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# Are there bubbles in exchange rates? Some new evidence from G10 and emerging market economies

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ABSTRACT

The existence, or otherwise, of bubbles has become a topical issue in economics and finance, particularly following the Global Financial Crisis. Using the generalized sup ADF (GSADF), unit root tests of Phillips et al. (2015a, PSY) we investigate evidence for exchange rate bubbles in some G10, Asian and BRICS countries from Mar.1991-Dec.2014. We conclude that the US$-Mexican Peso crisis of 1994–95 was a bubble. Of particular interest to financial market trading, is that newly emerging countries, with relatively shallow financial markets, may be more likely to exhibit bubbly behavior in foreign exchange markets than more mature G10 countries.

## 1. Introduction

Despite theoretical arguments against the existence of bubbles for finitely lived assets in rational markets, experiences from the Global Financial Crisis have once again put the possibility that financially driven bubbles exist, at least empirically, back into the spotlight where a simple and straightforward definition of a bubble is a deviation of the market price from (the asset's) fundamental value. With the advent of the 'unit root' revolution of the 1970 s, testing for bubbles using time series data has, until recently, typically focused on cointegration-type methods, testing for the existence of a single, long-run, linear cointegrating relationship between the 'price' and its 'fundamental' value. Early applications of such methods e.g. Kearney and MacDonald (1990), have recently been extended see, for example, Maldonado et al. (2016), to consider periodically collapsing bubbles.

It is, however, the recent developments in 'right-tailed only' unit root tests (e.g., Phillips et al., 2011, PWY, Phillips et al., 2015b, Phillips et al., 2015a, PSY) which have been designed to specifically test for the presence of bubbles (where several bubble episodes may exist and be identified within a timeline, punctuated by 'no-bubble' periods), that have become one of the most popular and most thoroughly researched tests for bubbles. Applications of the PSY approach have been broad e.g., housing markets, agricultural prices and energy prices and extensive (see e.g., Phillips and Yu, 2011, Homm and Breitung, 2012, Etienne et al., 2014, Greenaway-McGrevy and Phillips, 2015, Harvey et al., 2016, Shi et al., 2016)[1]. Bettendorf and Chen (2013) and Jiang et al. (2015) used the PSY method to test for the existence of bubbles in exchange rates, but their examples involved only one each of a bilateral rate with the US dollar, i.e., the Sterling-US Dollar and Chinese RMB-US Dollar exchange rates, respectively. Their results suggest that the explosiveness identified in the nominal exchange rate is likely driven by either exchange rate fundamentals (the relative prices of traded goods or nontraded goods) or the formation of 'rational bubbles'.

The results of three other studies; Jirasakuldech et al. (2006); Maldonado et al. (2012); and Maldonado et al. (2016), are worth emphasizing here as their results, bear some comparison with ours. Jirasakuldech et al. (2006) investigate the existence of bubbles in bilateral exchange rates between the US Dollar and five currencies including the South African Rand using three different approaches and provide no evidence of bubbles in all currency pairs[2]. Maldonado et al. (2012) apply a model, which is extended from Van Norden (1996) model, to the exchange rate of the Brazilian Real to the US Dollar during March 1999-February 2011. Maldonado et al. (2016) also examine the bubbles in the exchange rate of the BRICS countries currency relative to the US Dollar using the bubble model developed in Maldonado et al. (2012) and conclude the presence of rational bubbles for all countries.

This paper has two main aims and contributions. Firstly, we apply

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[1] Chan et al. (2001) and Roche (2001) also investigate the existence of bubbles in the housing markets of Hong Kong and Dublin using alternative approaches.
[2] The asymptotic properties of these authors' methods have not been investigated, or at least published.

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the generalized sup ADF (GSADF) test of Phillips et al. (2015a, PSY) to investigate the presence of exchange rate bubbles in a very wide range of countries, in particular, some G10 and a range of emerging markets countries (including some Asian and the BRICS). This allows us to consider whether exchange rate bubbles might be more likely to arise in certain countries (perhaps those with less well developed trading relationships or those where governments retain a role in trading behavior), rather than in the highly developed countries of for example, the UK and US. The second aim is to study the importance of model formulation issues highlighted by Phillips et al. (2014) in right-tailed unit root tests. In particular, the model specification for constructing the null hypothesis with/without an intercept is considered. By comparing two model formulations, our results show the inclusion of the intercept term for model specification under the null hypothesis affects the theory and date-stamping strategy of the PSY approach. This also allows us to show, quite clearly, situations where the typical use of the PSY approach fails to distinguish (without further analysis) periods of collapse from periods of recovery, where it is only the former case that relates to the growth and ultimate collapse of a bubble.

The importance of the results presented here are significant and messages of the paper multi-faceted. We have applied the popular and well-researched method of PSY to the widest and most extensive range of exchange rates currently undertaken. Readers can ascertain, in one single paper, the current evidence on explosive behaviour in both/either of the *nominal exchange rates*, and *exchange rate fundamentals*. Using the examples and time periods identified here, researchers can try and ascertain the particular, perhaps idiosyncratic reasons why bubbles arose. The paper presents, for the first time, empirical evidence on the importance of how the PSY test might/should be applied in practice, in particular, the importance of identifying genuine bubbles from the often observationally equivalent periods of 'collapse and recovery' and also the need to consider whether the conclusions of the test should be based upon 'with' or 'without intercept' results. This message goes beyond the particular example of this paper and constitutes a potential pitfall for both (some) published papers and (if ignored) future research.

The paper is therefore organized as follows. Section 2 provides a review of the theory of the role of fundamentals in determining the nominal exchange rate. This section summarises a well known theoretical literature and no claims for novelty are made here. To aid exposition, we follow the notation and terminology of Bettendorf and Chen (2013) and Jiang et al. (2015). Section 3 provides a brief description of the GSADF and SADF tests of Phillips et al. (2015a) and Phillips et al. (2011). Section 4 describes the data. Section 5 provides empirical results for G10 and emerging markets countries and Section 6 concludes.

## 2. Exchange rates: theoretical background

Following the work of Bettendorf and Chen (2013) and Jiang et al. (2015), we also define the economic fundamental for the nominal exchange rate as the price differential ($f_t$):

$$f_t = p_t - p_t^*, \tag{1}$$

where $p_t$ and $p_t^*$ denote the domestic and foreign price indices in logarithm. Engel (1999) shows that the price index for a domestic country can be expressed as a combination of traded and non-traded goods as following:

$$p_t = (1-\alpha)p_t^T + \alpha p_t^N, \tag{2}$$

where $p_t^T$ and $p_t^N$ denote the traded and non-traded goods price indices in logarithm, respectively. The foreign price index can be defined in a similar way:

$$p_t^* = (1-\beta)p_t^{T*} + \beta p_t^{N*}. \tag{3}$$

The price differential ($f_t$) therefore can be decomposed into the traded goods component ($f_t^T$) and the non-traded goods component ($f_t^N$):

$$p_t - p_t^* = (p_t^T - p_t^{T*}) + \alpha(p_t^N - p_t^T) - \beta(p_t^{N*} - p_t^{T*}). \tag{4}$$

The producer price index (PPI) is adopted here to measure the price level of traded goods and the traded goods component is constructed from following Engel (1999):

$$f_t^T = \ln(PPI_t) - \ln(PPI_t^*). \tag{5}$$

The relative non-traded goods component is constructed from the aggregate consumer price indices (CPI) relative to aggregate PPI:

$$f_t^N = \ln(CPI_t) - \ln(PPI_t) - (\ln(CPI_t^*) - \ln(PPI_t^*)). \tag{6}$$

## 3. Method

Phillips et al. (2011) proposed a sup ADF (SADF) test based procedure that can test for evidence of price exuberance and date stamp its origination and collapse. Homm and Breitung (2012) conducted simulation studies to show that the SADF test is an effective bubble detection algorithm. One highlight of this new approach is the ability to capture periodically collapsing bubbles of Evans (1991). However, as discussed in Phillips et al. (2015a), the SADF test has limited ability to detect the presence of multiple bubbles. The SADF test is recursively applied to the sample data and is implemented as follows. We apply the Augmented Dickey-Fuller (ADF) test to a time series $x_t$ for the null of a unit root against the alternative of explosive behavior. The following autoregressive specification for $x_t$ is estimated by least squares:

$$x_t = \mu_x + \delta x_{t-1} + \sum_{j=1}^{J} \phi_j \Delta x_{t-j} + \varepsilon_{x,t}, \quad \varepsilon_{x,t} \sim \text{NID}(0, \sigma_x^2), \tag{7}$$

for some given value of the lag parameter J, where NID denotes independent and normally distributed. The null hypothesis of this test is $H_0$: $\delta = 1$ and the alternative hypothesis is $H_1$: $\delta > 1$. Eq. (7) is estimated repeatedly using subsets of the sample data incremented by one additional observation at each pass in the forward recursive regression. The window size $r_w$ expands from $r_0$ to 1, where $r_0$ is the smallest sample window width fraction. The starting point $r_1$ is fixed at 0, and the end point of each sample ($r_2$) equals $r_w$ and changes from $r_0$ to 1. The SADF statistic is therefore defined as the sup value of the ADF statistic sequence:

$$\text{SADF}(r_0) = \sup_{r_2 \in [r_0, 1]} ADF_0^{r_2}$$

Unlike the SADF test, the GSADF test is extended by using a more flexible window size and has great power in detecting the presence of multiple bubbles. The end point $r_2$ varies from $r_0$ (the minimum window size) to 1. The start point $r_1$ is also allowed to vary from 0 to $r_2 - r_0$. The GSADF statistic is the largest ADF statistic over range of $r_1$ and $r_2$. The key difference between the SADF and GSADF is the window size of starting point $r_1$. The GSADF statistic is therefore defined as:

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

In general, a number of factors can affect the bubble detection results for example, the full sample/subsample, the minimum window size $r_0$, the lag length, and model specification under the null hypothesis. Firstly, the bubble detection results may differ if the GSADF test is applied to a subsample of (truncated) data rather than the full sample. This phenomenon is more obvious for the SADF test. Secondly, as stated in Phillips et al. (2015a), the asymptotic GSADF distribution depends on the smallest window size $r_0$. The minimum window size $r_0$ needs to be large enough to allow initial estimation, but it should not be too large to miss the chance of detecting an early bubble period. We therefore follow Phillips et al. (2015a) and let $r_0 = 0.01 + 1.8/\sqrt{T}$, where T is number of observation[3]. They recommend this rule for empirical use as it provides satisfactory size and power performance. Thirdly, the choice of the lag length is also crucial. If the lag order is over-specified, then the size distortion would be more severe for the GSADF test than the SADF test. A small fixed lag order approach is used in this study as suggested by Phillips et al. (2015a). The finite critical values are obtained from Monte Carlo simulation with 2000 replications. Finally, the model specification under the null hypothesis plays an important role in assessing the evidence of bubbles. Phillips et al. (2014) have investigated different formulations of the null and alternative hypothesis in the right-tailed unit root test of Phillips et al. (2011). These formulations use various specifications of the regression models (e.g., with/without an intercept or with/without a trend) for constructing the empirical tests to assess the evidence of explosiveness. Model specification was shown to affect both the finite sample and the asymptotic distributions and they suggested an empirical model specification with an intercept only for practical use. The model specification issue is not discussed in either Bettendorf and Chen (2013) or Jiang et al. (2015).

A number of studies have followed Phillips et al.'s (2014) suggestion to include an intercept in the right-tailed unit root test. Hence, many empirical papers have reported rejections of the null suggesting periods of rapid increase in prices associated with a growing bubble, when in fact the data identifies a 'collapse' or a 'collapse and recovery' phase and not a bubble. Visual inspection can usually resolve these cases, although it also seems that false (positive) bubbles also seem to be reported when an intercept is included. An example of 'collapse episode' and 'collapse and recovery episode' can be seen in Fig. 1 below. The backward SADF statistic (blue line) and its 95% critical value (red line) for Fig. 1a suggests a number of 'bubbles' as the test statistic exceeds the relevant critical value. However, the plot of the actual data (green line) shows that the data is continuously declining (a collapse period and not a series of bubbles). Fig. 1b presents data and test results consistent that relate to a 'collapse and recovery' episode and a genuine 'bubble'. In this paper, we consider two different model specifications for the null hypothesis in the right-tailed unit root tests (a model without an intercept as in Eq. (8) and a model with an intercept in Eq. (9)) to explore the evidence of bubbles and compare the results obtained from both formulations. The model specification is explained as follows. In PWY of Phillips et al. (2011), the null hypothesis is:

$$H_{01}\colon y_t = y_{t-1} + \varepsilon_t, \quad \varepsilon_{x,t} \sim \mathrm{NID}(0, \sigma^2). \tag{8}$$

The second specification for the null is obtained from Diba and Grossman (1988):

$$H_{02}\colon y_t = \alpha + y_{t-1} + \varepsilon_t, \quad \text{where}\,\alpha\,\text{is the constant.} \tag{9}$$

## 4. Data

The monthly exchange rates for some G10, Asian and BRICS countries are obtained from Quandl (https://www.quandl.com/) and the IMF International Financial Statistics, and these exchange rates are the end of period rates. We consider the following G10 currencies (e.g., British Pound (GBP), Canadian Dollar (CAD), Japanese Yen (JPY), Norwegian Krone (NOK), Swedish Krona (SEK), Swiss Franc (CHF)) and test for the existence of exchange rate bubbles. We also consider the US Dollar against several emerging market exchange rates in Asia including the Indonesian Rupiah (IDR), Korean Won (KRW), Malaysian Ringgit (MYR), Philippine Peso (PHP), Singapore Dollar (SGD) and Thai Baht (THR). In addition, we test for the existence of exchange rate bubbles in the US Dollar against several other emerging markets currencies: Brazilian Real (BRL), Indian Rupee (INR), South African Rand (ZAR), Colombian Peso (COP) and Mexican Peso (MXN). The CPI and PPI are obtained from the IMF International Financial Statistics. The monthly sample data used for our analysis are from March 1991 to December 2014[4]. All series have been transformed into logarithms.

## 5. Results

We present our results in four sections. Sections 5.1, 5.2 5.3 and 5.4 provide the empirical results for G10, Asian, BRICS and other emerging markets countries, respectively.

### 5.1. Results for G10 countries

Results for the G10 exchange rates are presented in Tables 1, 2, 3 using different model specifications (with/without an intercept) under the null hypothesis[5]. Under the model specification 'without an intercept", no strong evidence of explosiveness is detected in these currency pairs. If the model specification allows an intercept term, we do not find significant evidence of explosive behavior in these currencies except for the Sterling-Swiss Franc (GBP/CHF) and Sterling-Japanese Yen (GBP/JPY) based on the test statistic. We therefore only discuss the bubble-detection results for these two exchange rates.

#### 5.1.1. GBP/CHF

The left panel of Fig. 2 compares the backward SADF statistic with the 95% critical value sequences for the nominal exchange rate $s_t$, the relative ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$ and the relative ratio of the exchange rate to the non-traded goods fundamental $s_t - f_t^N$ using a model specification with an intercept for assessing the evidence of bubbles, respectively. The right panel of Fig. 2 presents bubble detection results for $s_t$, $s_t - f_t^T$ and $s_t - f_t^N$ using a model specification without an intercept. Table 1 suggests the existence of explosive behavior in the nominal exchange rate $s_t$ at the 1% significance level, which indicates the existence of explosive subperiods. Fig. 2a compares the backward SADF statistic with 95% critical value sequences for the nominal exchange rate $s_t$. Multiple episodes can be identified in Fig. 2a including 1995M05-1995M07, 2008M02-2008M04, 2008M09-2009M01 and 2011M05-2011M08, and most of these episodes are just 'collapse' episodes.

Fig. 2c and e display the backward SADF statistic sequences for $s_t - f_t^T$ and $s_t - f_t^N$, respectively. We find a 'collapse and recovery' episode between 2008M09 and 2009M01 in both figures. In addition, a 'collapse and recovery' episode from 2011M04 to 2011M09 and a 'collapse' episode from 1995M02 to 1996M01 are also identified in Fig. 2e. On a close inspection of the date-stamping outcomes using a model specification with an intercept, we find little evidence of bubble. One of the take home messages is that the rejection of the null hypothesis under the assumption 'with an intercept' in the PSY approach could lead to false positive identification of bubbles. In this example, the PSY approach identifies several 'collapse' episodes but not bubbles.

However, under the null hypothesis without an intercept term, we find no significant evidence of explosiveness in all three series ($s_t$, $s_t - f_t^T$ and $s_t - f_t^N$) as the null hypothesis of explosive behavior cannot be rejected at the 10% significance level. Moreover, the backward SADF statistic sequences no longer detect the 'collapse and recovery' episode

[3] We use this rule for choosing $r_0$ for most exchange rates except the US Dollar against the Mexican Peso.

[4] The modern Brazilian Real was introduced in 1994. The sample data for Brazil from June 1994 to December 2014 is used for our analysis. The data for Mexico and the Philippines ranges from January 1993 to December 2014.

[5] The critical values for the null hypothesis with an intercept: 1.8569 (90%), 2.0977 (95%), 2.6217 (99%). The critical values for the null hypothesis without an intercept: 3.1247 (90%), 3.5343 (95%), 4.2359 (99%).

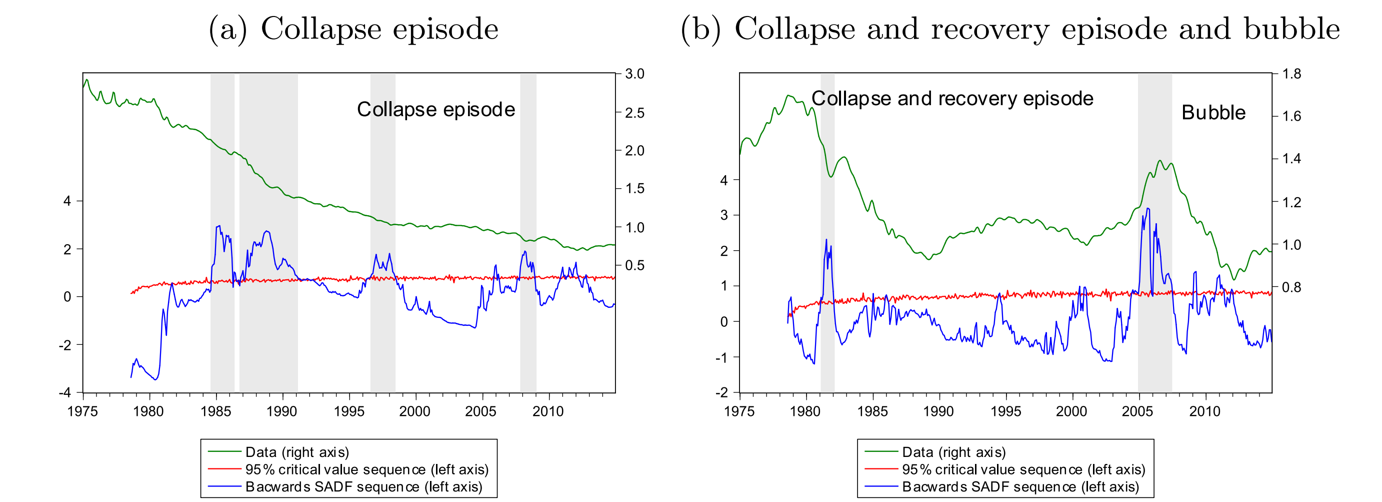


**Fig. 1.** Examples of the collapse episode, the collapse and recovery episode and the bubble. (a) Collapse episode. (b) Collapse and recovery episode and bubble. (For interpretation of the references to color in this figure, the reader is referred to the web version of this article.)

**Table 1**
The GSADF test for exchange rate in G10 countries.

| Exchange rate | Test Stat under $H_0$ with an intercept | Episodes | Test Stat under $H_0$ without an intercept | Episodes |
|---|---|---|---|---|
| **GBP/CAD** | | | | |
| $s_t$ | 1.9283[a*] | 13M12-14M05 | 1.9787 | |
| $s_t - f_t^N$ | 1.8906* | 13M12-14M05 | 2.1902 | 98M07-99M01, 14M01-14M04 |
| $s_t - f_t^T$ | 1.7400 | 13M12-14M03 | 2.0057 | |
| **GBP/CHF** | | | | |
| $s_t$ | 2.9084[b***] | 95M05-95M07, 08M02-08M04 08M09-09M01, 11M05-11M08 | 2.0548 | 97M11-98M04 |
| $s_t - f_t^N$ | 2.3762[c**] | 95M02-96M01, 08M09-09M01 11M04-11M09 | 2.0789 | 07M05-07M08 |
| $s_t - f_t^T$ | 2.6425*** | 96M10-97M08, 08M11-09M01 | 2.6425 | 97M11-98M07, 99M10-00M05 |
| **GBP/JPY** | | | | |
| $s_t$ | 3.0534*** | 08M10-09M03 | 3.0184 | 97M10-98M09, 07M05-07M07 |
| | | 13M11-14M01 | | 14M04-14M12 |
| $s_t - f_t^N$ | 2.5985** | 06M12-07M02, 07M04-07M07 | 3.0699 | 97M11-98M10, 06M10-07M11 |
| | | 08M10-09M03, 13M11-14M01 | | 14M04-14M12 |
| $s_t - f_t^T$ | 2.8423*** | 96M10-97M04, 98M03-98M09 | 3.3178* | 96M10-97M05, 97M10-98M10 |
| | | 08M09-09M02, 13M11-13M12 | | 06M12-07M10, 14M04-14M12 |
| **GBP/NOK** | | | | |
| $s_t$ | 1.2835 | 97M05-97M08 | 1.9141 | |
| $s_t - f_t^N$ | 0.9729 | 97M06-97M08 | 2.1358 | 00M08-00M11 |
| $s_t - f_t^T$ | 1.3922 | 97M06-97M08, 08M04-08M09 | 2.2619 | |
| | | 10M01-12M04 | | |

[a*] indicates significance at the 10% level.
[b***] indicates significance at the 1% level.
[c**] indicates significance at the 5% level.

in 2008-2009. These results suggest that the intercept term can potentially affect the asymptotic distributions of the PSY approach.

### 5.1.2. GBP/JPY

Under the null hypothesis 'with an intercept', Table 1 provides strong evidence of explosive behavior in the nominal exchange rate $s_t$ for GBP/JPY at the 1% significance level. As shown in Fig. 3a, there is an episode between 2008M10 and 2009M03 and $s_t$ remains explosive even if both exchange rate fundamentals are accounted for. If we look at all three series ($s_t$, $s_t - f_t^T$ and $s_t - f_t^N$) in Figs. 3a, c and e, all three series are declining and then recovering between 2008M10 and 2009M03 and rather than growing are collapsing. We may regard this special type of episodes as a 'collapse and recovery' episode but not a bubble. There is a short-lived bubble during 2013M11-2014M01 in Fig. 3a. Both the relative prices of traded goods ${f_t}^T$ and the relative prices of non-traded goods ${f_t}^N$ play no role in explaining the explo-

**Table 2**
The GSADF test for exchange rate in G10 countries.

| Exchange rate | Test Stat under $H_0$ with an intercept | Episodes | Test Stat under $H_0$ without an intercept | Episodes |
|---|---|---|---|---|
| **GBP/SEK** | | | | |
| $s_t$ | 1.1704 | 95M10-95M11, 08M02-08M04 | 2.2646 | 98M06-98M12, 99M03-99M06 99M01-00M04, 00M08-02M04 |
| $s_t - f_t^N$ | 0.5572 | | 2.6073 | 98M07-98M12, 99M11-02M10 |
| $s_t - f_t^T$ | 1.6099 | 95M10-95M11 | 2.6115 | 98M05-00M01 |
| **CAD/JPY** | | | | |
| $s_t$ | 0.6021 | | 2.3830 | 97M11-98M09, 07M04-07M11 |
| $s_t - f_t^N$ | 0.8551 | 94M02-94M08, 95M02-95M06 | 2.6121 | 97M12-98M08, 05M10-08M01 |
| $s_t - f_t^T$ | 0.6871 | | 2.6392 | 97M11-98M09, 05M09-07M12 |
| **CAD/NOK** | | | | |
| $s_t$ | 1.6490 | 02M07-03M01 | 1.9936 | |
| $s_t - f_t^N$ | 1.0078 | 00M08-00M10 | 2.1232 | |
| $s_t - f_t^T$ | 1.0078 | | 1.5926 | |
| **CAD/SEK** | | | | |
| $s_t$ | 0.5654 | | 2.3567 | 01M05-01M08 |
| $s_t - f_t^N$ | 0.8100 | | 2.6194 | 01M02-02M01 |
| $s_t - f_t^T$ | 0.1236 | | 1.9971 | 01M05-01M07 |
| **CHF/CAD** | | | | |
| $s_t$ | 0.4434 | | 0.9985 | |
| $s_t - f_t^N$ | 0.8767 | 95M01-95M07 | 1.2186 | |
| $s_t - f_t^T$ | 0.4805 | | 0.5891 | |

**Table 3**
The GSADF test for exchange rate in G10 countries.

| Exchange rate | Test Stat under $H_0$ with an intercept | Episodes | Test Stat under $H_0$ without an intercept | Episodes |
|---|---|---|---|---|
| **CHF/JPY** | | | | |
| $s_t$ | 0.5931 | | 2.2867 | 03M03-03M07, 06M11-08M08 |
| $s_t - f_t^N$ | 0.3783 | | 2.3967 | 03M03-03M07, 06M04-08M09 |
| $s_t - f_t^T$ | 0.7452 | | 2.5739 | 02M12-03M09, 06M06-08M08 |
| **CHF/NOK** | | | | |
| $s_t$ | 1.5892 | 96M12-97M03 | 2.5214 | 94M07-96M09 |
| $s_t - f_t^N$ | 1.3422 | 93M11-94M03, 95M02-95M05 10M11-12M04 | 3.1743[a*] | 94M07-96M10, 10M11-12M11 13M06-14M12 |
| $s_t - f_t^T$ | 3.0592[b***] | 96M10-97M04 | 2.0150 | 94M06-96M01 |
| **CHF/SEK** | | | | |
| $s_t$ | 1.8713* | 93M11-94M01, 01M08-01M11 08M11-09M03 | 2.6662 | 93M11-95M10 |
| $s_t - f_t^N$ | 1.8988* | 93M11-94M03, 95M02-95M06 01M09-01M10, 08M11-09M03 | 3.0832 | 93M11-96M05, 00M08-03M05 05M04-06M09, 08M09-12M06 |
| $s_t - f_t^T$ | 1.0940 | 08M11-08M12 | 1.7937 | 95M02-95M05 |
| **NOK/JPY** | | | | |
| $s_t$ | 1.0718 | 08M11-09M01 | 2.4754 | 02M11-03M07, 07M01-07M11 |
| $s_t - f_t^N$ | 1.2103 | 08M10-09M01 | 1.7607 | |
| $s_t - f_t^T$ | 1.0280 | 96M09-97M02, 08M04-08M09 | 2.8433 | 96M05-97M03, 02M12-03M03 05M03-08M09 |
| **NOK/SEK** | | | | |
| $s_t$ | 0.5022 | | 1.9339 | |
| $s_t - f_t^N$ | 0.6544 | | 1.7466 | |
| $s_t - f_t^T$ | 0.4901 | | 1.6884 | |

[a*] indicates significance at the 10% level.
[b***] indicates significance at the 1% level.

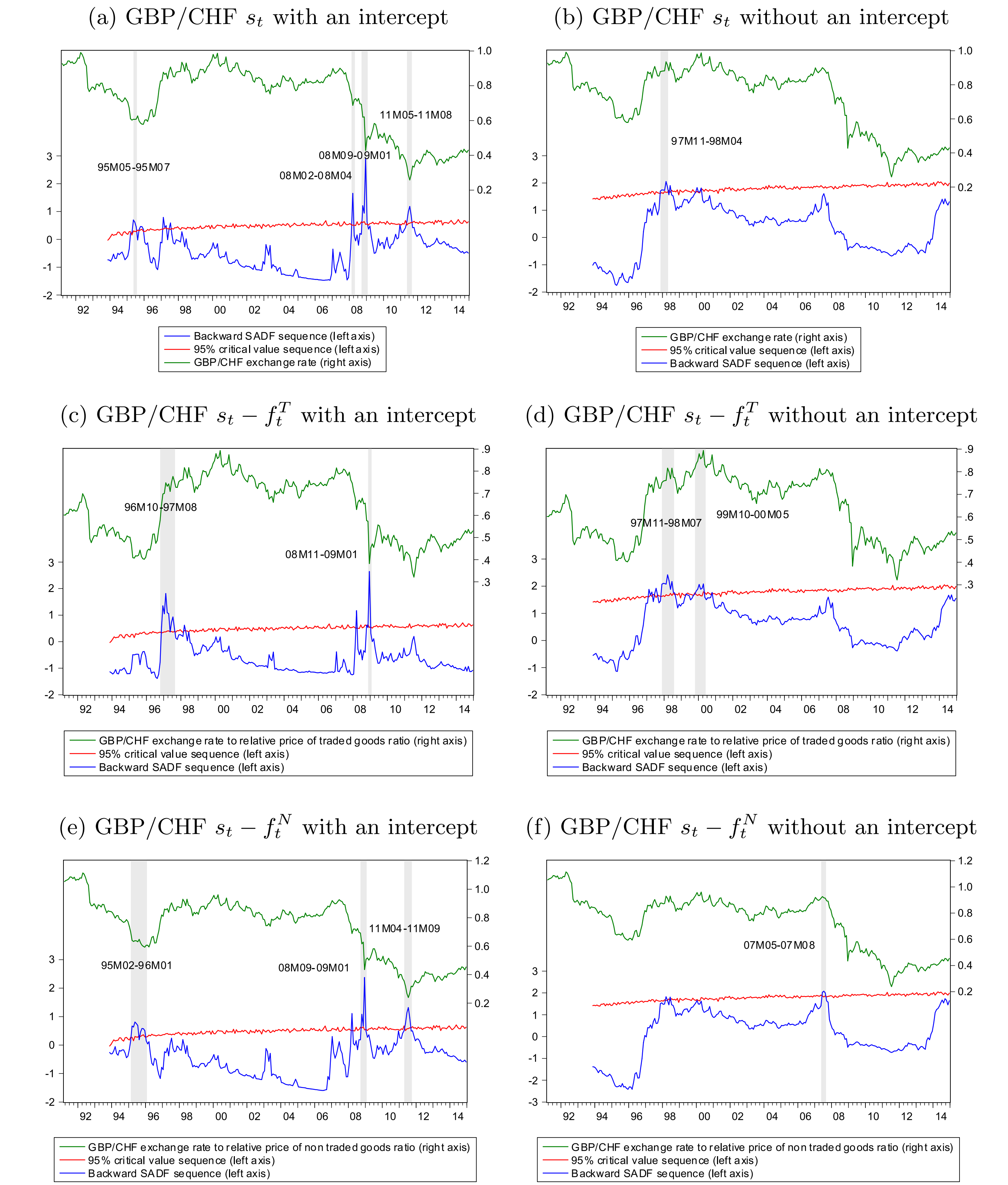


**Fig. 2.** Dating strategy for GBP/CHF nominal exchange rate $s_t$, the relative ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$ and the relative ratio of the exchange rate to the non-traded goods fundamental $s_t - f_t^N$. (a) GBP/CHF $s_t$ with an intercept. (b) GBP/CHF $s_t$ without an intercept. (c) GBP/CHF $s_t - f_t^T$ with an intercept. (d) GBP/CHF $s_t - f_t^T$ without an intercept. (e) GBP/CHF $s_t - f_t^N$ with an intercept. (e) GBP/CHF $s_t - f_t^N$ without an intercept.

siveness, suggesting evidence of rational bubbles during this period. Overall, we find no significant evidence of bubbles in $s_t$ although the test statistic suggests explosive bubble-like behaviors.

By comparing the left panel of Fig. 3 and right panel of Fig. 3, we obtain different date-stamping strategies for GBP/JPY using the two model specifications. Under the model specification of the null

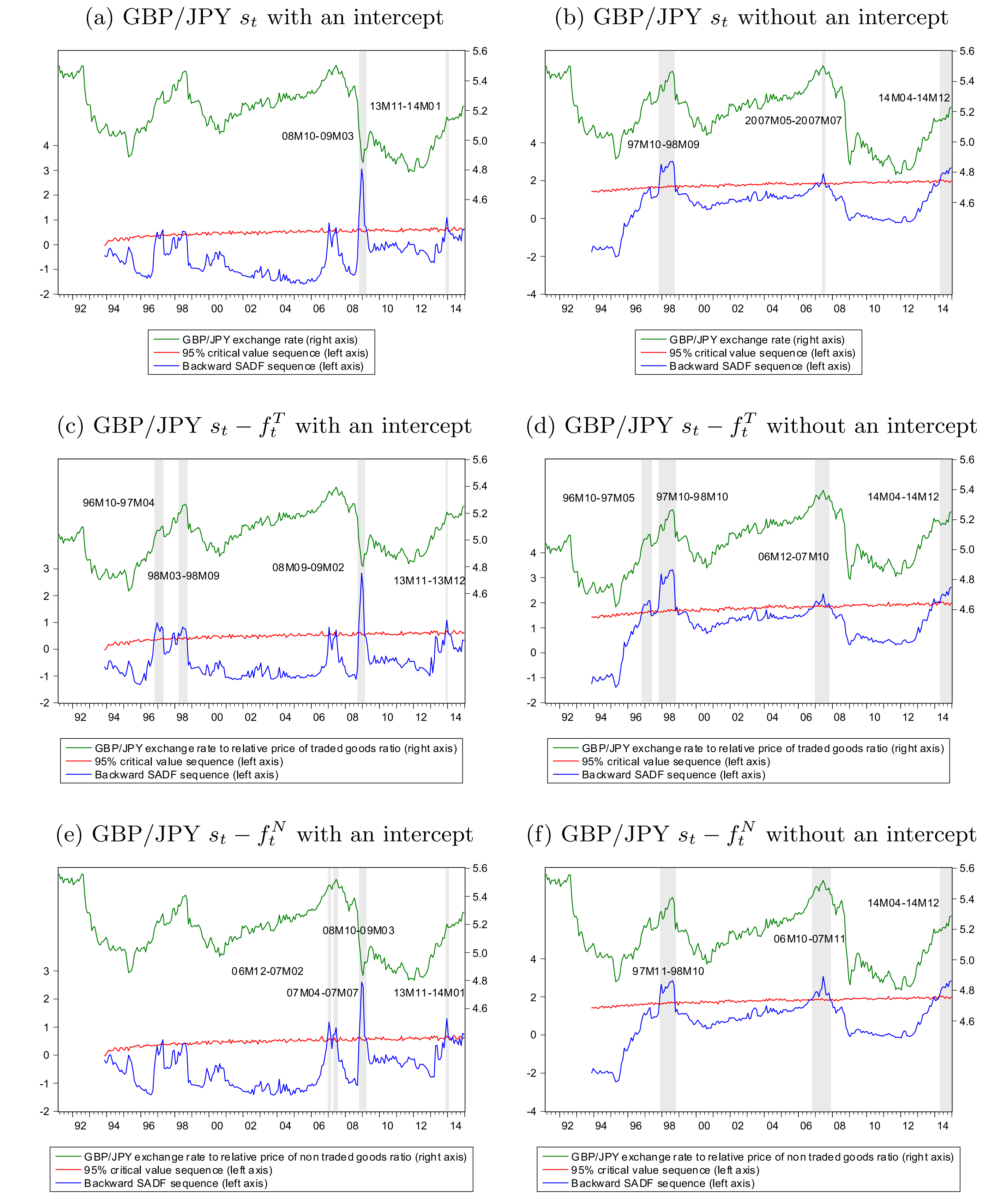


**Fig. 3.** Dating strategy for GBP/JPY nominal exchange rate $s_t$, the relative ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$ and the relative ratio of the exchange rate to the non-traded goods fundamental $s_t - f_t^N$. (a) GBP/JPY $s_t$ with an intercept. (b) GBP/JPY $s_t$ without an intercept. (c) GBP/JPY $s_t - f_t^T$ with an intercept. (d) GBP/JPY $s_t - f_t^T$ without an intercept. (e) GBP/JPY $s_t - f_t^N$ with an intercept. (f) GBP/JPY $s_t - f_t^N$ without an intercept.

hypothesis 'without an intercept', the null hypothesis of no explosive behavior cannot be rejected at the 10% significance level for $s_t$ and $s_t - f_t^N$ while the null hypothesis of no explosive behavior in $s_t - f_t^T$ is rejected at the 10% level. Three episodes have been identified from $s_t$ in Fig. 3b: 1997M10-1998M09, 2007M05-2007M07 and 2014M04-2014M12. All episodes identified from the right panel of Fig. 3

correspond to a 'genuine' bubble. The episode between 2014M04 and 2014M12 suggests that the Sterling-Japanese Yen exchange rate is experiencing a bubble. The nominal exchange rate series remains explosive after both traded and non-traded goods components are taken into account in Figs. 3d and f. We do not detect the 'collapse and recovery' type of episodes between 2008M10 and 2009M03. Our findings indicate some evidence of rational bubbles in $s_t$ as they are not explained by exchange rate fundamentals.

### 5.2. Results for Asian countries

In this section, we consider the existence of exchange rate bubbles in several Asian currencies with particular interest during the 1997 Asian Financial Crisis period. The 1997 Asian Financial Crisis originated in Thailand in July 1997 when the Thai Baht was allowed to float and soon spread to most Southeast Asian countries including Indonesia, Malaysia, the Philippines, Singapore and South Korea.

#### 5.2.1. Thai Baht (THB)

The Baht was pegged at 25 to the US Dollar between 1986 and 1995. In May 1997, a major speculative attack took place against the Baht. Due to the lack of foreign currency to defend the currency, the Thai government was forced to float against US Dollar in July 1997. The Baht depreciated to 55 to the US Dollar by the end of January of 1998 losing more than 50% of its value.

According to Table 4, the null hypothesis of no explosive behavior for USD/THB is rejected at the 1% significance level under the assumption of model specification with an intercept. As shown in Fig. 4a, there is a bubble during 1997M07-1998M02 and a 'collapse and recovery' episode in 2008. However, the explosiveness in 1997–1998 is driven by neither the relative prices of traded goods nor non-traded goods. As indicated in both Figs. 4c and e, the exchange rate $s_t$ is still explosive even if ${f_t}^T$ and ${f_t}^N$ are considered, respectively. We therefore conclude that neither the relative prices of traded goods nor non-traded goods could explain the explosive behavior during 1997–1998, which suggests the existence of rational bubbles. A 'collapse and recovery' episode in 2008 can be found in the left panel of Fig. 4, which is likely related with the Global Financial Crisis (GFC). An additional 'collapse and recovery' episode is observed during 2010 in Fig. 4c.

The right panel of Fig. 4 provides the date-stamping strategy under the model specification without an intercept. All three series ($s_t$, $s_t - f_t^T$ and $s_t - f_t^N$) are no longer explosive as the null hypothesis cannot be rejected at the 10% level. We no longer find any collapse episodes in Figs. 4b, d and f. However, we observe a bubble from 1997M09 to 1998M02 in all three figures, which is related to the Asian Financial Crisis.

#### 5.2.2. Indonesian Rupiah (IDR)

Following the collapse of the Baht, Indonesia widened the Rupiah currency trading band from 8% to 12% in July 1997. In August 1997, the managed floating exchange rate was abandoned and the Rupiah was allowed to float freely. The nominal exchange rate remained almost constant before the 1997 Asian Financial Crisis but it had some initial falls immediately after the crisis occurred. The Rupiah traded at 2600 to the US Dollar in July 1997 and it depreciated to 14900 per US Dollar in June 1998. The Indonesian Rupiah was one of the most volatile currencies during the East Asian currency crisis as it depreciated to near one-sixth of its pre-crisis level (Ito, 2007).

Under the model specification with an intercept, the null hypothesis of no explosive behavior in the Dollar-Indonesian Rupiah exchange rate is rejected at the 1% significance level as listed in Table 4. We find multiple bubbles in $s_t$ including 1994M08-1996M08, 1996M11-1998M09 and 2013M07-2014M02 from Fig. 5a. The first episode in $s_t$ is driven by the relative prices of traded goods ${f_t}^T$ as $s_t$ is no longer explosive once ${f_t}^T$ is taken into account. On the other hand, ${f_t}^T$ also contribute to explaining some explosiveness in 1998 and 2013. These results seem to suggest that the relative prices of traded goods have explained the majority of the movements in the nominal exchange rate. Additionally, a 'collapse and recovery' episode is observed in Fig. 5c between 2008M03 and 2008M08.

Bubble detection results under the model specification 'without an intercept' are provided in the right panel of Fig. 5. We find significant

**Table 4**
The GSADF test for exchange rate in emerging markets countries.

| Exchange rate | Test Stat under $H_0$ with an intercept | Episodes | Test Stat under $H_0$ without an intercept | Episodes |
|---|---|---|---|---|
| **USD/THB** | | | | |
| $s_t$ | 7.9539[a]*** | 97M07-98M02, 08M01-08M05 | 2.8066 | 97M09-98M02 |
| $s_t - f_t^N$ | 8.1865*** | 97M08-98M02, 08M02-08M04 | 2.7707 | 97M09-98M02 |
| $s_t - f_t^T$ | 4.6063*** | 95M03-95M07, 97M07-98M02<br>08M01-08M05, 10M08-10M12 | 2.4169 | 97M10-98M02 |
| **USD/IDR** | | | | |
| $s_t$ | 9.1720*** | 94M08-96M08, 96M11-98M09<br>13M07-14M02 | 15.7484*** | 93M11-98M02, 98M05-98M08<br>13M08-14M12 |
| $s_t - f_t^N$ | 11.0643*** | 95M04-98M09, 13M08-14M02 | 4.6668*** | 94M06-98M02, 98M05-98M08<br>13M09-14M12 |
| $s_t - f_t^T$ | 8.6602*** | 97M07-98M02, 08M03-08M08<br>13M08-13M09 | 2.0424 | 97M10-98M01 |
| **USD/KRW** | | | | |
| $s_t$ | 9.9778*** | 95M03-95M08, 96M12-98M02<br>08M08-08M11, 09M01-09M02 | 4.5216*** | 93M11-95M04, 96M05-98M02 |
| $s_t - f_t^N$ | 9.5177*** | 95M02-95M08, 97M01-98M03<br>04M11-05M05, 05M12-06M06<br>08M08-08M11, 09M01-09M02 | 2.3598 | 93M11-94M05, 97M02-98M02 |
| $s_t - f_t^T$ | 9.9778*** | 95M03-95M08, 97M09-98M02<br>08M08-08M11 | 2.9672 | 93M11-94M11, 97M08-97M12<br>08M09-08M11 |

[a]*** indicates significance at the 1% level.

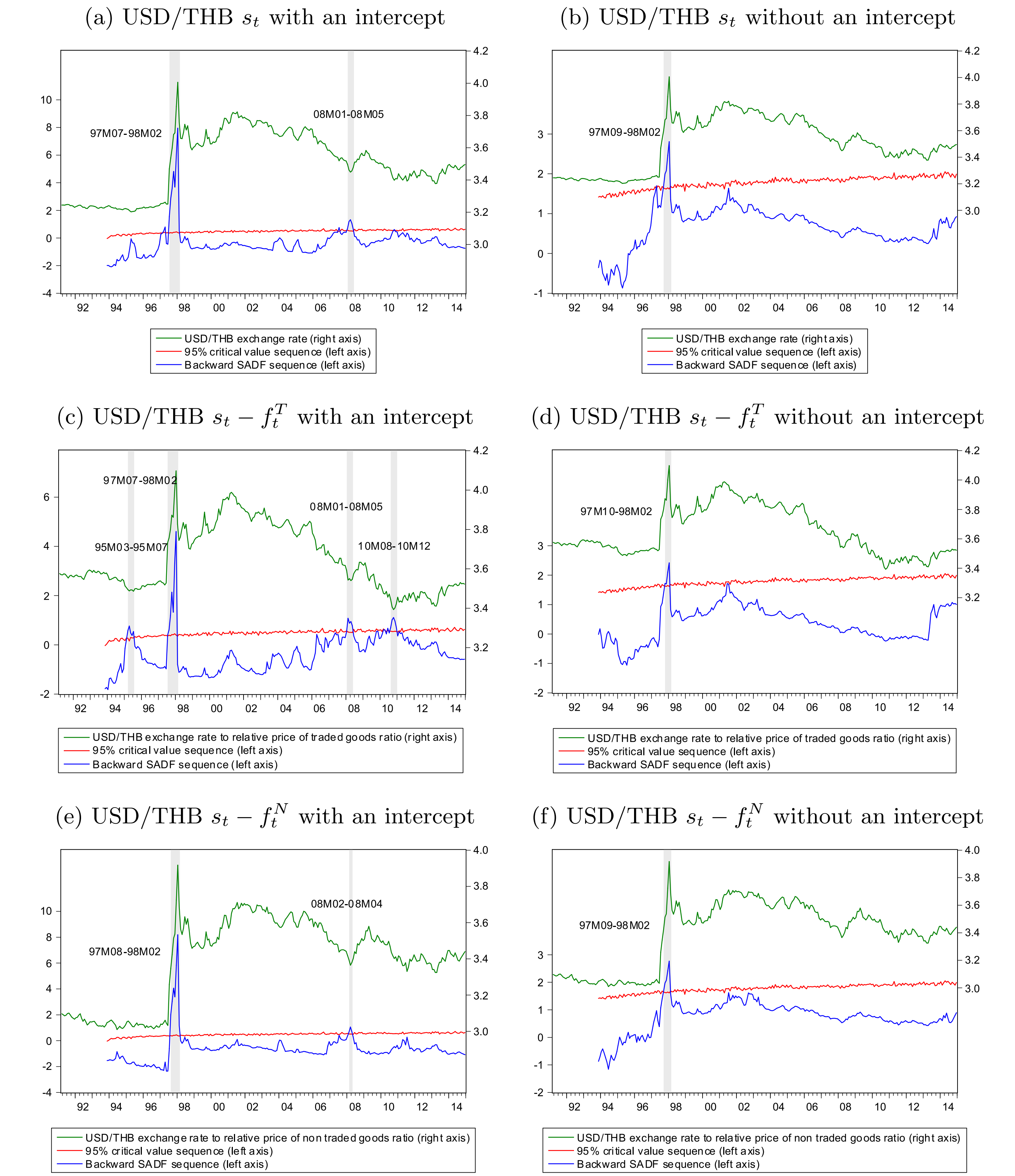


**Fig. 4.** Dating strategy for USD/THB nominal exchange rate $s_t$, the relative ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$ and the relative ratio of the exchange rate to the non-traded goods fundamental $s_t - f_t^N$. (a) USD/THB $s_t$ with an intercept. (b) USD/THB $s_t$ without an intercept. (c) USD/THB $s_t - f_t^T$ with an intercept. (d) USD/THB $s_t - f_t^T$ without an intercept. (e) USD/THB $s_t - f_t^N$ with an intercept. (f) USD/THB $s_t - f_t^N$ without an intercept.

evidence of bubbles in $s_t$ at the 1% significance level with three explosive subperiods including 1993M11-1998M02, 1998M05-1998M08 and 2013M08-2014M12 in Fig. 5b. The most recent episode (2013M08-2014M12) suggests that USD/IDR exchange rate is experiencing a bubble. The ratio of $s_t - f_t^N$ is also significant at the 1% level, which indicates strong evidence of explosive subperiods in Fig. 5f (e.g., 1994M06-1998M02, 1998M05-1998M08 and 2013M09-2014M12). On the other hand, the null hypothesis of no explosive bubbles for

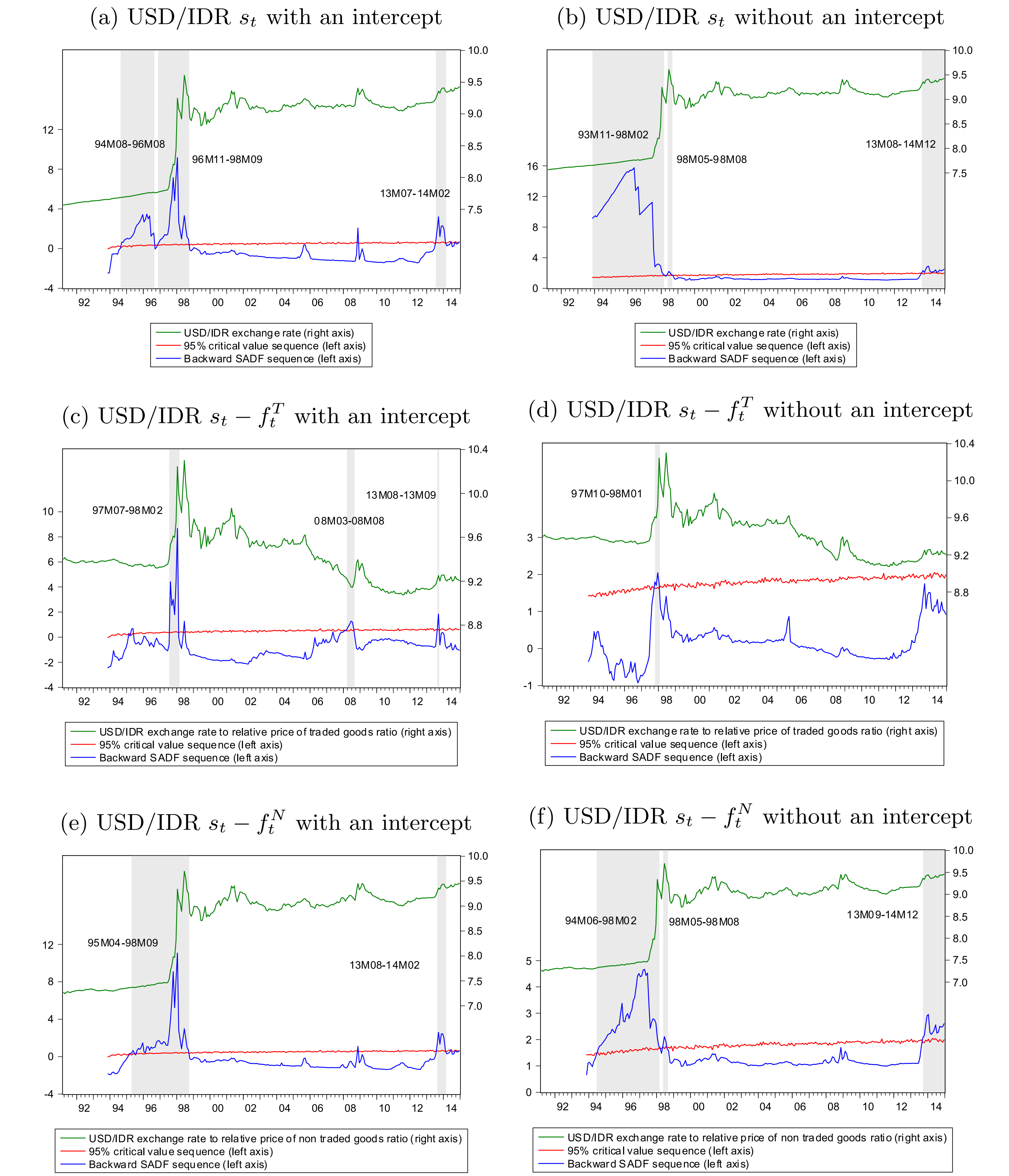


**Fig. 5.** Dating strategy for USD/IDR nominal exchange rate $s_t$, the relative ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$ and the relative ratio of the exchange rate to the non-traded goods fundamental $s_t - f_t^N$. (a) USD/IDR $s_t$ with an intercept. (b) USD/IDR $s_t$ without an intercept. (c) USD/IDR $s_t - f_t^T$ with an intercept. (d) USD/IDR $s_t - f_t^T$ without an intercept. (e) USD/IDR $s_t - f_t^N$ with an intercept. (f) USD/IDR $s_t - f_t^N$ without an intercept.

$s_t - f_t^T$ cannot be rejected at the 10% significance level. Unlike $f_t^N$, the relative prices of traded goods component $f_t^T$ plays an important role in explaining the volatility of exchange rates as suggested in Fig. 5d. Our empirical results from USD/IDR exchange rates suggest that the relative prices of traded goods $f_t^T$ have explained the majority of the movements in $s_t$, which are in line with conclusions drawn from Engel (1999) and Betts and Kehoe (2005).

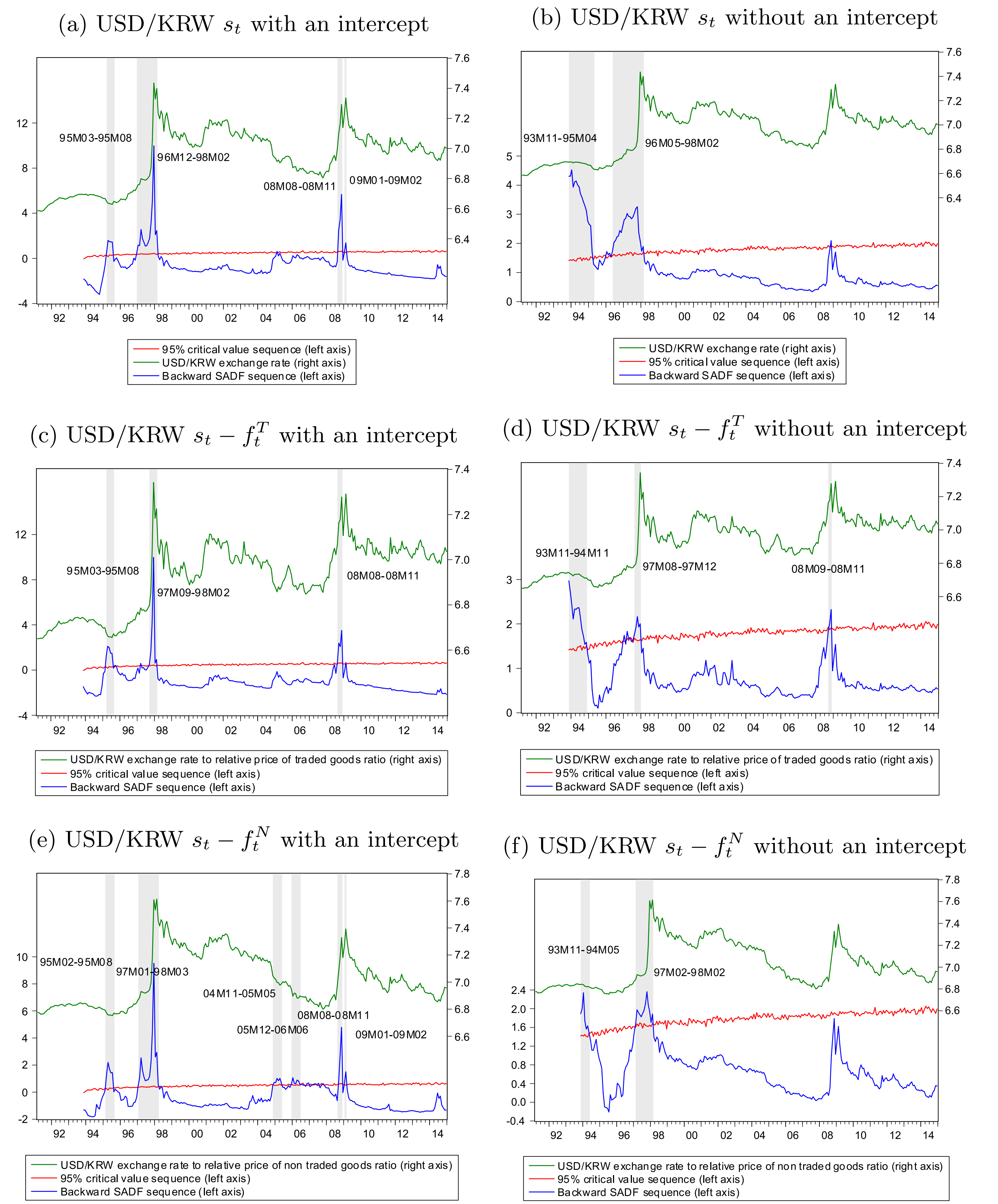


**Fig. 6.** Dating strategy for USD/KRW nominal exchange rate $s_t$, the relative ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$ and the relative ratio of the exchange rate to the non-traded goods fundamental $s_t - f_t^N$. (a) USD/KRW $s_t$ with an intercept. (b) USD/KRW $s_t$ without an intercept. (c) USD/KRW $s_t - f_t^T$ with an intercept. (d) USD/KRW $s_t - f_t^T$ without an intercept. (e) USD/KRW $s_t - f_t^N$ with an intercept. (f) USD/KRW $s_t - f_t^N$ without an intercept.

### 5.2.3. Korean Won (KWR)

The exchange rate between the Korean Won and US Dollar was one of the most affected pairs during the 1997 Asian Financial Crisis. The null hypothesis of no bubbles in USD/KWR under the model specification with an intercept is rejected for $s_t$, $s_t - f_t^T$ and $s_t - f_t^N$ at the 1% level and the corresponding bubble detection results are shown in

Table 4. Figs. 6a, c and e show the date-stamping outcomes in $s_t$, $s_t - f_t^T$ and $s_t - f_t^N$ under the model specification with an intercept, respectively. Four episodes are identified from Fig. 6a including 1995M03-1995M08, 1996M12-1998M02, 2008M08-2008M11 and 2009M01-2009M02. Firstly, a collapse and recovery episode is identified between March 1995 and August 1995 in all three series under the model with an intercept. Secondly, both $s_t$ and $s_t - f_t^N$ detect the explosiveness from the late 1996 or early 1997 to the early 1998 while $s_t - f_t^T$ suggests a bubble episode starting from September 1997 until the early of 1998. It appears that $f_t{}^T$ has partially explained the explosive behaviour from the early to mid 1997. These bubble episodes correspond to the 1997 Asian Financial Crisis where the Korean Won has depreciated sharply from the pre-crisis level of 800 per US Dollar to 1700 per US Dollar at the end of 1997. In order to avoid the worst case scenario of a sovereign default, the IMF provided a \$ 58.4 billion bailout plan to South Korea in December 1997 (Koo and Kiser, 2001). Thirdly, two more short-lived bubbles in 2008-2009 are likely related to the 2008 Global Financial Crisis. Both $f_t^T$ *and* $f_t^N$ have no effect in explaining the explosive behavior in $s_t$ in 2008 while $f_t^T$ can explain the explosiveness in early 2009. Our empirical results seem to suggest that the relative prices of traded goods $f_t^T$ play little role in explaining the exchange rate movements and the relative prices of non-traded goods $f_t^N$ contribute little in explaining the explosiveness in $s_t$.

As suggested in Table 4, the nominal exchange rate series $s_t$ remain explosive with two explosive subperiods (1993M11-1995M04 and 1996M05-1998M02) even if the intercept term is removed from the model specification under the null hypothesis. However, $s_t - f_t^T$ and $s_t - f_t^N$ series are non explosive as both series are not significant at the 10% level. Both $f_t{}^T$ and $f_t{}^N$ could not explain the majority of the explosiveness. We are more convinced by the fact that the episode between 1996M05 and 1998M02 is a bubble, which is caused by the Asian Financial Crisis. A short-lived bubble is also detected in Fig. 6f. These results are consistent with the early findings under the assumption of the inclusion of an intercept. The exclusion of an intercept for constructing the hypothesis affects the date-stamping strategy of the PSY approach.

### 5.2.4. Malaysian Ringgit (MYR)

For USD/MYR, we identify explosive behavior in $s_t$, $s_t - f_t^T$ and $s_t - f_t^N$ at the 1% level based on the model specification under the null hypothesis in Table 5. As indicated in Fig. 7a, multiple episodes can be observed in $s_t$ including 1997M08-1998M08, 2003M03-2003M06, 2006M02-2006M06 and 2006M11-2008M08. The Malaysian Ringgit traded at 2.5 US Dollar before the 1997 Asian Financial Crisis and it depreciated sharply to 3.8 US Dollars by the end of 1997. There is a bubble episode between August 1997 and August 1998 in both Fig. 7a and e while a shorter bubble episode from August 1997 to February 1998 is detected in Fig. 7c. Such a bubble corresponds to the 1997 Asian Financial Crisis. The relative prices of traded goods $f_t{}^T$ have partially explained the explosiveness in $s_t$ while such an explosive behavior is not driven by the relative prices of non-traded goods $f_t{}^N$.

It is perhaps noteworthy to compare findings from the GSADF test using the two model specifications. First, we find a spurious episode in 2003 for the nominal exchange rate $s_t$ in Fig. 7a. The Malaysian Ringgit was pegged to the US Dollar in September 1998 keeping the exchange rate around 3.8 per US Dollar until the end of 2005. Thus we would not expect any explosive behavior during this seven-year period. However, as shown in Fig. 7a, there is a spurious episode dated from March 2003 to June 2003 in the series. We could not explain the reason behind this 'collapse' episode. Second, we notice two 'collapse and recovery' episodes (2006M02-2006M06 and 2006M11-2008M08) in Fig. 7a for $s_t$. This spurious 'collapse' episode in 2003 and two 'collapse and recovery' episodes (2006M02-2006M06 and 2006M11-2008M08) are likely caused by the inclusion of an intercept in the model specification under the null hypothesis as seen by comparing Fig. 7a and b. Overall, under the assumption 'with an intercept', the PSY approach could lead to the false positive identification of bubbles as it cannot distinguish between 'collapse' type of episodes and bubbles.

However, we obtain different results if the intercept is excluded in the model formulation. The null hypothesis of no bubbles under model specification 'without an intercept' for $s_t$ and $s_t - f_t^N$ are rejected at the 5% significance level, which indicates strong evidence of bubbles. We find two explosive episodes (1997M09-1998M02 and 1998M05-1998M08) from $s_t$ in Fig. 7b and $s_t - f_t^N$ in Fig. 7f. The test statistics for $s_t - f_t^T$ is slightly lower than the 10% significance level. As exchange rate fundamentals ($f_t{}^T$ and $f_t{}^N$) could not explain the bubble in 1997–1998, we may conclude the evidence of rational bubbles. When the intercept term is removed from the model specification for the null hypothesis, the backward SADF statistic sequences and 95% critical value sequences do not "detect" the 'collapse' episode in 2003 and 'collapse and recovery' episodes any longer in the right panel of Fig. 7.

### 5.2.5. Philippine Peso (PHP)

Table 5 suggests that the null hypothesis of no explosive behavior in the nominal USD/PHP exchange rate $s_t$ is rejected at the 1% significance level based on the GSADF test. As shown in Fig. 8a, there is a bubble during 1997M08-1998M10 and a 'collapse and recovery' episode during 2006M12-2008M05 for $s_t$. The first explosive bubble is clearly related to the 1997 Asian Financial Crisis. It seems that $f_t{}^N$ could not explain this explosiveness while $f_t{}^T$ explains some movements in exchange rates. As can be seen in Fig. 8c, we find no evidence of explosiveness in the $s_t - f_t^T$ series for the second explosive period in 2007–2008, which is likely associated with the 2008 Global Financial Crisis. According to Fig. 8e, the exchange rate still remains explosive after the relative prices of non-traded goods are taken into account although the time duration of the explosive behaviour in the $s_t - f_t^N$ series is shorter than those from the $s_t$ series. On the other hand, we also observe three additional bubble periods from the $s_t - f_t^N$ series. Overall, the above results seem to suggest that the relative prices of traded goods play a crucial role in explaining the explosiveness in the nominal US Dollar-Philippine Peso exchange rate.

The exclusion of the intercept term for model formulation of hypothesis yields quite different results as indicated in the right panel of Fig. 8. The null hypothesis of no explosive behavior for $s_t$ and $s_t - f_t^T$ are not rejected at the 10% significance level while the hypothesis for $s_t - f_t^N$ is rejected at the 5%. The episode in 1997-1998 is identified in all three series ($s_t$, $s_t - f_N^T$ and $s_t - f_t^T$). There are two long-lasting episodes in $s_t$ (1999M07-2007M02) and $s_t - f_N^T$ (2000M03-2007M09) in Fig. 8b and f, respectively and these results are not expected and may be spurious. These two episodes are not detected under the model specification 'with an intercept'. It seems that the relative prices of traded goods $f_t{}^T$ explain the majority of exchange rate movements.

### 5.2.6. Singapore Dollar (SGD)

Unlike most Asian currencies, a managed floating exchange rate regime was adopted by the Singapore government in 1973 (Lu and Yu, 1999). In 1967, the Board of Commissioners of Currency of Singapore (BCCS) was established to issue currency. The Monetary Authority of Singapore (MAS) established in 1971 manages the Singapore Dollar against a trade-weighed basket of currencies. The Board of Commissioners of Currency of Singapore merged with the Monetary Authority of Singapore in October 2002.

As can be seen from Table 5, under the assumption 'with an intercept', we find strong evidence of explosive behaviour in all three series for USD/SGD at the 1% significance level. As shown in Fig. 9a, a bubble episode between 1997M09 and 1998M02 as well as several 'collapse and recovery' episodes (e.g., 1994M07-1995M08, 2007M09-2008M08 and 2011M01-2011M09) are observed in $s_t$. The bubble episode during 1997-1998 is associated with the 1997 Asian Financial Crisis. Neither $f_N{}^T$ or $f_t{}^T$ could explain the explosiveness during the Asian financial downturn, suggesting evidence of rational bubbles. More 'collapse' episodes have been found in Fig. 9e (e.g., 1994M07-

**Table 5**
The GSADF test for exchange rate in emerging markets countries.

| Exchange rate | Test Stat under $H_0$ with an intercept | Episodes | Test Stat under $H_0$ without an intercept | Episodes |
|---|---|---|---|---|
| **USD/MYR** | | | | |
| $s_t$ | 6.8802[a]*** | 97M08-98M08, 03M03-03M06<br>06M02-06M06, 06M11-08M08 | 3.3746** | 97M09-98M02, 98M05-98M08 |
| $s_t - f_t^N$ | 8.3895*** | 97M08-98M09 | 3.4557[b]** | 97M09-98M02, 98M05-98M08 |
| $s_t - f_t^T$ | 4.4348*** | 97M08-98M02, 07M12-08M05 | 2.9921 | 97M09-98M02 |
| **USD/PHP** | | | | |
| $s_t$ | 5.8052*** | 97M08-98M10, 06M12-08M05 | 2.8246 | 97M05-99M01, 99M07-07M02 |
| $s_t - f_t^N$ | 5.1539*** | 97M08-98M10, 00M07-02M03<br>07M10-08M07, 11M03-11M09<br>12M07-13M06 | 3.5298** | 97M09-98M03, 98M05-98M10<br>00M03-07M09 |
| $s_t - f_t^T$ | 3.3214*** | 97M08-98M02 | 2.2802 | 97M08-98M02, 14M01-14M11 |
| **USD/SGD** | | | | |
| $s_t$ | 4.7261*** | 94M07-95M08, 97M09-98M02<br>07M09-08M08, 11M01-11M09 | 3.1190 | 97M11-98M02 |
| $s_t - f_t^N$ | 3.7030*** | 94M07-95M11, 97M10-98M01<br>08M02-08M04, 08M11-09M01<br>10M08-11M09 | 2.5260 | 97M12-98M02 |
| $s_t - f_t^T$ | 3.0141*** | 97M07-98M02, 98M05-98M09 | 3.2448[c]* | 97M08-98M02, 98M05-98M09<br>14M10-14M12 |

[a]*** indicates significance at the 1% level.
[b]** indicates significance at the 5% level.
[c]* indicates significance at the 10% level.

1995M11, 2008M02-2008M04 and 2010M08-2011M09). Overall, we find significant evidence of bubbles during the 1997 Asian Financial Crisis.

It seems that the exclusion of the intercept for constructing the null hypothesis has affected the limit theory of the PSY approach. We obtain quite different results in the two model specifications. When the intercept is removed in the model specification of null hypothesis, $s_t$ and $s_t - f_t^N$ series are no longer explosive and the test statistics are lower than the 10% significance level. On the other hand, $s_t - f_t^T$ remains explosive at the 5% significance level. These results seem to suggest that there is little evidence of bubbles. The episode in 1997-1998 is explosive in Figs. 9b, d and f but it is short-lived. Once the intercept is removed, we no longer find 'collapse and recovery' type of episodes. Moreover, ${f_t}^T$ does not play an important role in explaining the majority of the movements in $s_t$.

### 5.3. Results for BRICS countries

We also look for evidence of explosive behavior in BRICS currencies including the Brazilian Real (BRL), Indian Rupee (INR) and South African Rand (ZAR) measured against the US Dollar[6]. Maldonado et al. (2016) adopt the cointegration approach and detect the presence of bubbles in BRICS economies' currencies against the US Dollar using data from 1999M03 to 2013M06. It would be interesting to make a comparison between our results and those of Maldonado et al. (2016).

#### 5.3.1. Brazilian Real (BRL)

The Brazilian Real was pegged to 1 US Dollar when it was initially introduced in July 1994. The Real appreciated against the US Dollar in the early years, but from July 1996, the Real depreciated against the US Dollar. By the end of 1998, the Real depreciated slowly against the US Dollar at a rate of 1:1.2. The Real was allowed to fluctuate within a narrow trading band until early 1999 such that its value was closely controlled by the government (Gruben and Welch, 2001). The adoption of the pre-set band provides some flexibility of the exchange rate, aimed at resolving the inflation problem. The Real was floated in January 1999 as the government was unable to hold the peg (Ferreira and Tullio, 2002). As a result, the Real further devalued to a rate of 1:2.

Based on Table 6, the null hypothesis of explosive behavior is rejected at the 5% significance level for the nominal USD/BRL exchange rate $s_t$. The first bubble period between June 1997 and March 1999 in Fig. 10a is associated with the devaluation of the Real. According to Ferreira and Tullio (2002), the price index for non-traded goods increased by 120 per cent, and the price index for traded goods increased by about 27 per cent between July 1994 and the end of 1998. Several short bubble episodes can be seen in Fig. 10a (e.g., 2001M07-2001M10, 2002M06-2002M07, 2002M09-2002M10) along with a 'collapse' episode during 2005M08-2005M11. We then investigate whether the explosiveness in $s_t$ is driven by rational bubbles or exchange rate fundamentals. According to Fig. 10c, $s_t - f_t^T$ suggests no evidence of rational bubbles as the ratio is no longer explosive. Thus ${f_t}^T$ plays a vital role in explaining the volatility of the nominal exchange rate. On the other hand, the ratio of $s_t - f_t^T$ is explosive as shown in Fig. 10e. Hence, ${f_t}^N$ has little contributions in explaining the explosiveness.

When the intercept is not used for constructing the hypothesis, $s_t$ and $s_t - f_t^N$ are still significant at the 1% level while the null hypothesis of no explosive bubbles in $s_t - f_t^T$ cannot be rejected at the 10% level. We find evidence of multiple bubbles in Fig. 10b (e.g., 1997M12-1999M02, 2001M08-2001M11 and 2002M05-2002M10). We cannot detect those 'collapse and recovery' episodes any more in the right panel of Fig. 10. Interestingly, there is a bubble episode between 2001M07 and 2003M03 in Fig. 10f, which is not identified before. It seems that ${f_t}^T$ has explained most movements in the exchange rate for both model formulations. Maldonado et al. (2016) report a lager

[6] Due to the lack of the PPI data for Russia, we could not test for the explosive behavior in the US Dollar-Russian Ruble exchange rate fundamentals. Jiang et al. (2015) investigated the explosive behavior in the Chinese RMB-US Dollar exchange rate. We therefore only include the three remaining countries in our analysis.

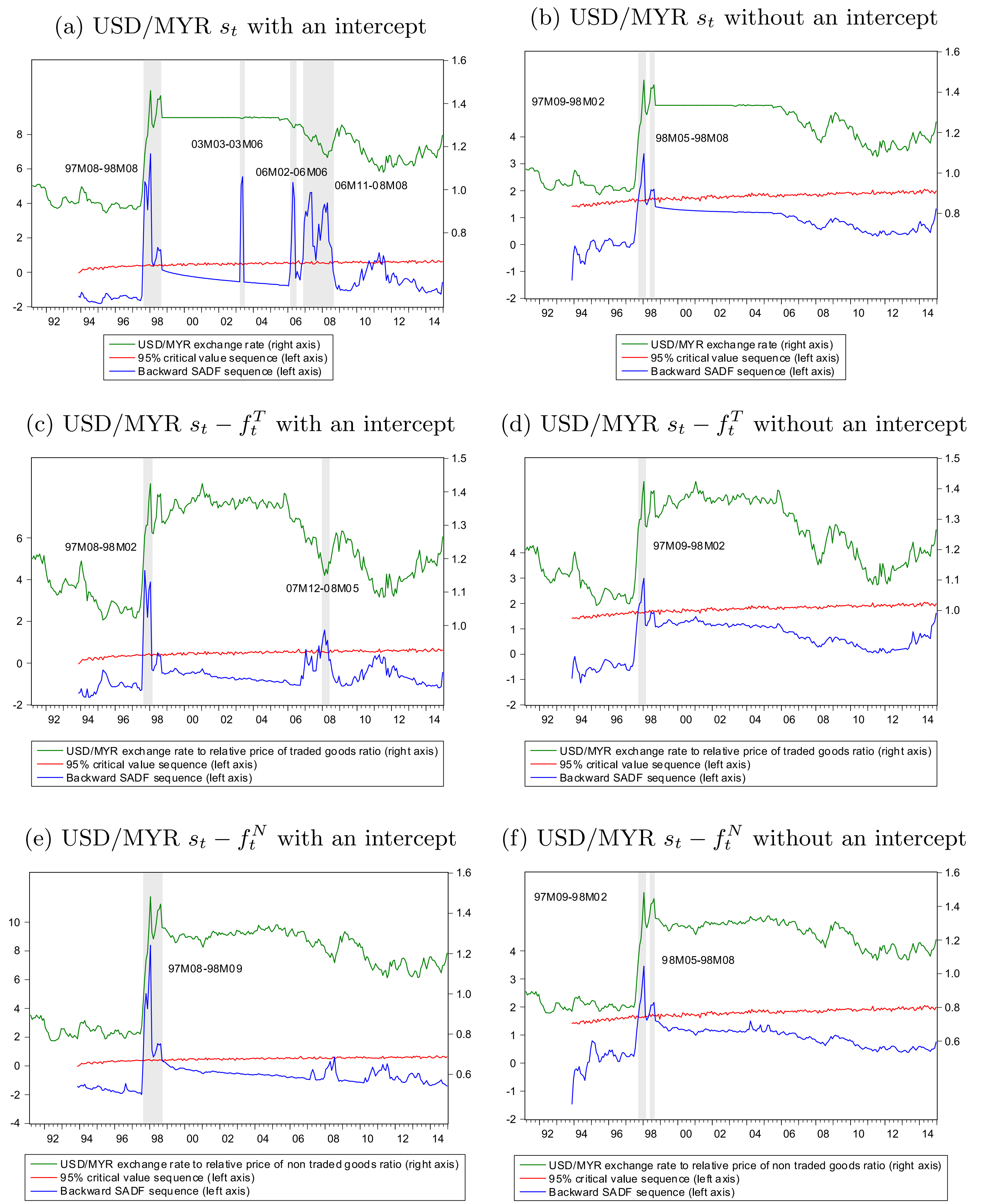


**Fig. 7.** Dating strategy for USD/MYR nominal exchange rate $s_t$, the relative ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$ and the relative ratio of the exchange rate to the non-traded goods fundamental $s_t - f_t^N$. (a) USD/MYR $s_t$ with an intercept. (b) USD/MYR $s_t$ without an intercept. (c) USD/MYR $s_t - f_t^T$ with an intercept. (d) USD/MYR $s_t - f_t^T$ without an intercept. (e) USD/MYR $s_t - f_t^N$ with an intercept. (f) USD/MYR $s_t - f_t^N$ without an intercept.

bubble from March 1999 to July 2007 and a short-lived bubble in 2008 for Brazil. However, our results for USD/BRL under both model formulations do not provide similar outcomes.

### *5.3.2. Indian Rupee (INR)*

Results for the nominal US Dollar-India Rupee exchange rate are presented in Table 6. The GSADF test suggests strong evidence of bubbles in $s_t$ as the null of no explosive behavior is rejected at the 1%

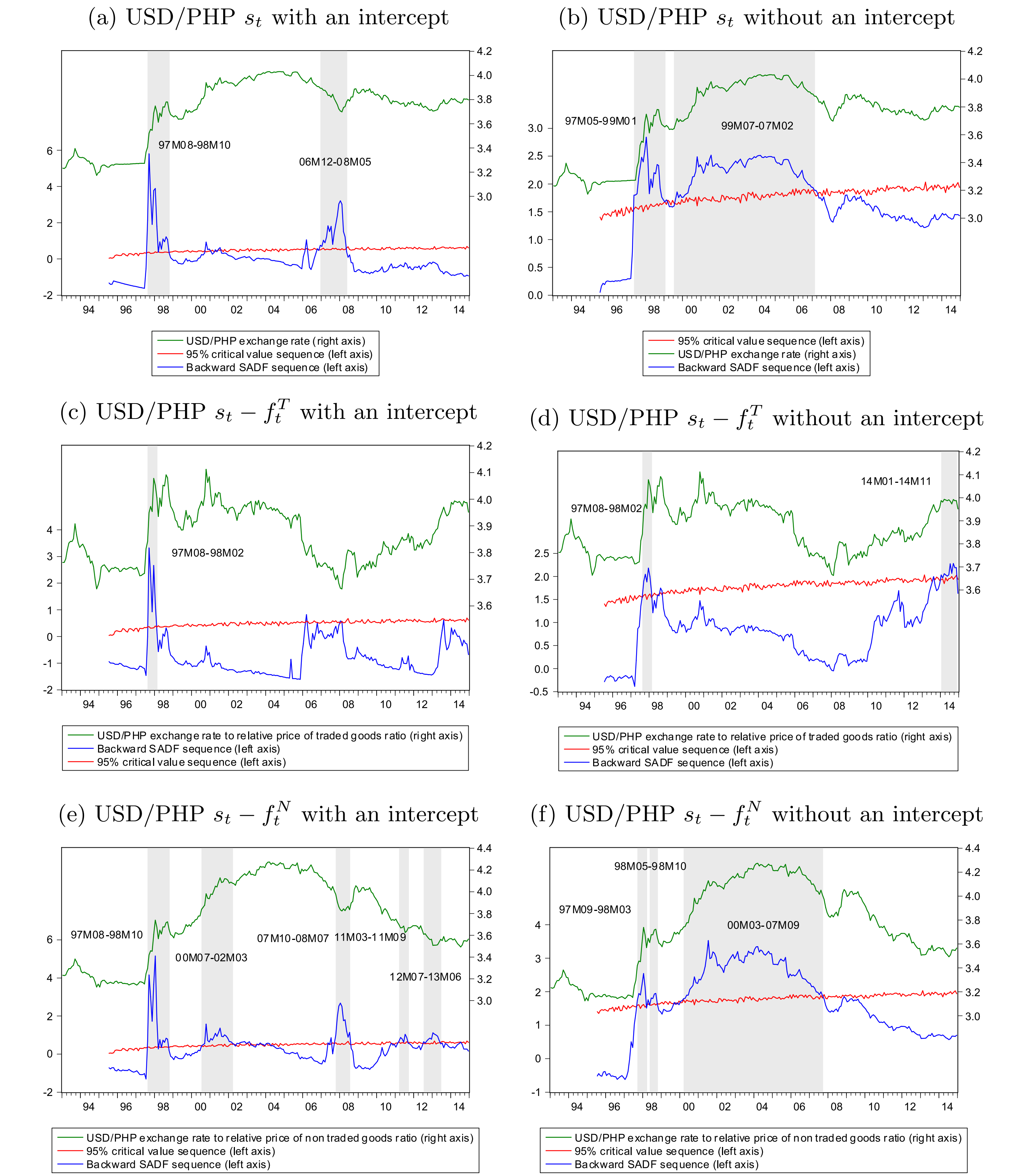


**Fig. 8.** Dating strategy for USD/PHP nominal exchange rate $s_t$, the relative ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$ and the relative ratio of the exchange rate to the non-traded goods fundamental $s_t - f_t^N$. (a) USD/PHP $s_t$ with an intercept. (b) USD/PHP $s_t$ without an intercept. (c) USD/PHP $s_t - f_t^T$ with an intercept. (d) USD/PHP $s_t - f_t^T$ without an intercept. (e) USD/PHP $s_t - f_t^N$ with an intercept. (f) USD/PHP $s_t - f_t^N$ without an intercept.

significance level. Fig. 11a shows the date-stamping results for $s_t$ and displays multiple periods of explosiveness including 1995M11-1996M02, 1998M03-1999M02, 2001M09-2002M05 and 2004M01-2004M04. The nominal exchange rate $s_t$ is no longer explosive in Fig. 11c once the relative prices of traded goods are accounted for. We find no episodes in Fig. 11c as the relative prices of traded goods explain the explosiveness in $s_t$. A ‘collapse and recovery’ episode between 2007M05 and 2008M04 is identified in Fig. 11e.

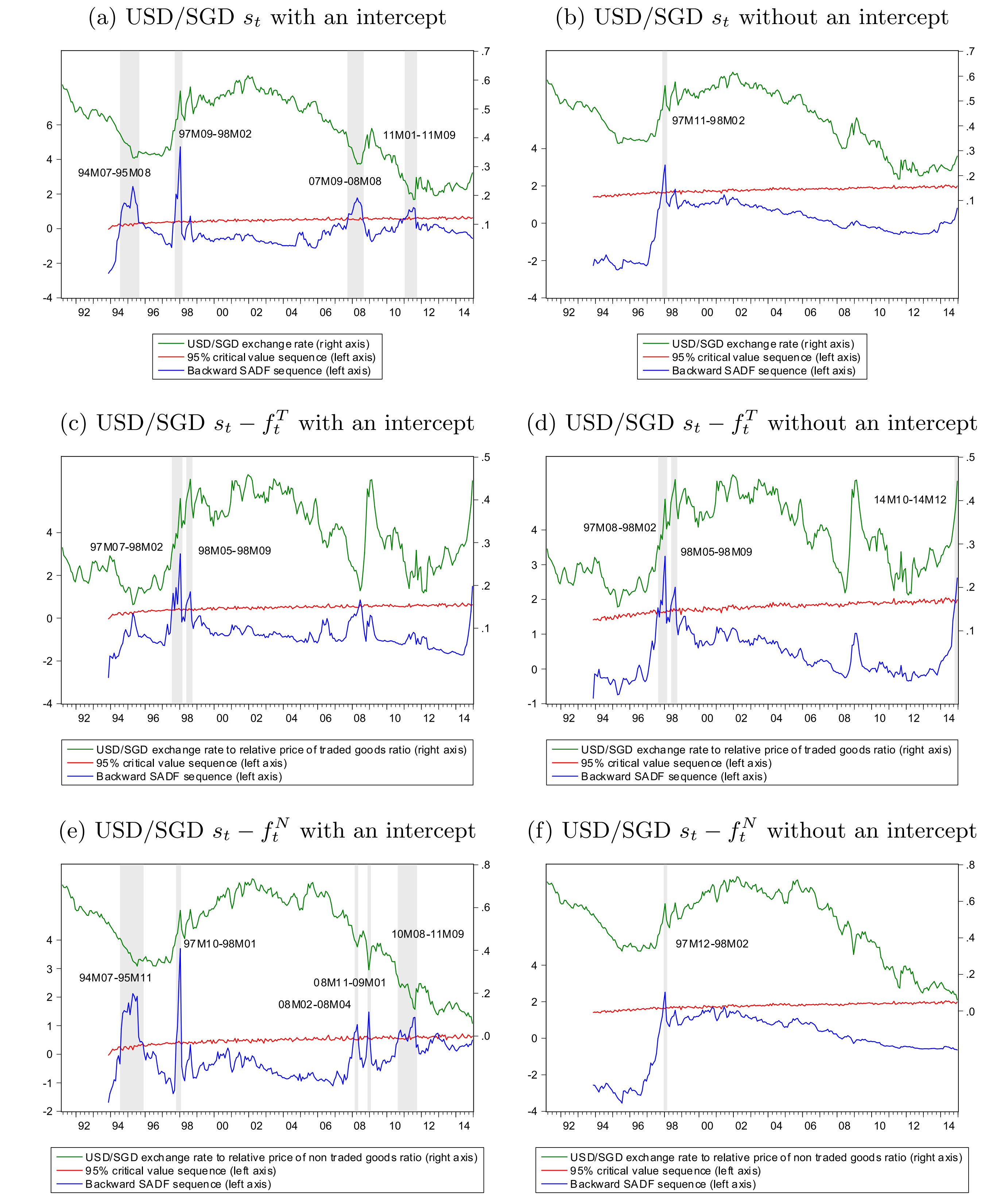


**Fig. 9.** Dating strategy for USD/SGD nominal exchange rate $s_t$, the relative ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$ and the relative ratio of the exchange rate to the non-traded goods fundamental $s_t - f_t^N$. (a) USD/SGD $s_t$ with an intercept. (b) USD/SGD $s_t$ without an intercept. (c) USD/SGD $s_t - f_t^T$ with an intercept. (d) USD/SGD $s_t - f_t^T$ without an intercept. (e) USD/SGD $s_t - f_t^N$ with an intercept. (f) USD/SGD $s_t - f_t^N$ without an intercept.

The date-stamping results for the model specification under the assumption of no intercept is quite different as shown in Figs. 11b, d and f. In Fig. 11b, we find a spurious bubble episode in $s_t$ from December 1993 to December 2014 and we do not expect such a long-lasting bubble. Similarly, a long-lasting episode between December 1993 and February 2007 is detected in $s_t - f_t^N$ of Fig. 11f. Similarly,

**Table 6**
The GSADF test for exchange rate in emerging markets countries.

| Exchange rate | Test Stat under $H_0$ with an intercept | Episodes | Test Stat under $H_0$ without an intercept | Episodes |
|---|---|---|---|---|
| **USD/BRL** | | | | |
| $s_t$ | 2.2281[a**] | 97M06-99M03, 01M07-01M10 02M06-02M07, 02M09-02M10 05M08-05M11 | 10.1813[b***] | 97M12-99M02, 01M08-01M11 02M05-02M10 |
| $s_t - f_t^N$ | 2.7464*** | 97M07-99M03, 99M08-99M12 01M04-01M12, 02M05-03M03 05M08-06M04 | 4.4563*** | 01M07-03M03 |
| $s_t - f_t^T$ | 0.8156 | | 1.8511 | 98M08-98M12 |
| **USD/INR** | | | | |
| $s_t$ | 2.7861*** | 95M11-96M02, 98M03-99M02 01M09-02M05, 04M01-04M04 | 4.0151** | 93M12-14M12 |
| $s_t - f_t^N$ | 1.3143 | 98M04-98M07, 07M05-08M04 | 3.1064 | 93M12-07M02 |
| $s_t - f_t^T$ | 0.7890 | | 1.9111 | |
| **USD/ZAR** | | | | |
| $s_t$ | 3.7159*** | 94M01-94M08, 96M03-97M01 98M04-98M10, 98M12-99M04 00M08-02M09 | 4.8427*** | 93M11-03M09 |
| $s_t - f_t^N$ | 4.9297*** | 94M02-94M08, 96M03-97M02 97M09-99M08, 00M08-02M11 | 5.0760*** | 93M11-03M09 |
| $s_t - f_t^T$ | 2.1865** | 98M06-98M08, 00M10-01M04 01M09-02M03 | 2.8881 | 96M03-96M12, 98M05-98M09 00M04-02M04 |

[a**] indicates significance at the 5% level.
[b***] indicates significance at the 1% level.

Maldonado et al. (2016) also identify a bubble in the US Dollar-India Rupee exchange rate from March 1999 to early 2007. Our results from $s_t - f_t^N$ overlap with those identified in Maldonado et al. (2016). Although the GSADF test statistic for $s_t$ and $s_t - f_t^T$ suggest evidence of bubbles, these results are spurious and we hardly believe the existence of genuine bubbles. These results demonstrate the importance of model specification in right-tailed unit root tests. When the intercept is excluded in the model formulation for constructing the null hypothesis, we could obtain some spurious and unexpected results (i.e., a spurious long-lasting episode). Thus it is important to assess a wide range of model specifications in the null.

#### *5.3.3. South African Rand (ZAR)*

We find strong evidence of bubbles in the nominal USD/ZAR exchange rate $s_t$ as shown in Table 6 as the null of no bubbles is rejected at the 1% significance level. Multiple bubbles periods are identified in Fig. 12a including 1994M01-1994M08, 1996M03-1997M01, 1998M04-1998M10, 1998M12-1999M04 and 2000M08-2002M09. According to Figs. 12c and e, the relative prices of traded goods $f_t^T$ have explained the majority of the movements in the nominal exchange rate. As both the relative prices of traded goods fundamentals and non-traded goods fundamentals cannot explain all the explosiveness in the nominal exchange rate, we therefore conclude the evidence of rational bubbles.

Comparing the left panel and right panel of Fig. 12, we obtain very different date-stamping results. Both $s_t$ and $s_t - f_t^N$ series remain explosive at the 1% significance level. However, $s_t - f_t^T$ is no longer explosive as $f_t^T$ could explain some explosiveness in $s_t$. More importantly, we find a long-lasting bubble episode from 1993M11 to 2003M09 in both $s_t$ and $s_t - f_t^N$ series and this episode is spurious. Maldonado et al. (2016) identify two bubbles in the US Dollar–South African Rand exchange rate. The first one is from March 1999 to early 2003 and the second one originates and collapses in 2009. Our bubble detection results from $s_t$ in Fig. 12b and $s_t - f_t^N$ in Fig. 12f overlap with the first bubble identified from Maldonado et al. (2016). However, our empirical results are not in line with those in Jirasakuldech et al. (2006), who find no evidence of bubbles in the US Dollar-South African Rand exchange rate between January 1989 and December 2004. Our results based on two model formulations indicate that the intercept term has greatly affected the asymptotic theory and the date-stamping strategy of the PSY approach. As discussed before, without considering the intercept in the null, the PSY approach no longer identifies 'collapse' episodes and 'collapse and recovery' episodes but this example shows that it could lead to spurious bubbles.

### *5.4. Results for other emerging markets countries*

In this section, we test for the existence of exchange rate bubbles in the US Dollar against Colombian Peso and Mexican Peso and the corresponding bubble detection results are provided in Table 7. The collapse of the Mexican Peso in 1994–95 was widely regarded as one of the exchange rate crises in the 20th century. Colombia has also experienced a banking crisis in late 1990s and followed by a currency crisis.

#### *5.4.1. Colombian Peso (COP)*

As shown in Table 7, the null hypothesis of no bubbles in the USD-COP exchange rate $s_t$ is rejected at the 10% level[7]. Fig. 13a illustrates two episodes (1997M09-2001M10 and 2002M07-2003M04). The first episode between the late 1990 s and early 2000 s is likely related with

[7] We let $r_0$=0.15 for the following analysis. If we let $r_0 = 0.01 + 1.8/\sqrt{T}$ and $T$ is 286, $r_0$ is approximately to 12%. We find that $r_0$ is not larger enough for initial estimation and therefore consider a larger $r_0$.

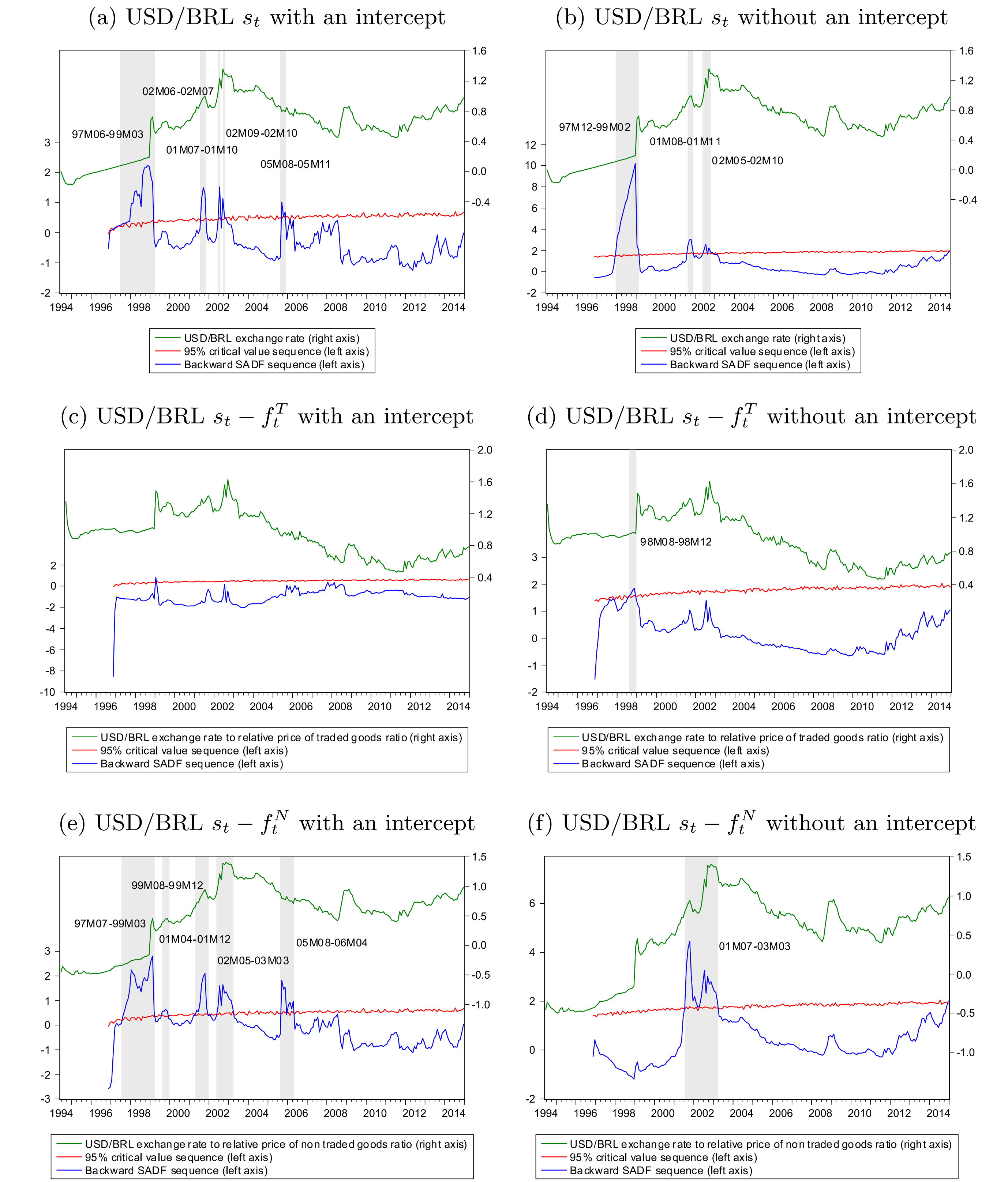


**Fig. 10.** Dating strategy for USD/BRL nominal exchange rate $s_t$, the relative ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$ and the relative ratio of the exchange rate to the non-traded goods fundamental $s_t - f_t^N$. (a) USD/BRL $s_t$ with an intercept. (b) USD/BRL $s_t$ without an intercept. (c) USD/BRL $s_t - f_t^T$ with an intercept. (d) USD/BRL $s_t - f_t^T$ without an intercept. (e) USD/BRL $s_t - f_t^N$ with an intercept. (f) USD/BRL $s_t - f_t^N$ without an intercept.

the Colombian Banking Crisis, see Gomez-Gonzalez and Kiefer (2009). The Colombian Banking Crisis during the late 1990s is also accompanied by a currency crisis, and the exchange rate regime is abandoned and is allowed to float freely in 1999 (Arias, 2000). $s_t - f_t^T$ is no longer explosive in Fig. 13c. On the contrary, the relative prices of non-traded goods fundamentals play little role in explaining the explosiveness of exchange rates as $s_t - f_t^N$ is still explosive. In addition, we spot two 'collapse' episodes in Fig. 13e (e.g., 2007M04-2007M07 and 2008M01-

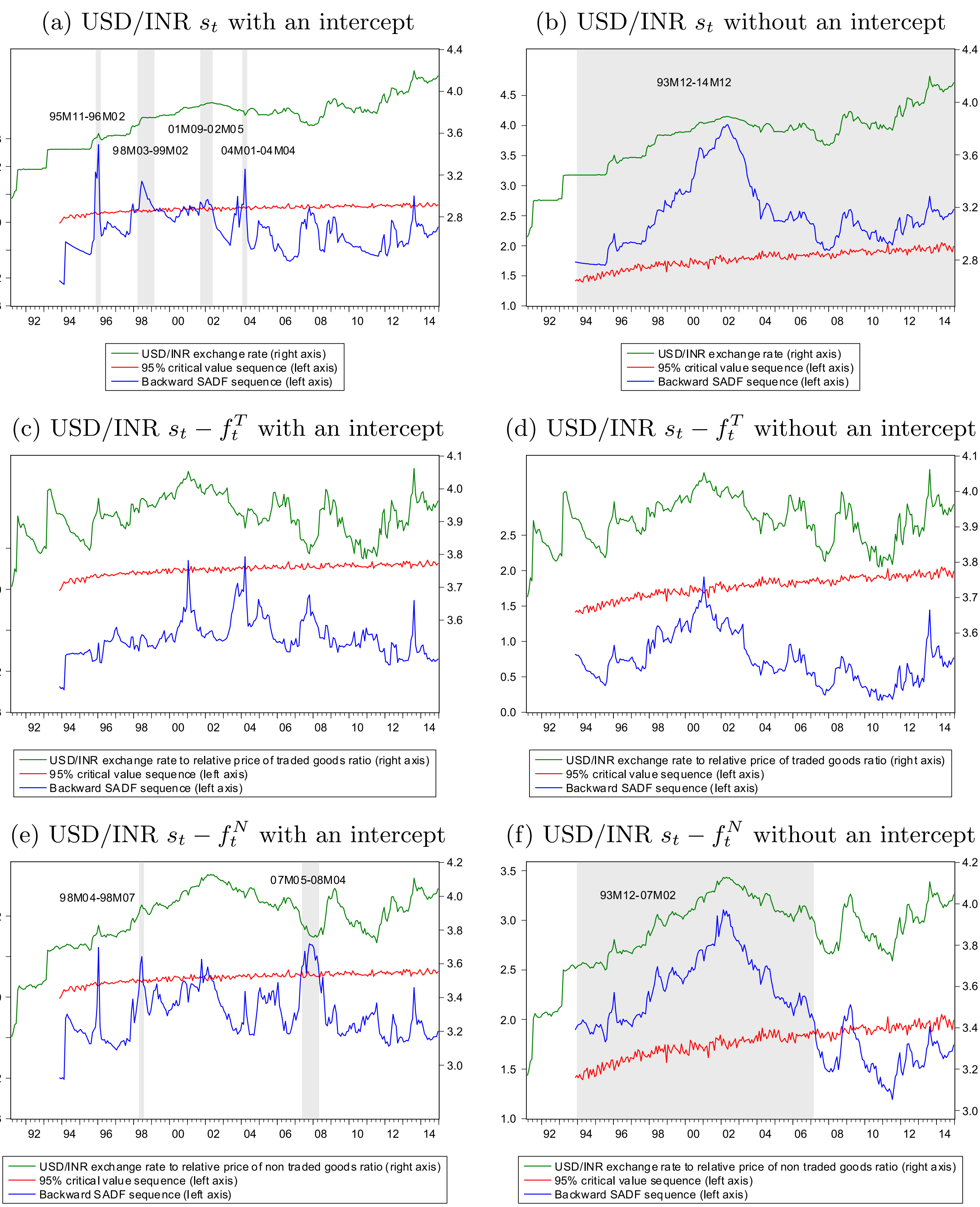


**Fig. 11.** Dating strategy for USD/INR nominal exchange rate $s_t$, the relative ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$ and the relative ratio of the exchange rate to the non-traded goods fundamental $s_t - f_t^N$. (a) USD/INR $s_t$ with an intercept. (b) USD/INR $s_t$ without an intercept. (c) USD/INR $s_t - f_t^T$ with an intercept. (d) USD/INR $s_t - f_t^T$ without an intercept. (e) USD/INR $s_t - f_t^N$ with an intercept. (f) USD/INR $s_t - f_t^N$ without an intercept.

2008M08).

Model formulation in the null hypothesis seems to have an impact on the PSY approach as detailed in Figs. 13b, d and f. The PSY approach detects two long-lasting episodes in Fig. 13b (1994M08-2014M12) and Fig. 13f (1995M06-2008M02) and these results are not expected and spurious. Thus the rejection of no bubbles in the null hypothesis under the assumption 'without an intercept' in the PSY could lead to some spurious episodes. Even if the GSADF test statistic for $s_t$ and $s_t - f_t^N$ indicate evidence of bubbles, we hardly believe the presence of genuine bubbles on a close inspection of the actual exchange rate series.

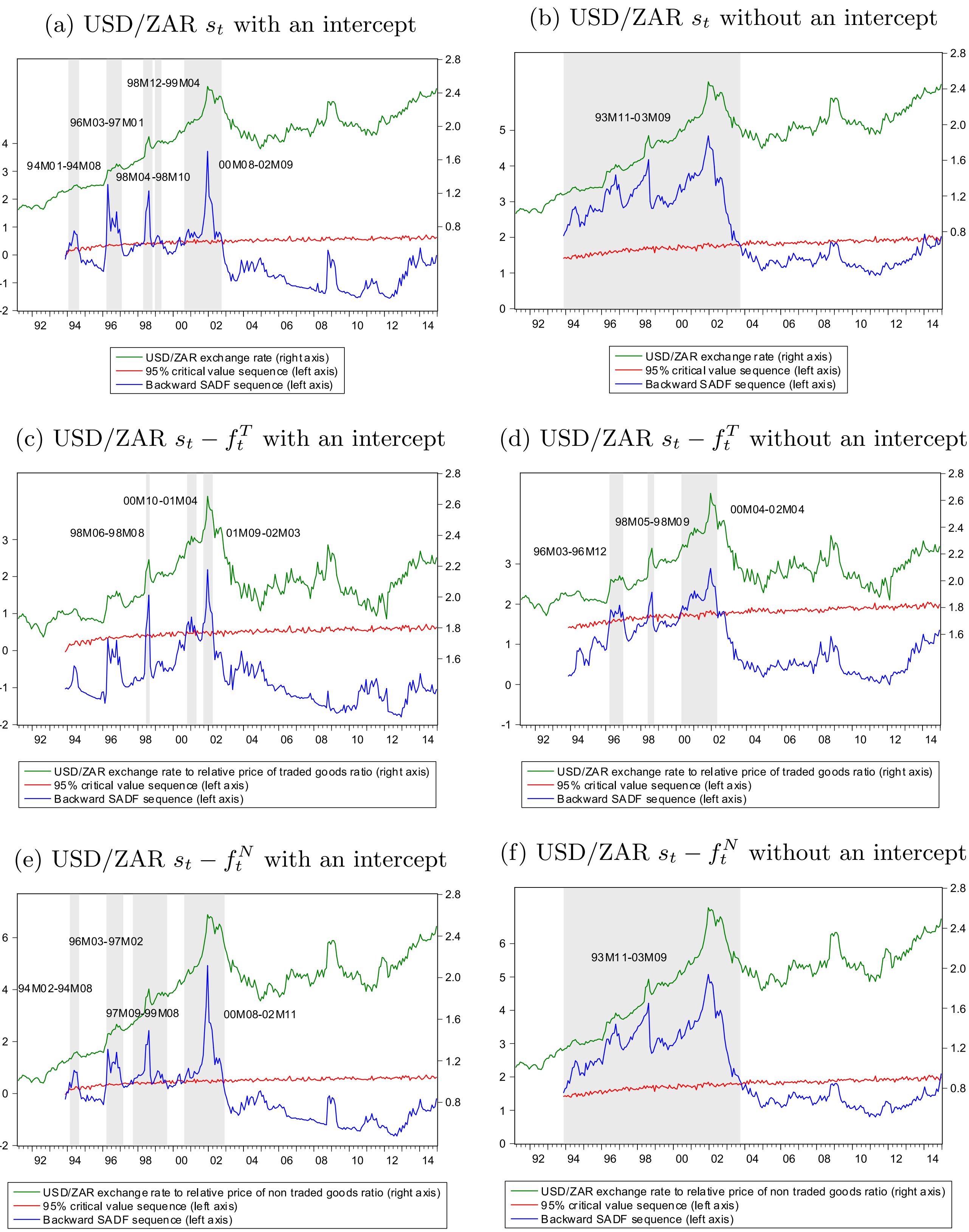


**Fig. 12.** Dating strategy for USD/ZAR nominal exchange rate $s_t$, the relative ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$ and the relative ratio of the exchange rate to the non-traded goods fundamental $s_t - f_t^N$. (a) USD/ZAR $s_t$ with an intercept. (b) USD/ZAR $s_t$ without an intercept. (c) USD/ZAR $s_t - f_t^T$ with an intercept. (d) USD/ZAR $s_t - f_t^T$ without an intercept. (e) USD/ZAR $s_t - f_t^N$ with an intercept. (f) USD/ZAR $s_t - f_t^N$ without an intercept.

### *5.4.2. Mexican Peso (MXN)*

The Mexican Peso was pegged to the US Dollar and the Peso was allowed to appreciate or depreciate against the US Dollar within a narrow target band. The Mexican central bank maintained the peg by frequently intervening in exchange rate markets (Whitt, 1996). As can be seen from Table 7, we find evidence of explosive behavior in the

**Table 7**
The GSADF test for exchange rate in emerging markets countries.

| Exchange rate | Test Stat under $H_0$ with an intercept | Episodes | Test Stat under $H_0$ without an intercept | Episodes |
|---|---|---|---|---|
| **USD/COP** | | | | |
| $s_t$ | 2.1757[a**] | 97M09-01M10, 02M07-03M04 | 5.4578[b***] | 94M08-14M12 |
| $s_t - f_t^N$ | 2.7464*** | 97M09-03M11, 05M11-06M03 | 4.9002*** | 95M06-08M02, 08M09-09M05 |
| | | 07M04-07M07, 08M01-08M08 | | |
| $s_t - f_t^T$ | 0.7397 | 94M08-94M12 | 2.1901 | 00M08-01M05, 02M07-03M04 |
| **USD/MXN** | | | | |
| $s_t$ | 3.5056*** | 94M02-94M04, 94M12-95M04 | 2.5653 | 98M08-98M11, 03M01-03M03 |
| $s_t - f_t^N$ | 3.3521*** | 94M02-94M04, 94M11-95M03 | 2.6254 | 98M08-99M03, 02M12-03M02 |
| | | 98M08-98M11, 08M04-08M08 | | 04M04-04M10 |
| $s_t - f_t^T$ | 1.8151 | 94M11-95M03 | 1.9643 | |

[a**] indicates significance at the 5% level.
[b***] indicates significance at the 1% level.

nominal Dollar-Mexican Peso exchange rate $s_t$ under the assumption of the intercept[8]. The null hypothesis of no bubbles in $s_t$ can be rejected at the 1% significance level. We observe two episodes from Fig. 14a (i.e., 1994M02-1994M04, 1994M12-1995M04).

Importantly, our results support the finding of explosiveness in USD/MXN between 1994 and 1995. The episode between 1994M12 and 1995M04 cannot be explained by two exchange rate fundamentals, which indicates the presence of rational bubbles. The 1994 Mexican currency crisis is one of the most well-known exchange rate crises in the literature. The North American Free Trade Agreement (NAFTA) came into force at the beginning of 1994 and was signed by Canada, Mexico and the US. The agreement aimed at encouraging foreign investors to take advantage of Mexican's access to the US market and lowering trade barriers between two countries (Whitt, 1996). However, in fewer than 12 months, the crisis exploded in December 1994, when the Mexican government suddenly devalued the Peso by 15%. Devaluation of the Peso led to a deep crisis in Mexico's financial services sector (Wilson et al., 2000). Thus the USD/MXN crisis of 1994-1995 is a bubble, which is of particular interest. However, when the intercept is removed from model formulation under the null hypothesis, all three series ($s_t$, $s_t - f_t^N$ and $s_t - f_t^T$) are not explosive. The null hypothesis of no bubbles cannot be rejected at the 10% level, suggesting no bubbles in USD/MXN. Although there are short-lived episodes in Figs. 14b and f during 1994–1995, we could not conclude that the crisis of 1994-1995 is a bubble when the intercept term is excluded in the null.

## 6. Conclusion

In this paper, we test for the explosiveness in the nominal exchange rate and if it is identified, investigate the cause of the explosiveness. We then explore whether the explosiveness in the nominal exchange rate is driven by rational bubbles or exchange rate fundamentals. We concur with Bettendorf and Chen (2013), that explosiveness in the asset price does not, on its own, imply the existence of rational bubbles, where it is necessary to consider the role played by economic fundamentals in asset prices. Following the recent work of Bettendorf and Chen (2013) and Jiang et al. (2015), we use the GSADF test of Phillips et al. (2015a, PSY) to investigate the evidence of exchange rate bubbles for both G10 and emerging markets countries (including some Asian and BRICS countries). The results can be summarized as follows.

[8] We let $r_0$=0.05 for the following analysis. This is due to the fact that the sample data starts from January 1993 and we would like to test for the evidence of exchange rate bubbles during Mexican currency crisis in 1994–1995. We also carry out an analysis by letting $r_0 = 0.01 + 1.8/\sqrt{T}$ and do not find significant evidence of bubbles.

Results for some G10 cross rates as presented in Tables 1, 2, 3 suggest, no evidence of bubbles in most exchange rate pairs with only a few exceptions. Under the assumption 'with an intercept', the GSADF test statistic for the Sterling-Swiss Franc and Sterling-Japanese Yen seems to suggest evidence of bubbles as the test statistic is significant at the 1% or 5% level in Table 1. In fact, the PSY identifies several 'collapse' episodes rather than bubbles as it cannot distinguish between 'collapse' episodes and bubbles if the intercept term is included in the null. Hence, we find little evidence of bubbles in these two exchange rate pairs.

Some interesting results are obtained from the Asian currencies. First, in line with the theory of Engel (1999) and Betts and Kehoe (2005), the relative prices of traded goods play an important role in explaining the majority of the movements in the US Dollar-Philippine Peso, US Dollar-Indonesian Rupiah and US Dollar-Singapore Dollar (under the model specification 'with an intercept') exchange rates. Second, our results indicate that the exchange rate movements between Korea, Malaysia, Thailand and the US cannot be explained by the theory of Engel (1999) and Betts and Kehoe (2005). We conclude that exchange rate fundamentals (the relative prices of traded goods and non-traded goods) do not explain the explosiveness in the US Dollar-Thai Baht and US Dollar-Korean Won exchange rates, which confirm the presence of rational bubbles. Unlike existing studies, our empirical results also suggest that the relative prices of traded goods don't explain most movements in the US Dollar-Malaysian Ringgit exchange rate under two model specifications. Last, we find evidence of bubbles or rational bubbles in several Asian currencies during the 1997 Asian Financial Crisis and also identify several 'collapse' episodes and 'collapse and recovery' episodes.

Our results from the three BRICS countries (e.g., Brazil, India and South African) suggest that the relative prices of traded goods account for the majority of the movements in exchange rates, which confirms Engel (1999) and Betts and Kehoe (2005). Overall, we find evidence of bubbles for these currencies but some evidence obtained from the model specification 'without an intercept' is spurious (e.g., Indian Rupee and South African Rand).

We also find evidence of explosive behavior in the US Dollar-Colombian Peso exchange rate but the evidence obtained from the model specification 'without an intercept' is spurious. The explosiveness in the US Dollar-Colombian Peso seems to be explained by the relative prices of traded goods. Moreover, we find significant evidence of explosive behavior in the US Dollar-Mexican Peso exchange rate as well. Our results also support the hypothesis that there is a bubble in the US Dollar-Mexican Peso exchange rate during the 1994–1995 Mexican currency crisis and this finding should be of some considerable interest.

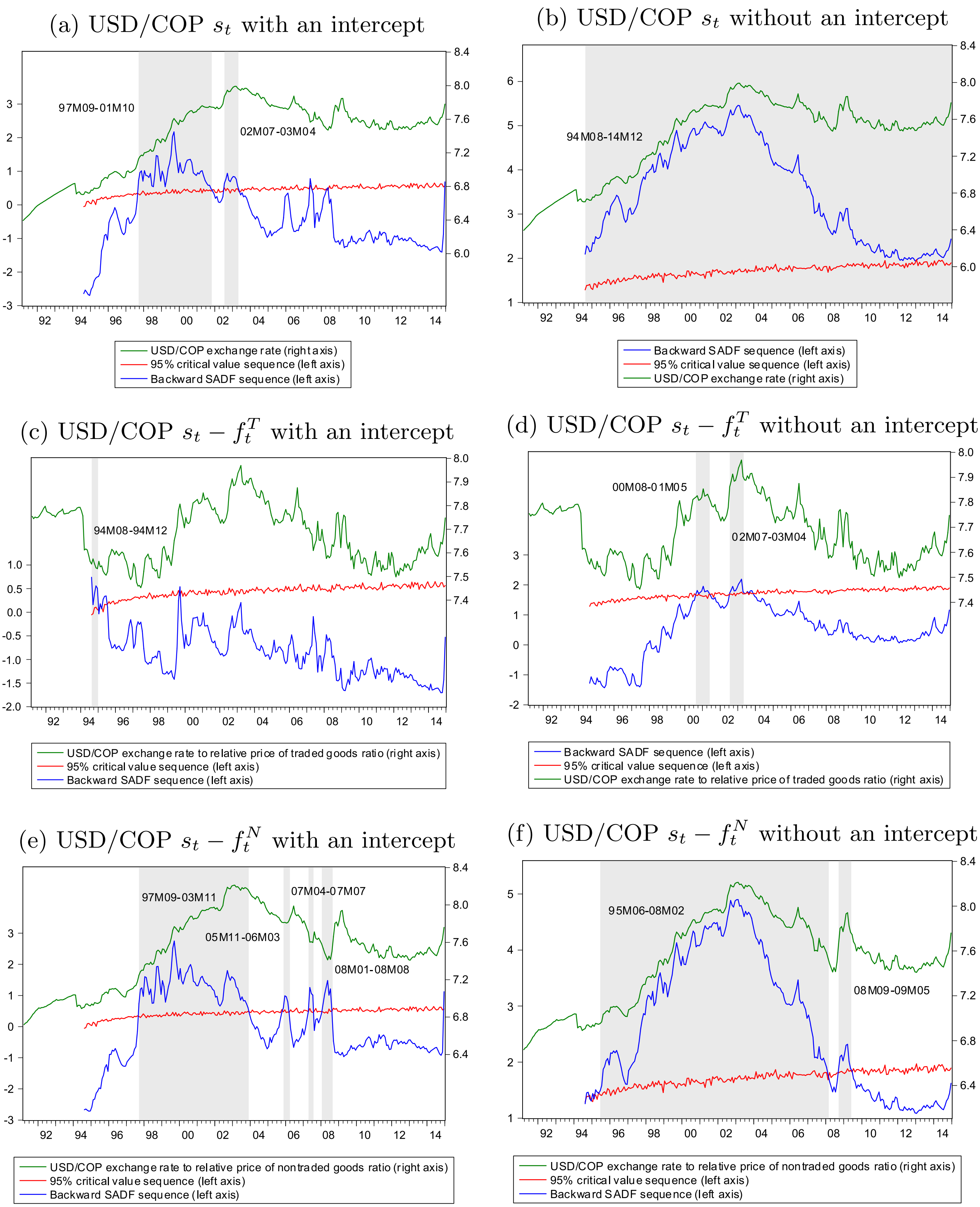


**Fig. 13.** Dating strategy for USD/COP nominal exchange rate $s_t$, the relative ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$ and the relative ratio of the exchange rate to the non-traded goods fundamental $s_t - f_t^N$. (a) USD/COP $s_t$ with an intercept. (b) USD/COP $s_t$ without an intercept. (c) USD/COP $s_t - f_t^T$ with an intercept. (d) USD/COP $s_t - f_t^T$ without an intercept. (e) USD/COP $s_t - f_t^N$ with an intercept. (f) USD/COP $s_t - f_t^N$ without an intercept.

Overall, we obtain quite different results when using a model specification 'without an intercept' in the null hypothesis. Firstly, the null hypothesis of no explosive bubbles is frequently not rejected as the critical values become larger under the model specification without an intercept. Secondly, when the intercept term is included in the model formulation for constructing the null hypothesis, we will identify both 'collapse' episodes, 'collapse and recovery' episodes and potential bubbles as the PSY cannot distinguish between the 'collapse' type of episodes and bubbles. Thirdly, if the null hypothesis involves no intercept, the 'collapse' type of episodes will not be identified by the

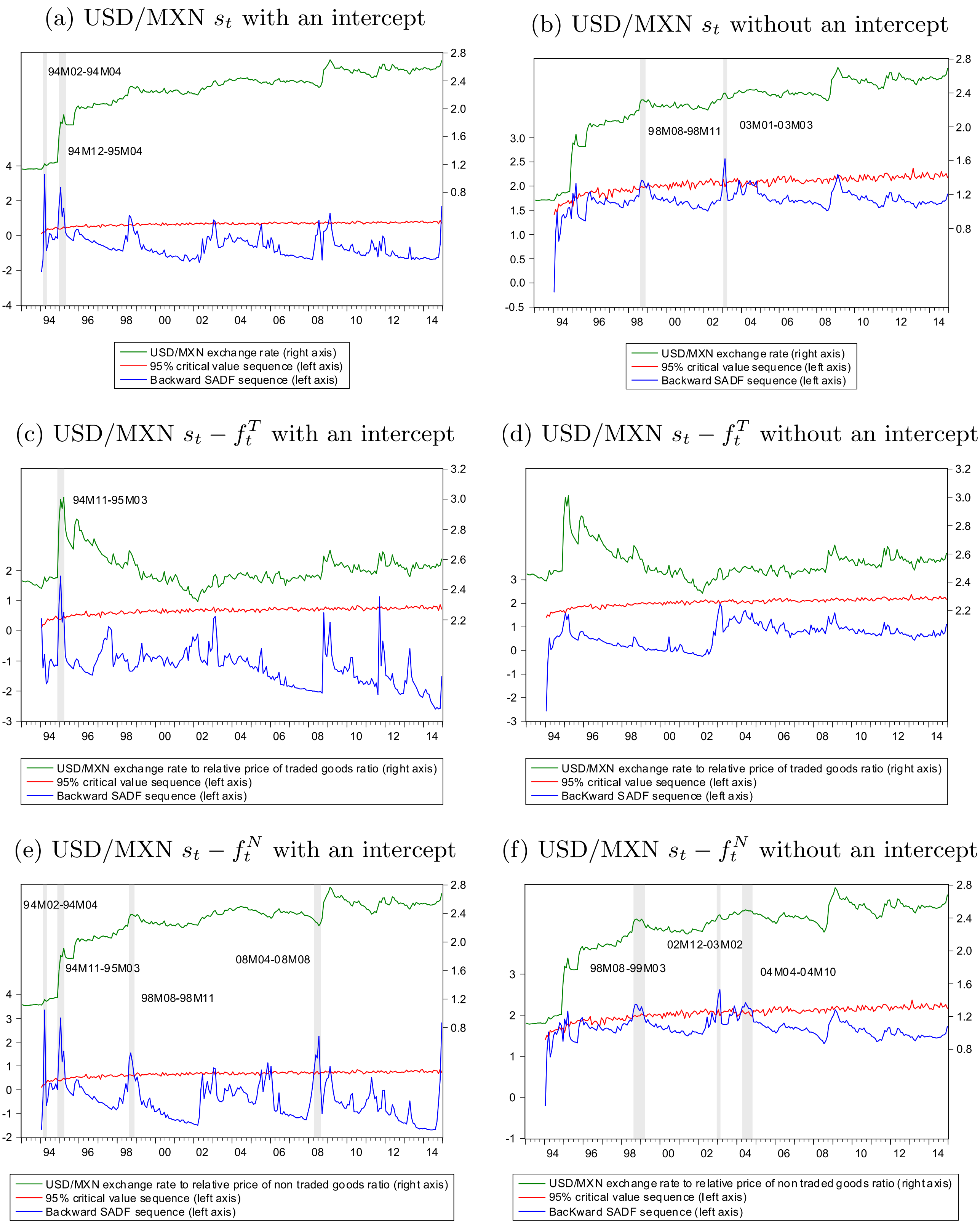


**Fig. 14.** Dating strategy for USD/MXN nominal exchange rate $s_t$, the relative ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$ and the relative ratio of the exchange rate to the non-traded goods fundamental $s_t - f_t^N$. (a) USD/MXN $s_t$ with an intercept. (b) USD/MXN $s_t$ without an intercept. (c) USD/MXN $s_t - f_t^T$ with an intercept. (d) USD/MXN $s_t - f_t^T$ without an intercept. (e) USD/MXN $s_t - f_t^N$ with an intercept. (f) USD/MXN $s_t - f_t^N$ without an intercept.

PSY approach but some episodes may be spurious (e.g., Philippine Peso, Indian Rupee, South African Rand and Colombian Peso). In short, the intercept term affects the asymptotic theory and date-stamping strategy of the PSY approach. The inclusion of the intercept demonstrates the practical importance in right-tailed unit root tests. It is of great importance to assess a wide range of specifications in the null and make a suitable choice. Finally, it seems that newly emerging economies are more likely to exhibit bubbles in the exchange rate than more mature countries, perhaps because their monetary policy stances are somewhat looser than for example, those in the G10.

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