology proposed by Phillips *et al.* (2011a, 2011b, 2013). Section VI presents our empirical results and some policy implications. Section VII concludes the article.

II. Literature Review

Different series of theoretical set-ups for bubbles detection are presented in literature (Thompson and Hickson, 2006; Brunnermeier and Oehmke, 2012). Brunnermeier (2009) broadly divide the literature into four groups. There are four main strands of models that identify conditions under which bubbles can exist. The first class of models assumes that all investors have rational expectations and identical information (Flood and Garber, 1980; Blanchard and Watson, 1982; Tirole, 1982; West, 1987a; Diba and Grossman, 1988b; Evans, 1991; McKelvey and Palfrey, 1992; Kindleberger, 2000). These models generate the testable implication that bubbles have to follow an explosive path. In the second category of models, investors are asymmetrically informed and bubbles can emerge under more general conditions because their existence need not be commonly known (Tirole, 1982; Allen and Gorton, 1993; Allen *et al.*, 1993; Brunnermeier, 2001). A third strand of models focuses on the interaction between rational and behavioural traders. Bubbles can persist in these models since limits to arbitrage prevent rational investors from eradicating the price impact of behavioural traders (DeLong *et al.*, 1990; Shleifer and Vishny, 1997; Abreu and Brunnermeier, 2003; Brunnermeier and Nagel, 2004; Temin and Voth, 2004). In the final class of models, bubbles can emerge if investors hold heterogeneous beliefs, potentially due to psychological biases, and they agree to disagree about the fundamental value (Miller, 1977; Harrison and Kreps, 1978; Ofek and

Richardson, 2003; Scheinkman and Xiong, 2003). The first two groups of models analyse bubbles within the rational expectations paradigm, but differ in their assumption whether all investors have the same information or are asymmetrically informed. A third group of models focuses on the interaction between rational and nonrational (behavioural) investors. In the final group of models, traders' prior beliefs are heterogeneous, possibly due to psychological biases, and consequently they agree to disagree about the fundamental value of the asset. An extension of the bubble debate is whether they should be studied as rational or behavioural. Recent studies of exuberance in asset prices have suggested that the field of psychology should have an equal importance in explaining the detection and causes of bubbles (e.g. see Hirshleifer, 2001; Abreu and Brunnermeier, 2003; Vissing-Jorgensen, 2004). Despite this recent shift in focus to irrational and behavioural bubbles, some economists still believe rational bubbles are more plausible than their irrational counterparts (e.g. see LeRoy, 2004).

The occurrence of rational bubbles signifies that no long-run relationships exist between nominal exchange rate and prices. A vast amount of research has been devoted to investigating the presence of rational bubbles in exchange rate over the past three decades. For example, Evans (1986) finds significant evidence of bubbles in the Sterling–dollar exchange rate in the early 1980s, while Meese (1986), West (1987b) and Wu (1995) yield mixed results. Elwood *et al.* (1999) applied state–space models and Monte Carlo experiments to the Japanese and German exchange rate series and did find strong evidence of a deviation from white noise that is consistent with the existence of a stochastic rational bubble which burst between April and May of 1990, when it was a period of considerable turmoil in the Japanese and German financial markets. Maldonado *et al.* (2012) set-up three models to test the significance of speculative bubbles that may have occurred in the period that is considered. The results could not reject the null hypothesis that rational bubbles exist at the 5% significance level. Moreover, Bettendorf and Chen (2013) apply the sequential unit root tests and find strong evidence for explosive behaviour in the nominal Sterling–dollar exchange rates. While other literatures doubt on the claims that exchange rate has been driven by speculative bubbles. Jirasakuldech *et al.* (2006) investigate the presence of rational

speculative bubbles in the exchange rates of the British pound, the Canadian dollar, the Danish krone, the Japanese yen and the South African rand against the US dollar. The unit root test shows that the exchange rates and fundamental variables – money supply, income and interest rates – are integrated of order one, indicating no rational speculative bubbles while the cointegration test indicates evidence of a long-run relationship between the exchange rate series and the fundamental variables, which suggests that rational expectations bubbles do not affect these exchange rates.

There have been only a small number of studies covering Chinese exchange rate bubbles so far. The only literature studying China is Maldonado *et al.* (2014), who test the occurrence of rational bubbles in the exchange rate of Brazil, Russia, India, China and South Africa (the 'BRICS' countries group) against the US dollar and consider bubbles of the periodically recurring variety. They assume that the exchange rate is affected by the fundamental value following a modified PPP relation which takes into account interest rate differentials. The results for China were significant and may be interpreted as evidence that models for the fundamental rate that include asset markets (via the interest rates) are more adequate to model speculative bubbles than the ones which do not include them explicitly.

In terms of the literature, studies make decomposition of the nominal exchange rates by relative price of tradable goods and relative price of nontradable goods following Engel (1999). In this line of research, Betts and Kehoe (2001, 2006, 2008) extend Engel's original results by looking at a broad cross section of country pairs, consider a broader set of decompositions and propose a model with endogenous tradability of goods to account for the exchange rate dynamics between the US and Mexico. Burstein *et al.* (2006) study Engel's decomposition using import and export prices, and conclude that with such indices used as trade prices the verdict is more favourable for the traditional trade theory of real exchange rate determination. Mendoza (2000) study the bilateral real exchange rate between the US and Mexico across different nominal exchange rate regimes and find significant differences across regimes. Drozd and Nosal (2009) use value-added deflators and present the results that the relative contribution of nontradable goods in the data is higher – but still small in comparison to the model. Our work

is following Engel (1999) on the decomposition of the nominal exchange rates by tradable goods component and nontradable goods component. Relative to these papers, our contribution is to demonstrate how the standard international macro models fare in light of these findings.

As regards methodology, empirical studies have, for the most part, employed cointegration techniques to examine the relationship between exchange rate and prices. Among the most notable of these techniques is the widely employed Johansen cointegration test (Johansen, 1988; Johansen and Juselius, 1990), which is based on the linear autoregressive model and assumes that the underlying dynamics are in a linear form. From a theoretical perspective, there is no sound reason to assume that economic systems are intrinsically linear (see Barnett and Serletis, 2000). In fact, numerous studies have empirically demonstrated that financial time series, such as exchange rate, exhibit nonlinear dependencies (see Basci and Caner, 2005; Norman and Phillips, 2009; Lee and Chou, 2013). In addition, substantive evidence from the Monte Carlo simulations in Bierens (1997, 2004) has indicated that, inherent to the conventional Johansen cointegration framework, there is a misspecification problem when the true nature of the adjustment process is nonlinear and that the speed of adjustment varies with the magnitude of the disequilibrium. A recursive method, supremum ADF (SADF), has also been proposed by Phillips *et al.* (2011b) which can detect exuberance in asset price series during an inflationary phase. However, the Phillips *et al.* (2011b) recursive method is especially effective when there is a single bubble episode in the sample data as in the 1990s NASDAQ episode analysed in Phillips *et al.* (2011b) and in the 2000s US house price bubble analysed in Phillips and Yu (2011). Therefore, given the possibility of multiple bubbles within the same sample period, this study investigates whether multiple bubbles exist in the Chinese RMB–dollar exchange rate using a new test, generalized sup ADF (GSADF) test, proposed by Phillips *et al.* (2013). The major advantage of the approach is that it allows one to account for the nonlinear structure and break mechanisms while investigating the existence of multiple bubbles. We utilize the GSADF test to investigate the existence of multiple bubbles in the Chinese RMB–dollar exchange rate given that Chinese exchange rate markets are experiencing higher volatility after 1994 than before

when China had started a managed floating exchange rate system since 1994. It is often marked by frequent and erratic structural changes, which are usually driven by various policy events and events of global importance. To the best of our knowledge, this is the first study to test for bubbles in the Chinese RMB–dollar exchange rate with this method.

III. Theoretical Background

Following Bettendorf and Chen (2013), we assume the exchange rate is determined by current and expected values of fundamentals. We thus assume the following present value model of exchange rate in accordance with Engel and West (2005) and Leon-Ledesma and Mihailov (forthcoming):

$$s_t = (1 - \alpha) \sum_{j=0}^k \alpha^j E_t(f_{t+j}) + \alpha^{k+1} E_t(s_{t+k+1}) \quad (1)$$

where s_t is the nominal exchange rate, and f_t is the market fundamental at time period t . α denotes the discount factor. We assume that the exchange rate will only depend on future expected fundamentals in the long run, thus

$$\lim_{k \rightarrow \infty} \alpha^{k+1} E_t(s_{t+k+1}) = 0 \quad (2)$$

However, if the transversality condition does not hold, the exchange rate may be subject to an explosive rational bubble. Assuming that the bubble follows an AR(1) process, it can be written as

$$\eta_t = \frac{1}{\alpha} \eta_{t-1} + \varepsilon_t \quad (3)$$

where $\frac{1}{\alpha} > 1$, as the bubble is an explosive process. Errors are captured by $\varepsilon_t \sim NID(0, \sigma^2)$. Therefore, we can write the exchange rate as

$$s_t = s_t^f + \eta_t \quad \text{or} \quad s_t^f = s_t - \eta_t \quad (4)$$

where s_t^f denotes the discounted sum of all future economic fundamentals and η_t the bubble component. We assume that s_t^f is linearly dependent on the economic fundamental f_t . In line with Engel and West (2005), we also assume that f_t is I(1). According to the PPP model, the economic fundamental for the nominal exchange rate is the price differential

$$f_t = p_t - p_t^* \tag{5}$$

where p_t denotes the log level of the domestic price index. Asterisks denote foreign counterparts. We follow Engel (1999) to decompose the price index into indexes of nontraded and traded goods, and to consider a price index for a country as a weighted average of traded and nontraded goods prices

$$p_t = (1 - \beta)p_t^T + \beta p_t^N \tag{6}$$

p_t^T denotes the log of the traded goods price index, p_t^N the log of the nontraded goods price index and β the share of the nontraded goods component. For the foreign country, one can also write

$$p_t^* = (1 - \gamma)p_t^{T*} + \gamma p_t^{N*} \tag{7}$$

It follows that the price differential (f_t) can be decomposed into two components, the traded goods component (f_t^T) and the nontraded goods component (f_t^N).

$$p_t - p_t^* = (p_t^T - p_t^{T*}) + \beta(p_t^N - p_t^N) - \gamma(p_t^{N*} - p_t^{T*}) \tag{8}$$

The PPI is the most broadly available and frequently used index to represent the price level of traded goods. Though there are some producer goods that are not traded, PPI is measured at the production site and thus excludes marketing and other nontraded consumer services. Thus we construct the traded goods component using the PPI following Engel (1999):

$$f_t^T = p_t^T - p_t^{T*} = \ln(PPI_t) - \ln(PPI_t^*) \tag{9}$$

The relative nontraded goods component is constructed from the aggregate CPI relative to aggregate PPI:

$$f_t^N = \beta(p_t^N - p_t^T) - \gamma(p_t^{N*} - p_t^{T*}) = \ln(CPI_t) - \ln(PPI_t) - (\ln(CPI_t^*) - \ln(PPI_t^*)) \tag{10}$$

In the following section, we demonstrate how explosiveness can be detected in the nominal RMB–dollar exchange rate $-s_t$, and the ratio of the exchange rate relative to the two types of economic fundamentals, using recursive right-tailed unit root tests proposed by Phillips *et al.* (2011a), Phillips *et al.* (2011b) and Phillips *et al.* (2013), respectively.

IV. Data

In this article, we apply monthly nominal exchange rate, CPI and production price indexes from China and the US for our empirical study. Data are all from IMF International Financial Statistics. China devalued the yuan by 33% overnight to 8.7% to the dollar as part of reforms to embrace a ‘socialist market economy’ in January 1994, and fixed the yuan around 8.28 to the dollar from April 1994, which are milestones in the exchange rate regime reforms of China. After that, another important time is July 2005 when China revalued the yuan by 2.1% and revised rules governing its currency, and the exchange rate had shifted to ‘a managed floating exchange rate based on market supply and demand with reference to a basket of currencies’. From then on China has implemented several reforms to broaden the flexibility of RMB. Therefore, we chose data from 1995 M7 to ending period 2013 M10 due to data availability after 1994 reform, containing the period before and after the exchange rate reform in 2005. Thus we could find out whether there are explosive bubbles under the peg exchange rate regime or managed floating exchange rate regime. Following the study of Bettendorf and Chen (2013), we use nominal exchange rate (lex), nominal exchange rate – relative prices of traded goods ratio

(lex_ft) and nominal exchange rate – relative prices of nontraded goods ratio (lex_fnt) for our analysis. The software we use for our study is GAUSS 10.0.3. If we look at the top panel of Figs 1–3, we can see the plots of these three data series and there might be evidence of explosive bubbles in these three data

series. These results motivate us to empirically investigate whether there are explosive bubbles in the RMB–dollar exchange rate using both the SADF and GSADF tests proposed by Phillips *et al.* (2011a), Phillips *et al.* (2011b) and Phillips *et al.* (2013), respectively.

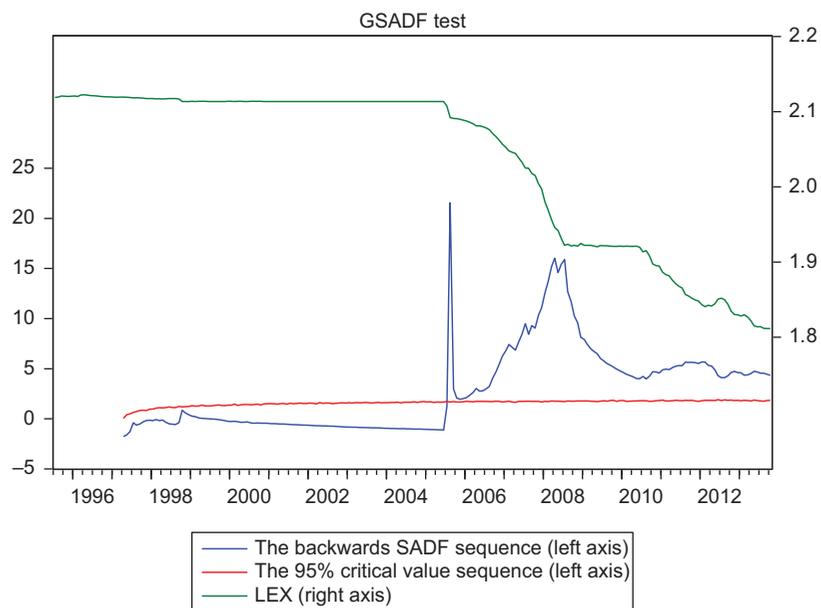


Fig. 1. GSADF test for nominal exchange rate (lex)

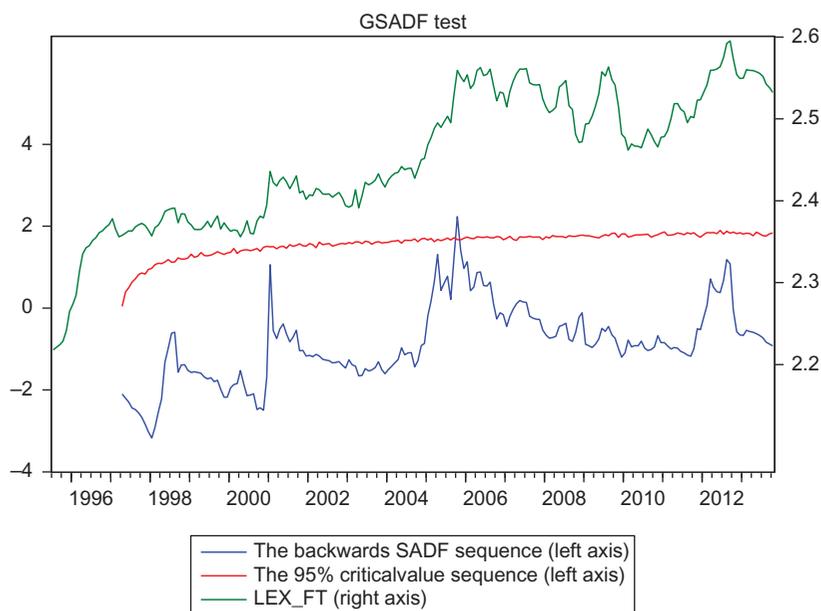


Fig. 2. GSADF test for exchange rate to relative price of traded goods ratio (lex_ft)

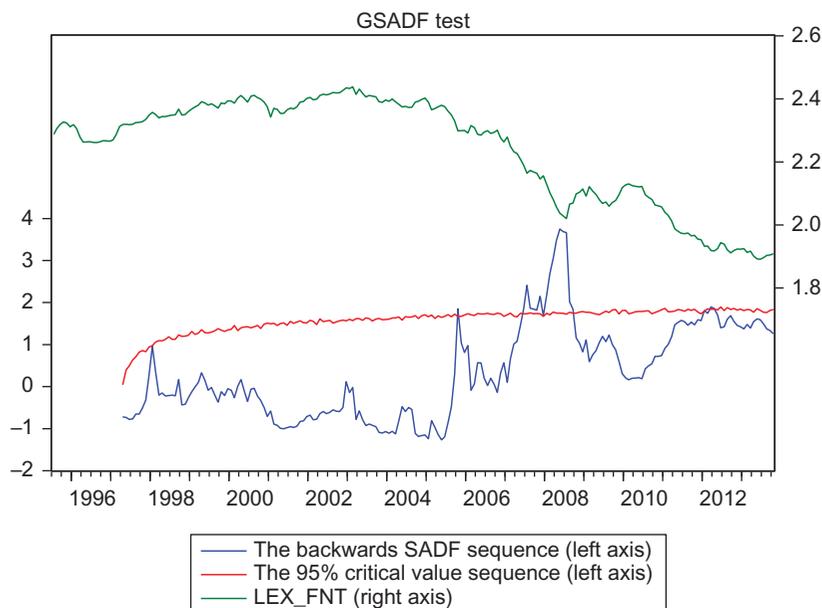


Fig. 3. GSADF test for exchange rate to relative price of nontraded goods ratio (lex_fnt)

V. Methodology¹

Over the past several decades, identifying a bubble in real time has proved to be challenging in the econometric literature. Econometric techniques suffered from finite sample bias. For example, conventional unit root and cointegration tests may be able to detect one-off exploding speculative bubbles but are unlikely to detect periodically collapsing bubbles. In other words, efforts to identify significant warning signs of future price bubbles have been impeded by the necessity to spot multiple starting and ending points. The reason is that conventional unit root tests are not well equipped to handle changes from I(0) to I(1) and back to I(0). This makes detection by cointegration techniques harder due to bias and kurtosis problem (Evans, 1991).

Recently, an innovative and persuasive approach to identification and dating multiple bubbles in real time has been pioneered by Phillips and Yu (2011) and Phillips *et al.* (2011a, 2011b, 2013). The idea is to spot speculative bubbles as they emerge, not just after they have collapsed. Their point of departure is the observation that the explosive property of bubbles is very different from random walk behaviour. Correspondingly, they have developed a new recursive econometric methodology interpreting mildly explosive unit roots as a hint for bubbles.

Considering the typical difference of stationary versus trend stationary testing procedures for a unit root, we usually restrict our attention to regions of ‘no more than’ a unit root process, i.e. an autoregressive process where $\delta \leq 1$. In contrast, Phillips and Yu (2011) model mildly explosive behaviour by an autoregressive process with a root δ that exceeds unity but is still in the neighbourhood of unity. The basic idea of their approach is to calculate recursively right-tailed unit root tests to assess evidence for mildly explosive behaviour in the data. This is a right-tailed test and therefore differs from the usual left-tailed tests for stationarity. More specifically, consider the following autoregressive specification estimated by recursive least squares:

$$x_t = \mu + \delta x_{t-1} + \sum_{j=1}^J \phi_j \Delta x_{t-j} + \varepsilon_t \quad (11)$$

Here, $x_t \in \{s_t, s_t - f_t^T, s_t - f_t^N\}$, where $s_t - f_t^T$ refers to the log of nominal exchange rate relative to traded goods component, and $s_t - f_t^N$ refers to the log of nominal exchange rate relative to nontraded goods component. The usual $H_0: \delta = 1$ applies, but unlike the left-tailed tests which have relevance for a stationary alternative, Phillips and Yu (2011) have H_a :

¹ This methodology section is in line with Phillips and Yu (2011) and Phillips *et al.* (2011a, 2011b, 2013).

$\delta > 1$, with $\delta = 1 + c/k_n$, where $c > 0$, $kn \rightarrow \infty$ and $kn/n \rightarrow 0$, and these allow for their mildly explosive cases. Phillips and Yu (2011) argue that their tests have discriminatory power because they are sensitive to the changes that occur when a process undergoes a change from a unit root to a mildly explosive root or vice versa. This sensitivity is much greater than in left-tailed unit root tests against stationary alternatives. But this is not all as we know that bubbles usually collapse periodically. Therefore, conventional unit root tests have limited power in detecting periodically collapsing bubbles (Evans, 1991). In order to overcome this shortcoming, Phillips and Yu (2011) have suggested using the supremum of recursively determined ADF t -statistics. The estimation is intended to identify the time period where the explosive property of the bubble component becomes dominant in the price process. The test is applied sequentially on different subsamples. The first subsample contains observations from the initial sample and is then extended forward until all observations of the complete sample are included in the tests. The beginning of the bubble is estimated as the first date when the ADF t -statistic is greater than its corresponding critical value of the right-sided unit root test. The end of the speculative bubble will be determined as the first period when the ADF t -statistic is below the aforementioned critical value.

Following Phillips *et al.* (2011a, 2011b, 2012), we can calculate a sequence of ADF tests. Let $\hat{\delta}_\gamma$ denote the OLS estimator of δ_γ and $\hat{\sigma}_{\delta,\gamma}$ the usual estimator for the SD of $\hat{\delta}_\gamma$ using the subsample $\{x_1, x_2, \dots, x_{[\gamma T]}\}$. The forward recursive ADF test of H_0 against H_a is given by

$$SADF(r_0) = \sup_{r_2 \in [r_0, 1]} \{ADF_{r_2}\} \quad (12)$$

where $ADF = \frac{\hat{\delta}_\gamma - 1}{\hat{\sigma}_{\delta,\gamma}}$. Here, the ADF statistic is computed for the asymmetric interval $[r_0, 1]$. r_0 is the minimum window width. In most applications, r_0 will be set to start with a sample fraction of reasonable size. However, there is one limitation of the SADF test; the starting point is fixed as the first observation of the sample. This implies that in the presence of two bubbles the second bubble may not be detected if it is dominated by the first bubble. Therefore, Phillips *et al.* (2011b) also apply a rolling version of the SADF test, where the starting window moves over the sample. However, the size of the

starting window is still fixed, which limits the power of the test. Phillips *et al.* (2012) have suggested employing the GSADF test as a dating mechanism. The GSADF diagnostic is also based on the idea of sequential right-tailed ADF tests, but the diagnostic extends the sample sequence to a more flexible range. Instead of fixing the starting point of the sample, the GSADF test changes the starting point and ending point of the sample over a feasible range of windows. Phillips *et al.* (2012) demonstrate that the moving sample GSADF diagnostic outperforms the SADF test based on an expanding sample size in detecting explosive behaviour in multiple bubble episodes and seldom gives false alarms, even in relatively modest sample sizes. The reason is that the GSADF test covers more subsamples of the data. The GSADF is able to detect potential multiple bubbles in the data and thus overcomes the weakness of the SADF test:

$$GSADF(r_0) = \sup_{r_2 \in [r_0, 1], r_1 \in [0, r_2 - r_0]} \{ADF_{r_1}^{r_2}\} \quad (13)$$

For more details about both SADF and GSADF tests, interested readers are referred to Phillips and Yu (2011) and Phillips *et al.* (2011a, 2011b, 2013).

VI. Empirical Results and Policy Implications

Tables 1–3 report the empirical results of ADF, SADF and GSADF tests for the natural logarithm of the monthly RMB–dollar exchange rate, RMB–dollar exchange rate – relative price of traded goods ratio and RMB–dollar exchange rate – relative price of nontraded goods ratio for the full sample from July 1995 to October 2013. Figures 1–3 plot the

Table 1. The ADF, SADF and GSADF tests for nominal exchange rate (lex)

	lex	90%	95%	99%
ADF	4.33***	−0.413	−0.026	0.630
SADF	16.00***	1.12	1.41	2.02
GSADF	21.55***	2.91	2.29	3.01

Notes: Critical values of both tests are obtained from Monte Carlo simulation with 2000 replications (sample size 220). The smallest window has 22 observations. *** indicates significance at 1% level.

Table 2. The ADF, SADF and GSADF tests for exchange rate-relative price of traded goods ratio (lex_ft)

	lex_ft	90%	95%	99%
ADF	-2.77	-0.413	-0.026	0.630
SADF	-1.12	1.12	1.41	2.02
GSADF	2.24*	2.00	2.29	2.90

Notes: Critical values of both tests are obtained from Monte Carlo simulation with 2000 replications (sample size 220). The smallest window has 22 observations. * indicates significance at 10% level.

Table 3. The ADF, SADF and GSADF tests for exchange rate-relative price of nontraded goods ratio (lex_fnt)

	lex_fnt	90%	95%	99%
ADF	1.22	-0.413	-0.026	0.630
SADF	3.08***	1.120	1.410	2.020
GSADF	3.75***	2.000	2.290	2.900

Notes: Critical values of both tests are obtained from Monte Carlo simulation with 2000 replications (sample size 220). The smallest window has 22 observations. *** indicates significance at 1% level.

corresponding graphs of the GSADF test. For the expanding window GSADF test, the originating subsamples are selected to ensure estimation efficiency and consist of 220 observations for the monthly data. The critical values of both tests are obtained from Monte Carlo simulation with 2000 replications, simulations, with the data generating process specified to be a random walk mode. We choose the smallest window r_0 by 22 observations, corresponding to around 1% of the data (Phillips *et al.*, 2013). The optimal lag length to account for serial correlation is determined by the method suggested by Campbell and Perron (1991), with the results also compared to the optimal lag length chosen by Schwarz information criterion and AIC.

From Tables 1–3 we find that the null hypothesis of no bubble are rejected for both SADF and GSADF tests for lex, lex_ft and lex_fnt with the exception of the SADF tests for lex_ft. Phillips *et al.* (2012) demonstrate that the moving sample GSADF diagnostic outperforms the SADF test based on an expanding sample size in detecting explosive behaviour in multiple bubble episodes and seldom gives

false alarms, even in relatively modest sample sizes. The reason is that the GSADF test covers more subsamples of the data. Based on this argument, we can conclude that there is evidence of multiple bubbles in the Chinese RMB-dollar exchange rate.

To locate specific bubble periods, we compare the backward SADF statistic sequence with the 95% SADF critical value sequence, which were obtained from Monte Carlo simulations with 2000 replications. Figure 1 displays results for the date-stamping strategy over the period for the RMB-dollar nominal exchange rate (lex). We can see that there is evidence of multiple bubbles in lex. If we look at Fig. 1, there is evidence of bubble occurring in the earlier period of 2005 for lex, which is precisely the period of exchange rate regime reform undertaken in China in July 2005. It seems that the Chinese RMB-dollar exchange rate is in the bubble period. The bubble does not collapse at this moment. The reason might be that the Chinese exchange rate regime reform conducted in 2005 when the RMB-dollar exchange rate appreciated for 2% for once and the exchange rate was allowed for managed floating from then on. Since then the reform progress is irreversible and rapid economic development urges further appreciation. Figure 1 also indicates that there is another bubble occurring in lex in the period of 2008 when subprime crisis exploded. But there is no evidence of explosive bubbles in exchange rate before 2005 as the exchange rate was fixed at that time.

The explosiveness in the nominal exchange rate could be driven either by rational bubbles or explosive fundamentals. Table 3 shows the results for the ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$. The exchange rate remains explosive after the relative prices of traded goods are accounted for at 10% significance level. But the null hypothesis of no bubble cannot be rejected at both 5% and 1% significance levels. Figure 2 displays that there is evidence of small bubble during the period of 2005–2006 in lex_ft. Thus the relative prices of traded goods f_t^T play no role in explaining the explosiveness during 2005–2006 in the nominal exchange rate, but explaining the explosiveness during 2007–2008 in the nominal exchange rate. According to Engel (1999), the reason might be that during 2005–2006 the PPI in China had

fallen by 46%, whereas in the US the decrease in the PPI was 78%. The movement in the price difference is still not enough to account for much of the real exchange rate movement. However, during 2007–2008, the PPI in China had fallen by 135%, whereas in the US the decrease in the PPI was 700%. Therefore, the larger movement in these prices can account for the movement of the RMB–dollar exchange rate (we use year-on-year PPI data for both China and the US). Engel (1999) points out that one potential problem with the measures of traded goods prices is that there are extensive producer services – distribution, retailing, advertising and so on – whose costs are included in the traded goods price. Since these marketing services are largely nontradable, this hypothesis could revive the possibility that the relative price of nontradable goods can account for much of the exchange rate movements. However, Engel (1999) has further examined the OECD output prices and found evidence that weighs in against this hypothesis. He finds that the traded goods component accounts for almost all the exchange rate movements. The marketing and distribution prices should not be an important component of these output prices.

For lex_fnt , Table 3 shows the results for the ratio of the exchange rate to the nontraded goods fundamental $s_t - f_t^N$. Both SADF and GSADF show that the exchange rate remains explosive after the relative prices of traded goods are accounted for at 1% significance level. Figure 3 displays the result of the GSADF test graphically. There is a small bubble in lex_fnt during the period of 2005–2006, rather than similar to the bubble in nominal exchange rate not collapsing from 2005, the bubble in lex_fnt start and collapse quickly. Regarding lex_fnt , we find that there is evidence of another bubble during the 2007–2008 subprime crisis periods and collapse in 2008. Therefore, the explosive behaviour in the nominal RMB–dollar exchange rate may not be driven by the relative prices of nontraded goods between the US and China. This finding is grossly at odds with the predictions of the traditional theories of exchange rate that movements in the exchange rate are driven by productivity differentials through the relative price of nontradable goods (Balassa, 1964; Samuelson, 1964; Mendoza, 1991; Rebelo and Vegh, 1995; Stockman and Tesar, 1995).

What can possibly explain why the movements in the relative price of traded goods are so important for the exchange rate? One possibility is terms-of-trade movements. If there are fluctuations in the relative prices of goods that constitute the traded goods price indexes and if the weights these goods receive are different in the US and foreign price indexes, then the relative traded goods price indexes will fluctuate. The second reason might be that the average income per capita in China is low relative to the US. As Table 3 of Rogoff (1996) demonstrates, the behaviour of exchange rates for low-income countries might be very different. The absolute purchasing power of the currencies of low-income countries is much higher than that of high-income countries, although within the two groups there seems to be little relationship between income and purchasing power. It seems that the disparity between low-income and high-income countries' purchasing powers reflects differences in the relative prices of nontraded goods. Although our results attribute the movements in the traded goods price indexes to failures of the law of one price, the conclusion would be consistent with the findings of a number of studies (Engel, 1993, 1999; Froot *et al.*, 1995; Rogers and Jenkins, 1995; Engel and Rogers, 1996, 1998; Knetter, 1997; Chanthapun, 2005; Betts and Kehoe, 2006; Burstein *et al.*, 2006; Chaban, 2006). Thirdly, according to Diba and Grossman (1988a) if the fundamental value and the actual exchange rate are both explosive, then this may not necessarily indicate a bubble in the exchange rate, but instead reflect some regime switching or time-varying fundamentals.

The existence of bubbles has some implications on the economy. The effects of bubbles can vary depending on different factors and the outcome might be distinct as well for every single agent. Bubbles can harm an economy, but it may simply generate a strong temporary deviation from a price tendency, or it could even benefit the economy (Jiménez, 2011). However, there is clear empirical evidence that every single bubble generates a redistribution of wealth, directly or indirectly, among the various agents in the economy. Tirole (1985) tests the impact of economic bubbles on modelled economies. Later on, Grossman and Yanagawa (1993) expand Tirole's model to include economies that grow in the long run at an endogenous rate. The conclusions are

that ‘bubbles retard the growth of the economy, perhaps even in the long run, and reduce the welfare of all generations born after the bubble appears’. The fact that bubbles attract capitals that otherwise would have been allocated in more productive assets provides support for these theories. Moreover, Jiménez (2011) notes that in these models bubbles can only exist on nonaccumulable useless assets and the impact of both the emergence and the burst of the bubble, although they might be beneficiary for the current generation, have serious retards on economic growth for future generations. Bubbles might lead to economic distortions as well as financial and real economy instability, and have effects on current output growth, aggregate spending and expected inflation (Roubini, 2006a). The negative impacts of asset price booms and bursts on the economy have also been shown in a number of studies (Bordo and Olivier, 2002; Borio and Lowe, 2002; Helbling and Bayoumi, 2003).

These bubbles generated a clear redistribution of wealth from the middle class – which increased its debts by almost 50% on average – towards the ruling elites, generated an almost immediate unemployment of 10% on average, implied a bailout of more than \$800 billion in the US, more than doubled the national debt and plunged world Western economies into severe recessions (Jiménez, 2011). This is still to some extent experienced now. When the economy faces a bubble before the burst occurs, it is time to try to control the bubble, so just monetary policy is mainly relevant. However, after the bubble bursts, fiscal policy becomes relevant as well. As regards controlled bubble bursts, it seems to be empirical evidence supporting monetary policy tightening deterring bubbles while not causing any financial or economic crash. Roubini (2006b) states that

monetary policy should react to asset prices and should try to “prick” or “burst” asset bubbles. Bubbles that are growing excessively large lead to economic and investment distortions that are dangerous and likely to eventually trigger bubble bursts whose real and financial consequences are severe. Thus, optimal monetary policy should preemptively deal with asset bubbles rather than just mop up the mess that they cause after they burst.

Roubini (2006b) advocates that ‘asset prices should enter directly in the reaction function of the optimizing monetary authority, above and beyond

the direct effects that such asset prices have on expected inflation and current growth’. This view was also shared in Roubini (2006a). Further, Filardo (2000, 2001) theoretically demonstrates that even under uncertainty optimal monetary policy should react to asset prices. React to the overall asset price, regardless if there is uncertainty about a bubble component or not. Moreover, the need for adopting fiscal policy rules that secure a sound medium-term orientation of fiscal policies while leaving adequate short-term flexibility had been advocated by Jaeger and Schuknecht (2004). However, Jaeger and Schuknecht (2004) note that the success of such fiscal policy rules hinges largely on the credibly containing expenditure growth and preventing tax cuts during the ‘high temptation phases’ towards the end of a prolonged boom and at the onset of bust phases in asset prices. Following these arguments and unlike some pro-nonintervention authors (Bernanke, 2002, 2004; Greenspan, 2004; Kohn, 2004; Blinder and Reis, 2005) who argue that uncertainty about a bubble’s existence precludes any policy response, we are of the opinion that both monetary and fiscal policies are relevant for curtailing the exchange rate bubbles given that our study is able to detect the presence of multiple bubbles. However, caution needs to be taken when implementing such policies to avoid a longer-lasting negative impact on the economy.

VII. Conclusions

This study investigates whether there were explosive bubbles in the Chinese RMB-dollar exchange rate using GSADF tests proposed by Phillips *et al.* (2013). We have found several results as follows.

First, from Fig. 1, our empirical results indicate that there were explosive multiple bubbles in the Chinese RMB-dollar nominal exchange rate after the 2005 exchange rate regime reform, which was not shown in previous studies.

Second, we find neither the relative prices of traded goods nor the relative price of nontraded goods account for the explosive behaviour in the nominal exchange rate in 2005–2006 between the US and China, as from Figs 1–3 there is evidence of small bubbles in both the ratio of exchange rate to traded goods and the ratio of exchange rate to non-trade goods during 2005–2006. This bubble should

be caused by a structural break that the most influential exchange rate regime reform conducted in 2005. According to Engel (1999), the reason why neither of the two ratio accounts for the bubble might be during 2005–2006 the PPI in China had fallen by 46% whereas in the US the decrease in the PPI was 78%. The movement in the price difference is still not enough to account for much of the exchange rate movement.

Third, the explosiveness in exchange rate during 2007–2008 is determined by the relative prices of traded goods but not the relative price of nontraded goods, as from Figs 1–3 there is a small bubble in the ratio of exchange rate to nontraded goods but not a bubble in the ratio of exchange rate to traded goods. The reason might be that during 2007–2008 the PPI in China had fallen by 135% whereas in the US the decrease in the PPI was 700%. Therefore, the larger movement in these prices can account for the movement of the RMB–dollar exchange rate. Engel (1999) has further examined the OECD output prices and find evidence that the traded goods component accounts for almost all the exchange rate movements. The marketing and distribution prices should not be an important component of these output prices.

Our findings are in line with Engel (1999), Betts and Kehoe (2006, 2008) and Bettendorf and Chen (2013), who show that the relative prices of traded goods explain most of the movements in exchange rates. However, empirical evidence shows that bubbles generate a redistribution of wealth and can have negative financial and real consequences. Therefore, some tightening monetary and fiscal policies are required in China since these are the most efficient and effective under a bubble burst scenario.

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