Testing for Bubbles in Time Series Data Using Long Historical Series

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Abstract

The concept of asset price bubble has drawn a large amount of academic attention. A bubble is commonly defined as where the asset price is not justified by expectations of the fundamentals. The empirical characteristics of a bubble are widely regarded as a dramatic rise in asset prices in the early periods along with a sudden collapse, which have been observed in a wide range of financial markets including commodity, real estate and stock.

This thesis attempts to examine several well-noticed bubble episodes in history, for example, the Mississippi Bubble in 1719, the South Sea Bubble in 1719-20, Japan's asset price bubble during the 1980s and the more recent US real estate bubble in the early 2000s. One of the greatest contribution of this thesis is that it has provided econometric-based evidence to support the existence of these famous episodes for the first time using recently developed econometric tests. The origination and collapse dates for these identified episodes are coincided with the traditional view of famous bubbles.

This thesis not only tests for the existence of bubbles but also seeks to provide some empirical evidence to support the existence of bubble spillovers/contagion using the most-documented bubble episodes. The first example considered in this thesis is whether financial bubbles/crises spill-over to housing markets in Amsterdam, Norway and France. The second example is considered for a number of British financial organisations during the South Sea episode in 1719-20 (e.g., the South Sea Company, the Bank of England, East India Company, London Assurance, Million Bank, the Royal African Company and the Royal Exchange Assurance). The South Sea Bubble seems to migrate to other organisations as the British share market was speculative for that period. The timings of some identified episodes can be viewed as possible evidence of spillovers or contagion for this important period in England. The third example looks at the Japan's

real estate and stock markets during the 1980-90s. Based on econometric tools, the findings indicate signs of bubble migration from Japan's stock market to its real estate market for the first time. Even if this phenomenon has been observed in the literature, none of existing studies provide any solid evidences to support this claim.

This thesis investigates the long-run equilibrium relationship between house prices and economic fundamentals in the US based on quantile cointegration approaches for the first time. The presence of cointegration between house prices and fundamentals would imply convergence to a stable equilibrium relationship, suggesting a temporary run-up or drop in house prices would eventually come back to the equilibrium. By making use of quantile autoregressive distributed lag model (QARDL), this research explores the cointegration relationship across the whole conditional distribution.

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Contents

Chapter 1

Introduction

"if the reason that the price is high today is only because investors believe that the selling price is high tomorrow-when 'fundamental' factors do not seem to justify such a price-then a bubble exists."

— *(Stiglitz, 1990, p.13)*

1.1 Overview

Bubbles are one of the most beautiful concepts in economics and finance (Garber, 2001). Stiglitz (1990) introduces a simple definition of the bubble that may arise when asset prices in a market deviate from underlying fundamental values. The presence of an asset price bubble is not new to history despite a growing recent interest since the Global Financial Crisis. Dating back to as early as the 17th century in Netherlands, Tulipmania (1636-7) is commonly regarded as the earliest example of a bubble, where the contract prices of tulip bulbs were astonishingly high from 1636 and suddenly collapsed in 1637.[1] For example, Thompson (2007) reports that the contract price of tulips in February 1737 was 20 times higher than the price in November 1636 and May 1737. Furthermore, according to *The Economist*, a very rare tulip, Semper Augustu bulb, was initially valued at around

[1] A detailed review of the Tulipmania can be found at Garber (1989).

1,000 guilders in the 1620s and later valued at 5,500 guilders per bulb, which was equal to the cost of a luxurious house in Amsterdam.[2]

Shortly after the Dutch Tulipmania, the Mississippi Bubble in France and South Sea Bubble in England during 1719-1720 are regarded as next famous episodes of bubbles.[3] Both bubble episodes are some of the most influential stock price bubbles in history. The share prices of the Mississippi and South Sea Company achieves remarkable growth in values during the earlier period, however, both collapsed drastically and suddenly. Other than these two closely related episodes, the British Railway Mania of the mid-1840s is another well-known example of stock price bubbles in history, which resulted in a boom of railway construction in England and a financial panic in share prices. The Economist simply describes such an episode as 'arguably the greatest bubble in history' based on the scale of investment as a proportion of national income.[4] The railway shares peaked in the mid-1840s. By 1850, railway shares had fallen from their peak over 85%, see Chancellor (2000, p.149). These three historical bubble episodes during the 17th-19th centuries are still considered as the most famous examples and are frequently mentioned. As part of this thesis, I will consider the Mississippi Bubble and South Sea Bubble episodes in some detail.

Not only do the 17th-19th century highlight the presence of asset price bubbles, but also more recent experience reminds us of the existence of several other bubbles including but not limited to the German stock price bubble in 1927, the Wall Street Crash of 1929, Japanese asset price bubble of the 1980s-90s, the Dotcom bubble and the more recent US housing bubble during the 2000s. There have been some excellent studies on two bubble episodes during the 1920s, for example, see Voth (2003) for the German stock price bubble in 1927, Romer (1990), White (1990) and Rappoport and White (1994) for the Wall Street Crash of 1929. His-

[2] *The Economist*, "Economic history: Was tulipmania irrational?", 4 Oct, 2013.

[3] For a literature review on Mississippi and South Sea Bubbles, see Neal (1990), Murphy (1997), Temin and Voth (2004) and Frehen et al. (2013).

[4] *The Economist*, "The beauty of bubbles", 18 December 2008.

tory seems to suggest that the global financial system is more fragile after 1980 as several devastating bubbles occurred then. This thesis will look at some recent examples of bubbles including Japan's bubble period and the US real estate bubble.

During the second half of the 1980s, Japan's real estate and stock markets experienced their most severe episode of speculation. The dramatic and unsustainable rise in asset prices drew a lot of attention, while the collapse of asset prices has led to serious consequences. Two myths existed relating to Japan's exuberant stock and real estate markets. Firstly, that land prices could never fall; secondly, that stock prices could only rise (Malkiel, 2003). The economic losses from Japan's financial system during Japan's bubble economy were at least 20% of Japan's Gross Domestic Product (GDP) based on a conservative estimate from Hoshi and Kashyap (2004).

The Dotcom bubble of the 1990s or the NASDAQ bubble is a very recent example of speculation in stock prices, see Ofek and Richardson (2003), Ljungqvist and Wilhelm (2003), Pástor and Veronesi (2006) and Perez (2009). Perhaps the most famous phrase-'irrational exuberance' introduced by Alan Greenspan in 1996 is used to describe the significant surge in the stock market of the 1990s. According to Shiller (2005), the famous Greenspan speech in 1996 was given at the beginning of the most speculative growth in the US stock market history. The Dow Jones Industrial Average was 3600 at the beginning of 1994 but reached 10000 in March 1999 and peaked at 11722 in early 2000. The unusual rise and fall in the stock prices of technology stocks has led many academics and practitioners to describe such a phenomenon as a 'bubble'.

Several prominent economists acknowledged that the US experienced a speculative bubble in the housing market during the 2000s. Baker (2002) claimed no obvious explanation for the rapid rise in house prices and concluded the presence of a housing bubble in August 2002. In June 2005, Robert Shiller warned that "The market is in the throes of a bubble of unprecedented proportions that probably will end ugly".[5] He also argues that house prices could drop by up to 50% in the next decade. In August 2005, Paul Krugman pointed out the existence of a housing bubble in the coasts-areas due to a combination of high population density and land-use restrictions.[6] During the Economic Club Meeting of New York in 2005, Alan Greenspan concluded that: "we don't perceive that there is a national bubble, but it's hard not to see that there are a lot of local bubbles". However, Ben Bernanke argued that "house price increases largely reflect strong economic fundamentals including robust growth in jobs and incomes, low mortgage rates, steady rates of household formation, and factors that limit the expansion of housing supply in some areas".[7]

The majority of studies in the literature focus on some recent bubble episodes using short time series. In particular, most house price indices exist for a few decades. This thesis attempts to investigate explosive behavior using very long historical house price series for Amsterdam (1649-2010), Norway (1819-2014) and Paris (1650-2012). There are several financial bubbles and crises associated with these house price data, for example, the Amsterdam Banking Crisis of 1763, the Credit Crisis of 1772, the Kristiana Crisis of 1899-1905 and the Scandinavian crisis of 1984-1992 in Norway, the Crises of 1763 and 1772 in French. Theses house price series are therefore ideal candidates to apply recent bubble detection procedure to search for housing bubbles and potential contagion across different financial markets.

It has been shown that the Phillips, Shi, and Yu (2015, PSY) procedure can effectively detect the presence of multiple bubbles. There have been no empirical studies that make use of the PSY testing procedure to test for historical bubbles in the 18th century. It is not clear whether the PSY test can offer evidence to support the existence of those well-known bubble episodes. The question may arise here:

[5] Laing, Jonathan R, "The Bubblels New Home," Barron's, 20 June, 2005.

[6] Krugman, Paul, "That Hissing Sound," The New York Times, 8 August, 2005

[7] Bernanke, Ben S., "The Economic Outlook," Testimony before the Joint Economic Committee, 20 October, 2005.

can the recently developed right-tailed unit root tests detect historical bubbles during 18th centuries? Hence this thesis will attempt to apply the right-tailed unit root tests of PSY to seek econometric evidence to support the existence of the Mississippi and South Sea Bubbles. In addition, the share prices of a number of British financial organisations are also considered during the South Sea episode in 1720 as these shares also exhibit bubble-like behaviour.

Testing for bubbles in exchange rates is not new to the literature. Several well-published papers examine the presence of exchange rate bubbles including Huang (1981), Evans (1986), Meese (1986) and West (1987). For example, Huang (1981) finds evidence of bubbles in the US Dollar-Sterling, US Dollar-Mark, and Sterling-Mark exchange rates from March 1973 to March 1979 using the variance bounds test of Shiller (1980) and Leroy and Porter (1981). Evans (1986) presents evidence to support the presence of bubbles in the Sterling-Dollar exchange rate between 1981 and 1984. Similarly, Meese (1986) provided evidence of bubbles for the Dollar-Deutsche Mark and Sterling-Dollar exchange rate using the monthly data between 1973 and 1982. However, the conclusions drawn from West (1987) indicate no evidence of bubbles in the US Dollar-Mark exchange rate during the period 1974 to 1984. These classic studies on exchange rate bubbles focus on developed countries only. This thesis will, therefore, seek to identify the existence of bubbles in some G10 and a range of emerging markets countries (including some Asian and the BRICS). By selecting such a wide range of countries, this thesis 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 developed countries including the UK and US.

The US house price bubble during the 2000s is also considered in this study as it is one of the most recent bubble episodes. This thesis finds that the PSY testing procedure seems to identify false-positive bubbles under the regression model specification with an intercept. Empirical results from the US regional house prices seem to be in agreement with the talk given by Alan Greenspan in 2005, who argued that there was no sign of a nationwide housing bubble but a lot of local bubbles.

Financial bubbles are often related to spillover effects. Another key area to analyse is therefore market migration. This thesis will examine the existence of Japan's asset price bubble in its most inflated stock and real estate markets of the 1980s-90s. A a two-step testing procedure is adopted in this thesis to consider bubble migration effects. The PSY method is complemented by a new time-varying regression methodology of Greenaway-McGrevy and Phillips (2016) for analysing market contagion across markets. A striking feature of this great bubble is that Japan experiences at least two types of asset bubbles involving stocks and real estate, unlike most financial bubbles, which seem to relate a single type of asset bubble. As Japan experience speculative growth in stock and real estate markets, it seems an ideal candidate to examine evidence of bubble spillovers. This thesis demonstrates, for the first time in the literature, signs of bubble migration from Japan's stock market to its real estate market. We show this by using two new econometric tools.

Another area closely related to bubbles is to examine the long-run equilibrium relationship between asset prices and fundamentals. There have been many empirical studies on the US housing market based on cointegration analysis, for example, Gallin (2006), Mikhed and Zemčík (2009), Zhou (2010), Clark and Coggin (2011) and Duca et al. (2011). Cointegration between house prices and fundamentals is also important for understanding house price dynamics. The presence of cointegration between house prices and fundamentals would imply convergence to a stable long-run equilibrium relationship. In other words, cointegration would indicate whether a temporary run-up or drop in house prices can eventually back to the equilibrium. Many studies attempt to explore whether US house prices reflect their economic fundamentals using cointegration analysis. However, the existing studies are mainly based on conditional mean based cointegration analysis and do not reach a firm conclusion. This thesis will re-visit the hypothesis between US house prices and the relevant economic fundamentals using the recently proposed linear and non-linear quantile cointegration models. Conventional cointegration approaches focus on the conditional mean behaviour. Apart from the mean, other locations in the distribution affect the cointegration analysis. The estimated relationship from conventional cointegration approaches assumes to be held at the mean and even in the tails. This thesis will formally test for the long-run cointegration relationship at each of the quantiles by using formal testing procedures of the linear and non-linear quantile ARDL based models for the first time in the literature. By applying the linear and non-linear quantile autoregressive distributed-lag models of Cho et al. (2015) and Greenwood-Nimmo et al. (2011), quantile cointegration models allow us to study the entire conditional distribution of the dependent variable (US housing market) compared with those of conditional mean-based cointegration models. As the housing market is the core focus, we may interpret the conditional distribution of housing market with lower quantiles (at a relatively low price level) as a bear market, the median quantiles as the normal phase of the market, and the upper quantiles (at a relatively high price level) as a bull market.

1.2 Background and Research Questions

Speculation in the Mississippi and South Sea Bubbles seem to behave in a similar way. In particular, speculation in these historical events involved a company (the Mississippi Company and the South Sea Company) that expanded its balance sheet through corporate takeovers or acquisition of government debt, financed by successive issues of shares (Garber, 1990). The motivation of the Mississippi and the South Sea schemes was to refinance the national debts accumulated during the War of the Spanish Succession, see Hamilton (1947) and Dickson (1967). The Mississippi bubble is an economic bubble that resulted from John Law's 'system'. Law developed and adopted a 'system' to take over the French national debt accumulated by the wars of Louis XIV by using equity. The share price of the Mississippi Company soared to more than 10,000 livres per share by December 1719 and dropped significantly in 1720.

The South Sea Company was first founded in 1711. The South Sea Bubble also involved the South Sea Company that acquired some outstanding British government debt in 1720. However, the South Sea Company was not involved in takeovers of commercial companies. According to Garber (1990), the British debt in 1720 was worth approximately £50 million and £18.3 million of the debt was held by three of the largest companies: the Bank of England (£3.4 million), East India Company (£3.2 million) and South Sea Company (£11.7 million). In 1720, the South Sea Company had monopoly rights on British trade with the Spanish colonies of South America. The South Sea Company, Bank of England, and East India Company played a major role in the South Sea Bubble as they engaged in the debt-for-equity swap. The Bank of England was founded in 1694. The Bank was the first permanent stock and the longest-lived security in the London Stock Exchange (Neal, 1990). According to Carlos et al. (2006), Bank of England shares were one of few publicly available securities, and shares are reorganised as a stable asset as they are the least speculative during 1720. The original East India Company was founded in 1600 with a royal charter, and it became a permanent joint-stock company in 1657. The new East India Company was established in 1698. The old and new companies were merged by 1709.

In this thesis, I will also consider those British companies that were not involved in the debt-equity swap (e.g., London Assurance, Million Bank, the Royal African Company and the Royal Exchange Assurance). The Million Bank was founded in 1695 and held a large number of securities. A review of the Million Bank is provided by Scott (1911). The Royal African Company received a Royal charter in 1672 and then became the second largest of the joint-stock companies after the East India Company (Carlos et al., 2002). The company had the monopoly of English trade in Africa and engaged in the slave trade, but was not involved in the financing of the national debt. In 1720, the company issued stock that rose quickly in value. One of the innovations in 1720 was the establishment of marine insurance companies-Royal Exchange Assurance and London Assurance. The early growth of the British insurance industry is completed by the establishment of two insurance companies in 1720 along with their powers to include fire and life insurance in 1721 (Supple, 1970). Both insurance companies founded in 1720 are a major innovation in sharing-risk for foreign trade. The creation of the Royal Exchange Assurance is important not only in the development of insurance in the British history but also the share market in Britain. Such a company was needed to provide a secure service for the marine trade. The growth of the London insurance market was associated with the marked expansion of English foreign trade during the second half of the seventeenth century; and forms part of that remarkable period of financial activity culminating in the South Sea Bubble (John, 1958). Marine insurance played a vital role in facilitating the expansion of trade during the eighteenth and nineteenth centuries (Kingston, 2007).

One of the most researched topics in bubbles is the US real estate market. There are two periods of extreme movements in US house price during the 1980s and 2000s as suggested by Mayer (2011). House prices firstly rose and later plummeted in many US regions during the 1980s. Consequently, the boom-bust cycles occurred in California, Boston, Hawaii and New York, for example, Case (1986), Case and Shiller (1988), Case (1994) and Case and Shiller (1994). The second period of boom-bust cycles in the 2000s seems to be more widespread. The real US house prices increased by 5% per year between 1996 and 2006, and then dropped by 6.4% per year from 2007 to 2012 (Glaeser et al., 2017). A falling interest rate, income growth and reduced lending standards are generally considered as causes of the recent cycles, see Case and Shiller (2003) and McCarthy and Peach (2004). There has been considerable debate over the existence of a housing bubble in the 2000s by several prominent economists. Therefore, a number of studies investigate the US housing bubbles during the 2000s, for example, see Case and Shiller (2003), McCarthy and Peach (2004), Himmelberg et al. (2005) and Martin (2011).

Japan's asset price bubble is commonly referred as one of the greatest bubble episodes in recent history. There has been an unprecedented rise in Japan's land, real estate and stock prices. According to Malkiel (2003, p.77), the value of real estate and stocks increased over 75 and 100 times from 1955 to 1990. There are two 'myths' attributed to the inflated Japanese real estate and stock market that are of particular interest. These myths formed during a period where these markets were undergoing extraordinary increases in price levels. The first myth been that land prices could never fall; the second that stock prices could only rise. Both myths were fuelled by large amounts of cash savings (Malkiel, 2003, p.78). It is widely accepted that both markets are characterised by bubble-like behaviours, see, for example, Hardouvelis (1988), Lee (1995), French and Poterba (1991), and Chirinko and Schaller (2001). As discussed in Okina et al. (2001), the Nikkei 225 reached a historical high of 38,130 in December 1989 from a start of 12,598 in September 1985, and then dropped sharply to 14,309 in August 1992, which represented a more than 60% decline from the peak. When we consider land prices in September 1990, they were around four times higher than the level in September 1985, only to drop in 1999 by approximately 80% from the peak in 1990. The bursting of 'double bubbles' in Japan's stock and real estate markets led to a decade of economic stagnation with large amounts of government debt. The aftermath of economic stagnation in the 1990s after the bursting of Japanese asset price bubble is commonly referred to as "the Lost Decade". Consequently, Japan experiences slow economic growth, deflation and output persistently below potential during "the Lost Decade", see Kuttner and Posen (2001), Hayashi and Prescott (2002), Hoshi and Kashyap (2004) and Horioka (2006).

In summary, the thesis will consider the following research questions:

1. Are there bubbles in exchange rates?
    (a) Are exchange rate bubbles more likely to arise in emerging countries or developed countries?
    (b) Are results sensitive to model specification of methods used?

2. Do historical house prices in Amsterdam, Norway and Paris exhibit bubbles or explosive behavior?

    (a) Can long memory models detect the presence of bubbles or explosive behavior?

    (b) Do financial bubbles/crises spill-over into local or national markets?

    (c) Do housing bubbles coincide with major financial crises in Norwegian history?

3. Is there evidence of real estate bubbles in the US during the 1980s and 2000s?

    (a) Does the US experience a national housing bubble during the 2000s?

    (b) Is the 2000s housing bubble more widespread than the 1980s housing bubble?

4. Can the recently developed right-tailed unit root test identify historical stock market bubbles - the Mississippi and South Sea Bubbles?

    (a) Do a number of other 18th century financial organisations in Britain also experience bubble-like behaviour during the South Sea Bubble?

    (b) Do bubbles spill-over?

5. Did bubbles exist in Japan's stock and real estate markets in the 1980-90s?

    (a) Do Japanese asset price bubbles migrate from the stock market to the housing market or vice versa?

6. Do US house prices reflect economic fundamentals?

    (a) Is there a linear or non-linear long-run equilibrium relationship between house prices and economic fundamentals at the national and regional levels based on conventional cointegration tests?

(b) Is there a linear or non-linear long-run equilibrium relationship between house prices and economic fundamentals at the national and regional levels based on recently developed quantile cointegration models?

1.3 Significance and Contribution to the Research

This thesis has made several key contributions to the literature.

- Firstly, this thesis provides econometric-based evidence of several well-known bubble using the recent econometric tests.

  - This thesis first revisits the existence of exchange rate bubbles in both developed and emerging market countries using newly developed bubble detection methods due to Phillips et al. (2015). The main findings seem to suggest that newly emerging countries are more likely to exhibit exchange rate bubbles due to shallow markets or loose monetary policy. Moreover, there is little evidence to support exchange rate bubbles in G10 countries.

  - This thesis tests for housing bubbles or explosive behavior in historical house price series of Amsterdam, Norway and Paris. Little evidence of exuberance is found in the Amsterdam and Paris data. Of particular interest is that, evidence of exuberance is identified in Norwegian house prices. The identified episodes coincide with major financial crises in Norwegian history at the national and city levels. Both the long memory models and PSY approaches are used for identifying explosive behavior or bubbles.

  - The US real estate market is generally believed to be contain a bubble episode especially during the 2000s. This thesis concentrates on the US regional housing market between 1975 and 2014 with several aims. First, a major focus is to examine evidence for housing bubble(s) during the 1980s with particular emphasis on the States of California, Hawaii, Massachusetts and New York. Several other studies, for example, Case and Shiller (1988), Case and Shiller (1994), Case and Shiller (2003), Wheelock et al. (2006) consider US house price booms and busts or bubbles during the 1980s. However, most of these studies describe or graphically inspect some house price measure (e.g., house price, price-income ratio or price-rent ratio) without applying any econometric tests. Thus this study aims to fill the gap by providing econometric-based evidence of housing bubbles during the 1980s. This research also provides evidence to support the well-documented housing bubbles in the 2000s. Empirical results present evidence of a housing bubble in more than 20 states and the District of Columbia, during a period of speculation from the early 2000s to mid-2000s. It seems that the bubble of the 2000s is not a national phenomenon, but is more widespread than the 1980s. This conclusion seems to be the first empirical study to make a comparison in terms of their magnitude and coverage between the regional housing bubbles in the 1980s with the recent regional bubbles in the 2000s, which also contributes to the novelty of this study. Third, the importance of regression model formulation highlighted in Phillips, Shi, and Yu (2014) is also demonstrated by exploring the role of the intercept in the regression model of Phillips, Shi, and Yu (2015). Based upon the model specification with an intercept, the testing procedure can identify 'collapse' episodes, 'collapse and recovery' episodes and the potential bubbles. Whereas without an intercept in the regression model leads to identification of no 'collapse' episodes and 'collapse and recovery' episodes. An important finding under the regression model specification without the intercept in this thesis is that several States do not exhibit any bubble-like behaviors for the whole sample period.

- The most important findings is that for the first time in the literature, robust and powerful econometric-based evidence is found to support the existence of a South Sea bubble, using the right-tailed unit root tests of PSY. A number of other 18th century financial organisations in Britain are also considered during the South Sea episode and several significant results are presented. The South Sea Company experiences the first bubble-like behaviour in the British market. Furthermore, the South Sea episode is not the first one to burst and it lasts the longest period. Several other British share prices also exhibit bubble-like behaviour (e.g., East India Company, London Assurance, Million Bank, the Royal African Company and the Royal Exchange Assurance). There is little evidence of bubble-like behaviour for the Bank of England as it is widely treated as the least speculative stock among the major joint-stock companies. These results seem to suggest that the British share market was generally much more speculative during the South Sea episode than was previously thought, as the South Sea Company was not the only one experiencing bubble-like behaviour in its share prices. It seems to suggest that the South Sea episode does spillover to other British share prices.

– This thesis revisits the well-documented Japanese asset price bubble period with a particular focus on its irrationally exuberant stock and real estate markets using an appropriate bubble detection test. Empirical results present significant evidence of bubbles in Japan's stock price-earnings and house price-rent ratios. The timing of the origination and collapse dates of both stock and real estate bubbles coincide with the traditional view of Japan's asset price bubble for the period 1980s-90s. Such a finding is a useful addition to the literature for the most remarkable financial events in Japan's history, as it provides, for the first time, econometric-based evidence to support the existence of bubbles in Japanese asset prices from the late 1980s to early 1990s, which rose to extraordinary heights. Furthermore, using a recently developed time-varying regression based contagion methodology of Greenaway-McGrevy and Phillips (2016) to examine evidence of bubble spillovers between Japan's stock and real estate markets during the Japanese bubble economy period for the first time in a rigorous approach. The contagion results clearly show signs of migration from a stock price bubble to a real estate bubble, which is an addition to the literature. This finding may help to understand why the real estate bubble bursts after the stock price bubble as the bubble-like behaviour from the stock market not only migrates to, but fuels, the booming real estate market as well.

- Secondly, this thesis provides empirical guidelines for practitioners using the most popular PSY tests based on two model specifications (e.g., intercept vs. no intercept).

  - This thesis explores the relationship between model specification and the probability of a false positive, and discusses the advantages and disadvantages of the two specifications. Phillips et al. (2014) argue that without an intercept from the ADF regression inflates the probability of a false positive. They, therefore, recommend the regression specification with an intercept in practice. Under the assumption of an intercept in the regression, this specification certainly works well in many situations. A number of empirical papers follow this suggestion. However, in some cases, this specification also leads to a false positive identification of bubbles. The following section describes the pros and cons of the two approaches.

    Advantage of "with an intercept"

    * Under the specification with an intercept, the PSY procedure has shown great power in detecting multiple bubbles by a wide range of empirical studies.

    Disadvantage of "with an intercept"

* The PSY approach could detect some collapse episodes rather than a (positive) bubble. On page 156, Figure 1 in Chapter 5 presents the date-stamping outcomes based on the model "with an intercept". An example of 'collapse episode' and 'collapse and recovery episode' can be seen in Figure 1(a) and Figure 1(b). As can be seen from both figures, the PSY approach cannot distinguish between a 'collapse episode' or a 'collapse and recovery episode' and a genuine 'bubble'.
* The PSY approach is also applied to the Mississippi share prices and results are presented as Figure 2a on page 189 in Chapter 6. The null hypothesis of no explosive behaviour is rejected at the 1% level. However, upon closer inspection, the rejection is caused by a 'collapse and recovery' episode.
* Overall, the inclusion of an intercept could increase the probability of an incorrect rejection of the null hypothesis (false positive) by detecting a collapse episode or a collapse and recovery episode.

Advantage of "without an intercept"

* The PSY approach is applied to a house price-income ratio in Nevada based on two specifications, see Figure 6(e) on page 162 and Figure 7(e) on page 163 in Chapter 5. The results from Figure 6(e) based on the model with an intercept are spurious as some collapse episodes are identified. It seems that the specification "without an intercept" works well if data exhibits a steady downward trend like Figure 7(e) as the PSY approach does not detect collapse episodes any more.

Disadvantage of "without an intercept"

* It seems that the specification "without an intercept" does not work well if data exhibits an upward trend, for example see Figure 11b on page 91 and Figure 13b on page 94 in Chapter 3. This

specification does not always provide “correct” results as the results may be spurious.

General comments of the two approaches

* Both the two approaches could increase the probability of an incorrect rejection of the null hypothesis. Neither of the two approaches uniformly dominates the other in terms of inflating the risk of a false positive. More importantly, it is necessary to carry out post-test checking for collapse episodes/ collapse and recovery episodes.

- Thirdly, this thesis shows how to formally test for quantile coinetgration using two ARDL models.

  – This thesis employs both conditional mean- and quantial-based cointegration models to examine a long-run equilibrium relationship between US house prices and economic fundamentals at the national and selected state levels. A key contribution to the literature is that formal cointegration testing procedures for two quantile ARDL models of Cho et al. (2015) and Greenwood-Nimmo et al. (2011) are utilised. Cho et al. (2015) extends Pesaran and Shin’s (1998) autoregressive distributed-lag approach into quantile regression, resulting in a new quantile autoregressive distributed-lag (QARDL) model. Similarly, Greenwood-Nimmo et al. (2011) also develop a non-linear autoregressive distributed-lag quantile (NARDL-Q) regression model, which is extended from the non-linear autoregressive distributed lag (NARDL) model by Shin et al. (2014). However, a formal cointegration testing mechanism that could address the existence of the long run cointegration relationship at each quantile is not provided by Cho et al. (2015) and Greenwood-Nimmo et al. (2011). The research presented here makes use of the t-test and Wald test to jointly test for quantile cointegration at each of the quantile. The first testing procedure of Banerjee et al. (1998) is to carry out a t-test. The second testing procedure is a Wald test, which is used to test for quantile cointegration based on asymptotic theory. This research formally tests for the long run cointegration relationship at each of quantiles by two formal testing procedures of linear and non-linear quantile ARDL based models for the first time in the literature. Several interesting findings are presented using the US evidence. The US national house prices are not cointegrated with economic fundamentals using both conditional mean- and quantile ARDL based models. The empirical results also suggest that the cointegration relationship may be absent at the conditional mean-based model, but it may be established at the quantile based model. Under some instances, the cointegration may be established based on the conditional mean model, but there is no overwhelming evidence of cointegration based on a quantile model. Additionally, it is also possible to identify a cointegration relationship at the conditional mean and the relationship may also be held for the whole conditional distribution by quantile models.

1.4 Thesis Outline

This section overviews the outline of this thesis, which is made up of nine chapters. Each of the chapters in this thesis is briefly described as follows.

Chapter 1 provides the overview of this thesis and discusses background and research questions, and the major contributions of the study.

Chapter 2 presents the theoretical framework and existing studies of financial bubbles. This chapter begins with a definition and description of a financial bubble. A particular focus of this chapter is to re-visit empirical studies on US stock market based on various approaches as the US stock market is one of the most researched topics on bubbles.

Chapter 3 examines exchange rate bubbles for some G10 and a range of emerging markets countries including some Asian and the BRICS. The selection of sample countries considers whether exchange rate bubbles might be more likely to occur 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 developed countries. 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.

Chapter 4 explores evidence of explosive behavior in several historical housing price indices, in particular, the Herengracht index for Amsterdam (1649-2010), Norway (1819-2014) and Paris (1650-2012) using the PSY procedure and long memory models of Geweke and Porter-Hudak (1983), Robinson (1995) and Shimotsu (2010). In particular, this chapter also tests whether historical bubbles/crises spill-over to local or national housing markets. In addition, several major Norwegian cities are considered in this study including Oslo, Bergen, Trondheim and Kristiansand.

Chapter 5 examines the presence of bubbles in the US real estate markets at the state level for the period 1975-2014. One of the most empirically investigated economic phenomena on the subject of bubbles has been the potential bubble-like behavior in the US housing markets. A number of empirical studies in the literature suggest that the US market is widely regarded as one of the most speculative housing markets in the world as the US market experienced several period of speculation in history. This chapter makes a comparison in terms of their magnitude and coverage between the regional housing bubbles in the 1980s with the recent regional bubbles in the 2000s.

Chapter 6 investigates bubble-like behaviour in historical stock prices during the Mississippi and South Sea Bubbles in the 18th century. Of particular interest is that, these two bubbles are widely believed to be most speculative episodes and earliest examples in stock markets. In addition, this chapter also investigates a number of other 18th century financial organisations in Britain during the 1719-20, for example, Bank of England, London Assurance, Million Bank, East India Company, the Royal African Company and the Royal Exchange Assurance. The timings of these bubble episodes show signs of possible contagion during South Sea episode.

Chapter 7 investigates Japan's asset price bubble during its most inflated stock and real estate markets during this period in history. This chapter also uses the recently developed time-varying regression based contagion methodology of Greenaway-McGrevy and Phillips (2016) to examine evidence of bubble spillovers between Japan's stock and real estate markets during the Japanese bubble economy period for the first time in a rigours approach. This chapter reviews the existing empirical work on Japanese bubbles and bubble contagion using various approaches.

An important question that has been researched extensively in the literature, is whether US house prices reflect economic fundamentals. Chapter 8 explores a long-run equilibrium relationship between US house prices and economic fundamentals based on both conventional cointegration and quantile cointegration models. Conventional cointegration approaches focus on the conditional mean behaviour, which is perhaps not informative (Cho et al., 2015). It is assumed that the estimated relationship holds at the mean and even in the tails. Apart from the mean, other locations in the distribution also affect the cointegration analysis, which is well demonstrated by several empirical studies in the literature based on quantile regression model of Xiao (2009). Conventional cointegration tests focus on the conditional mean relationship between variables while quantile cointegrating models concentrate on a long-run relationship in a range of quantiles. The quantile-based cointegration tests are more informative as they provide additional information not only at the mean but also in the tails. This chapter will review the existing quantile cointegration models of Xiao (2009), Kuriyama (2016), Cho et al. (2015) and Greenwood-Nimmo et al. (2011) and the revelent empirical studies based on these models. Previous studies between US house prices and fundamentals based on cointegration analysis is also provided. The highlight of this chapter is that it applies linear and non-linear quantile based cointegration models to examine a quantile-varying cointegration relationship between house prices and fundamentals. Two testing procedures for a long-run relationship are also utilised and described in this chapter. The conclusion drawn from this chapter provides further evidence of a cointegration or non-cointegration relationship for the national and several key states, which is an additional contribution.

Chapter 9 provides the overall findings of this thesis by way of a concluding chapter. This chapter also suggests several avenues for future research.

1.5 Note on Publications

A number of journal articles have been produced from this thesis.

1. Hu, Y., & Oxley, L. (2017). Are there bubbles in exchange rates? Some new evidence from G10 and emerging market economies. ***Economic Modelling***, 64, 419-442.
(This publication is presented as Chapter 3.)

2. Hu, Y., & Oxley, L. (2018). Bubbles in US regional house prices: Evidence from house price-income ratios at the State level. ***Applied Economics***, 50, 3196-3229.
(This publication is presented as Chapter 5.)

3. Hu, Y., & Oxley, L. (2018). Do 18th century 'bubbles' survive the scrutiny of 21st century time series econometrics? ***Economics Letters***, 162 , 131-134.
(This publication is presented as Chapter 6.)

4. Hu, Y., & Oxley, L. (2018). Bubble contagion: Evidence from Japan's asset price bubble of the 1980s-90s. ***Journal of the Japanese and International Economies***, 50, 89-95.

(This publication is presented as Chapter 7.)

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Chapter 2

Literature Review

This chapter reviews the theoretical literature on bubbles; several different types of bubbles; the key econometric methodologies and the relevant empirical studies related to bubbles.

2.1 Theoretical Literature on Asset Bubbles

This section briefly summary of the theoretical aspects of bubbles. A summary of economic theory of asset bubbles can be found in Miao (2014).

Asset bubbles can be generated in overlapping generations models, see Samuelson (1958). The seminal work of Tirole (1985) studied conditions for the existence of asset bubbles using an overlapping generations model with inelastic labor supply. Tirole (1985) showed that asset bubbles can occur only if the economy is dynamically inefficient if there is an overaccumulation of capital. The criterion of dynamic efficiency in Tirole's (1985) work is introduced by Cass (1972).

Similarly, Weil (1987) suggested a model of stochastic bubbles using the same framework as Tirole (1985), and derived even stronger conditions for the existence of bubbles. If the economy is dynamically inefficient, stochastic bubbles exist only if the probability that they will persist next period is high. Weil (1987) also proved that the smaller the size of the inefficiency, the higher the required probability that the bubble would persist. Apart from the above mentioned studies, Several other studies on overlapping generations models of bubbles are worth of special mention here including Caballero et al. (2006), Caballero and Krishnamurthy (2006), Arce and López-Salido (2011) and Galí (2014).

Santos and Woodford (1997) also provided an intertemporal competitive equilibrium framework and established conditions for the non-existence of bubbles. Asset bubbles cannot exist in equilibrium with agents facing borrowing constraints if assets are in strictly positive supply and the present value of total future resources is finite.

Apart from the literature on overlapping generations models of bubbles, infinite-horizon models are also utilised to study bubbles. Kocherlakota (1992) and Kocherlakota (2008) are two excellent papers on bubbles based on infinite-horizon models. In particular, asset bubbles cannot be eliminated if some portfolio constraints can be introduced to limit the agents' arbitrage opportunities. Both papers study endowment economies to understand the asset pricing bubbles. Several papers study infinite-horizon models of production economies with bubbles, see Miao and Wang (2012), Wang and Wen (2012) and Miao and Wang (2014).

2.2 Types of Bubbles

2.2.1 Rational Bubbles

The following section draws on Campbell et al. (1997). The net return on a stock may be given as:

$$R_{t+1} = \frac{P_{t+1} - P_t + D_{t+1}}{P_t}, \tag{2.1}$$

where R_{t+1} denotes the return on the stock held from time t to time $t+1$; P_t denotes the price of a share of stock measured at the end of period t, or equivalently an ex-dividend price; D_{t+1} is the next period's dividend.

It is important to assume that the expected stock return is equal to a constant R:

$$E_t(R_{t+1}) = R, \tag{2.2}$$

where E_t denotes expectations conditional on information at time t.

Taking the expectation of Equation (2.1) based on available information at period t, we may rewrite Equation (2.2) as:

$$E_t(R_{t+1}) = \frac{E(P_{t+1} + D_{t+1}) - P_t}{P_t} = R. \tag{2.3}$$

Taking expectations of Equation (2.1), imposing Equation (2.2), and re-arranging, we obtain the equation relating the stock price to next period's expected stock price and dividend:

$$P_t = E_t\left[\frac{P_{t+1} + D_{t+1}}{1+R}\right] \tag{2.4}$$

Equation (2.4) is also known as an "expectational difference equation". It can be solved forward by repeatedly substituting out future prices and using the Law of Iterated Expectations - $E_t\left[E_{t+1}\left[X\right]\right] = E_t\left[X\right]$ to eliminate future-dated expectations. After solving forward K periods, we can now write as:

$$P_t = E_t\left[\sum_{i=1}^{K}\left(\frac{1}{1+R}\right)^i D_{t+i}\right] + E_t\left[\left(\frac{1}{1+R}\right)^K P_{t+K}\right]. \tag{2.5}$$

The second term on the right-hand side of Equation (2.5) is the expected discounted value of the stock price K periods from the present. Assuming that this term converges to zero as K increases:

$$\lim_{x\to\infty} E_t\left[\left(\frac{1}{1+R}\right)^K P_{t+K}\right] = 0. \tag{2.6}$$

Letting K increases in Equation (2.5), and the convergence assumption of Equation (2.6) holds. Then the expected present value, which is also the stock fundamental price (P_t^f), may be expressed as the expected discounted value of future dividends, discounted at a constant rate, which is presented as Equation (2.7):

$$P_t^f = E_t\left[\sum_{i=1}^{\infty}\left(\frac{1}{1+R}\right)^i D_{t+i}\right]. \tag{2.7}$$

Equation (2.7) shows that the stock price must be equal to the expected present value of future dividends.

If the convergence assumption of Equation (2.6) does not hold, there is no unique solution to Equation (2.4). In other words, if we relax this assumption, then the solution to Equation (2.4) can be written as:

$$P_t = P_t^f + B_t, \tag{2.8}$$

where $B_t = E_t\left[\frac{B_{t+1}}{1+R}\right]$. The second term B_t in Equation (2.8) is often called a *rational bubble*. In rational bubble models, the price, therefore, includes the fundamental component (P_t^f) and the bubble component (B_t). Several key empirical studies related to rational bubbles are summarised as Table 2.1 below.

2.2.2 Periodically Collapsing Bubbles

A very important class of rational bubbles is known as periodically collapsing bubbles. Evans (1991) defines the following rational bubbles that are always positive but periodically collapse:

$$B_{t+1} = (1+r)B_t u_{t+1}, \qquad \text{if} \quad B_t \leq \alpha \tag{2.9}$$

$$B_{t+1} = \left[\delta + \pi^{-1}(1+r)\theta_{t+1} \times (B_t - (1+r)^{-1}\delta)\right] u_{t+1}, \qquad \text{if} \quad B_t > \alpha \tag{2.10}$$

where δ and α are positive parameters with $0 < \delta < (1+r)\alpha$, u_{t+1} is an exogenous independently and identically distributed positive random variable with $E_t U_{t+1} = 1$, and θ_{t+1} is an exogenous independently and identically distributed Bernoulli process (independent of u) which takes the value 1 with probability π and 0 with probability $1 - \pi$, where $0 < \pi \leq 1$.

When $B_t \leq \alpha$, the bubble grows at mean rate $1 + r$. When $B_t > \alpha$, the bubble shifts into a phase in which it grows at the faster mean rate $(1 + r)\pi^{-1}$ while the bubble collapses with probability $1 - \pi$. When the bubble collapses, it falls to a mean value of δ and the bubble process begins again.

Table 2.1 – Empirical studies for testing rational bubbles in the literature.

Authors	Sample data	Method	Conclusion: a bubble?
Campbell and Shiller (1987)	S&P 1871-1986	Cointegration test of Engle and Granger (1987)	No
Diba and Grossman (1988)	S&P 1871-1986	Cointegration test of Engle and Granger (1987)	No
Lim and Phoon (1991)	S&P 1871-1986	Cointegration test of Engle and Granger (1987)	Yes
Craine (1993)	S&P 1871-1988	Augmented Dickey-Fuller unit root test of Dickey and Fuller (1979)	Yes
Han (1996)	S&P 1871-1986	Canonical cointegrating regressions of Park (1992) cointegration test of Johansen (1991)	Yes
Lamont (1998)	S&P 1947-1994	Cointegration test of of Horvath and Watson (1995)	No
Koustas and Serletis (2005)	S&P 1871-2000	Fractional integration/structual break test	Yes
Chang et al. (2007)	S&P 1871-2002	Johansen cointegration test	Yes
		Bierens (1997) nonparametric cointegration test	No

Periodically collapsing bubbles are not detectable by conventional unit root and cointegration tests. Table 2.2 below summaries several empirical studies related to the periodically collapsing bubbles of Evans (1991).

2.2.3 Intrinsic Bubbles

Froot and Obstfeld (1991) argue that US stock prices may be explained by the presence a specific type of rational bubble that depends exclusively on dividends (fundamentals). An intrinsic bubble solely relies on fundamentals alone, these bubbles will remain constant over time for a given level of fundamentals. Froot and Obstfeld (1991) also state that these intrinsic bubbles can cause asset prices to overreact to changes in fundamentals.

The following section is based upon Froot and Obstfeld (1991). Suppose that log dividends are generated by the geometric martingale:

$$d_{t+1} = \mu + d_t + \xi_{t+1}, \tag{2.11}$$

where μ is the trend growth in dividends. d_t is the log of dividends at time t, and ξ_{t+1} is a normal random variable with conditional mean zero and variance σ^2.

Assuming that period-t dividends are known when P_t is set, the present-value stock price is directly proportional to dividends:

$$P_t^{pw} = \kappa D_t, \tag{2.12}$$

where $\kappa = (e^r - e^{\mu+\sigma^2/2})^{-1}$ and r is a real interest rate.

Equation (2.12) is essentially a stochastic version of Myron Gordon's (1962) model of stock prices, where $P_t^{pw} = (e^r - e^{\mu})^{-1} D_t$. Equation (2.12) also requires $r > \mu + \sigma^2/2$.

Suppose that the function $B(D_t)$ as

$$B(D_t) = cD_t^{\lambda}, \tag{2.13}$$

Table 2.2 – Selected empirical studies in the literature related to periodically collapsing bubbles (PCBs).

Authors	Sample data	Method	Conclusion
Taylor and Peel (1998)	S&P 1871-1986.	Cointegration test.	No evidence of PCBs.
Hall et al. (1999)	Argentina's hyperinflation 1983M1-1989M11.	Markov-switching unit root test.	Existence of PCBs.
Psaradakis et al. (2001)	German hyperinflation 1918M121924M12.	Random-coefficient autoregressive models.	Existence of PCBs.
Bohl (2003)	S&P 1871-1995.	Momentum threshold autoregressive (MTAR) cointegration test.	No evidence of PCBs.
	S&P 1871-2001.	Momentum threshold autoregressive (MTAR) cointegration test.	Evidence of PCBs.
Payne and Waters (2007)	US equity REIT market 1972M01-2005M03.	Momentum threshold autoregressive (MTAR) cointegration test.	No evidence of PCBs.
	US equity REIT market 1972M01-2005M03.	Cointegration test.	Evidence of PCBs.

where λ is the positive root of the quadratic equation $\lambda^2\sigma^2/2 + \lambda\mu - r = 0$ and c is an arbitrary constant.

The basic stock price equation is now given as:

$$P(D_t) = P_t^{pw} + B(D_t) = \kappa D_t + cD_t^{\lambda}, \tag{2.14}$$

Equation (2.14) contains a bubble $B(D_t)$ (for $c \neq 0$). $P(D_t)$ is a function of dividends only and does not depend on time or any other extraneous variable. $B(D_t)$ is an example of an intrinsic bubble. Several empirical studies related to the intrinsic bubbles of Froot and Obstfeld (1991) are reported in Table 2.3 below.

2.2.4 Speculative Bubbles

Flood and Garber (1982) consider a persistent deviation of stock prices from the path determined by these fundamentals. Such deviations, if induced by self-fulfilling expectations, are called speculative bubbles. Harrison and Kreps (1978) argue that “investors exhibit speculative behavior if the right to resell a stock makes them willing to pay more for it than they would pay if obliged to hold it forever”. Empirical studies relating to speculative bubbles are presented in Table 2.4 below.

Table 2.3 – Selected empirical studies in the literature related to intrinsic bubbles.

Authors	Sample data	Method	Conclusion
Driffill and Sola (1998)	S&P 1900-1987	Markov-switching and intrinsic bubble model.	The occurrence of a bubble is low.
Ma and Kanas (2004)	S&P 1871-1996	Cointegration test.	Existence of intrinsic bubbles.
Black et al. (2006)	UK housing market 1973Q4-2004Q3	Intrinsic bubble model.	Existence of intrinsic bubbles.
Bidarkota and Dupoyet (2007)	S&P 1871-1988	Intrinsic bubble model with fat tails.	Existence of intrinsic bubbles.
Fraser et al. (2008)	New Zealand and UK housing market 1970Q1-2005Q4	Intrinsic bubble model.	Existence of intrinsic bubbles.
Chen et al. (2009)	S&P 1871-2004	Intrinsic bubble mode.	Existence of intrinsic bubbles.
Naoui (2011)	S&P 1871-2009	Intrinsic bubble mode.	Existence of intrinsic bubbles.

Table 2.4 – Selected empirical studies in the literature related to speculative bubbles.

Authors	Sample data	Method	Conclusion: a bubble?
Shiller (1981)	S&P 1871-1979.	Variance bounds test.	Yes.
Huang (1981)	Exchange rates 1973M03-79M03.	Variance bounds test.	Yes.
West (1987b)	S&P 1871-1980.	West's two step procedure.	Yes.
	Dow Jones Industrial Average 1928-1978.	West's two step procedure.	Yes.
Dezhbakhsh and Demirguc-Kunt (1990)	S&P 1871-1981.	West's two step procedure.	No.
	S&P 1871-1988.	West's two step procedure.	No.
Kearney and MacDonald (1990)	Australian Dollar-US Dollar 1984M01-86M12.	Variance bounds test.	No.

2.3 Bubble Detection Methods

This section discusses existing bubble detection approaches in the literature. Various bubble testing procedures have been developed, see Gürkaynak (2008) for a survey.

2.3.1 Variance Bounds Tests

The earliest work on testing for a bubble is the variance bounds test introduced by Shiller (1981) and Leroy and Porter (1981). Such tests were initially designed for assessing the reliability of the present value model. The market fundamental solution forms the basis of asset prices P_t:

$$P_t = E_t\left[\sum_{i=1}^{\infty}\left(\frac{1}{1+R}\right)^i D_{t+i}\right]. \tag{2.15}$$

The *ex post* rational price (P_t^*) is the present value of actual dividends:

$$P_t^* = \sum_{i=1}^{\infty}\left(\frac{1}{1+R}\right)^i D_{t+i}. \tag{2.16}$$

Let ε be an unobserved variable with a zero mean.

$$P_t^* = \sum_{i=1}^{\infty}\left(\frac{1}{1+R}\right)^i \left[E_t(D_{t+i}) + \varepsilon_i\right] = P_t + \sum_{i=1}^{\infty}\left(\frac{1}{1+R}\right)^i \varepsilon_{t+i}. \tag{2.17}$$

The variance bounds test relies on the assumption of:

$$V(P_t^*) = V(P_t) + \varphi V(\varepsilon_t) \geq V(P_t), \tag{2.18}$$

where φ is $\left[\frac{1}{(1+R)}\right]^2 / \left[1 - \left(\frac{1}{1+R}\right)^2\right]$.

Equation (2.18) places an upper bound on the variance of the observed price and suggests that the variance of the ex post rational price should be no smaller than the variance of observed prices. A violation of the variance bound indicates the presence of bubbles.

2.3.2 West's Two-Step Tests

West (1987a) proposed a simultaneous test for both model specification and bubbles by using a three-equation model. The first equation is related to Euler equations with no-arbitrage condition, which yields the discount rate. The second equation describes the dividend process by an ARIMA model. The third is used for checking the existence of bubbles. When such a procedure is applied to US stock data for 1871-1980, the test rejects the no bubble hypothesis.

The following notation is adopted from Gürkaynak (2008). Suppose that dividends are exogenous and follow a stationary AR(1) process of the form:

$$D_t = \phi D_{t-1} + u_t^D, \tag{2.19}$$

The autoregressive parameter can be estimated by OLS. The market fundamental stock price is:

$$P_t^f = \sum_{i=1}^{\infty} \left(\frac{1}{1+R} \right)^i E_t(D_{t+i}|\Omega) = \bar{\beta} D_t, \tag{2.20}$$

where

$$\bar{\beta} = \frac{\phi/(1+R)}{1-\phi/(1+R)}.$$

.

The actual price may consist of a bubble B_t. P_t is the sum of the market fundamental price, which is given as:

$$P_t = \beta D_t + B_t. \tag{2.21}$$

If a bubble exists, which is correlated with dividends, the estimate of β in Equation (2.21) is biased. If dividends can be represented by an autoregressive (AR) process, the relationship between dividends and stock prices can be estimated using Equation (2.19) and Equation (2.21). This is a test under the null hypothesis that when there are no bubbles, the actual relationship will not differ from the constructed fundamental prices, i.e., $\beta = \bar{\beta}$.

If a Hausman coefficient restriction test rejects the equality of $\bar{\beta}$ and a estimate of β indicates the presence of a bubble.

2.3.3 Unit Root/Cointegration Based Tests

Cointegration tests also play an important role in detecting bubbles. Many different cointegration based tests have been applied to examine the phenomenon of bubbles. One way of detecting bubbles is to use the residual-based cointegration tests to investigate the presence of bubbles. These approaches commonly apply the unit root tests to the residuals from the regression of the stock prices and dividends. If the residual has a unit root, then the stock prices and dividends are not cointegrated. In other words, the occurrence of bubbles implies that there is no long run relationship between stock prices and dividends. Hence, if stock prices and dividends are cointegrated, which may be taken as evidence of the absence of bubbles. Similarly, another approach to examine the existence of bubbles is to investigate the dividend-price ratio by unit root tests. A unit root of the dividend-price ratio implies the existence of bubbles. Such a test may be interpreted as a way to explore the long-run relationship between stock prices and dividends.

As discussed in Lim and Phoon (1991), the following relationship between stock and a rational bubble can be written as Equation (2.8):

$$P_t = P_t^f + B_t,$$

where P_t is the real stock price, F_t is the fundamental valuation of the stock at time t, B_t is the rational bubble.

The fundamental valuation of P_t^f is based on the anticipated stock dividend stream. P_t^f can be written as $\beta d_t + \theta_t + \upsilon_t$, where β is the discount factor on an infinite series of constant expected dividend at time t, θ_t is a discounted sum of expected dividend changes, and υ_t is the present value of other unobserved market fundamentals. Then the cointegrating model of stock price and dividends can then be written as:

$$P_t - \beta d_t = B_t + \theta_t + \upsilon_t \tag{2.22}$$

If the unobservable variables in the market fundamentals are stationary in levels, and dividends are stationary in first-difference, then the linear combination $P_t - \beta d_t$ is stationary if the bubble B_t does not exist. P_t and d_t are cointegrated with cointegrating vector $(1, -\beta)$ if θ_t and υ_t stationary and B_t is zero. Therefore, a long-run cointegration relationship between asset prices and fundamentals indicates the existence of (rational) bubbles.

2.3.4 Long Memory Models

A number of studies apply long memory models to test for stock price bubbles, for example see Koustas and Serletis (2005), Cuñado et al. (2005) and Cuñado et al. (2012). The presence of long memory in a series can be estimated by the long memory parameter d. Different semi-parametric and parametric methods have been proposed to estimate d in the literature.

Inferences on bubbles are based from the estimated value of persistence parameter, d, of the price-fundamental ratio. The bubble-like behavior in asset prices can be tested using the estimated values of the fractional integrating parameter, d, of the (log) price-fundamental ratio.

- If $0 \leq d \leq 0.5$, the series is covariance-stationary and mean reverting, implying no rational bubbles.
- If $0.5 \leq d \leq 1$, the series is non-stationary and mean reverting process, implying no rational bubbles.
- If $d \geq 1$, the series is a non-stationary explosive and not mean-reversion process, implying rational bubbles.

Hence, the value of d then shows the presence or absence of a bubble in the series.

2.3.5 Phillips Type Right-tailed Unit Root Tests

The recent development in right-tailed unit root tests of Phillips, Wu, and Yu (2011, PWY) and Phillips, Shi, and Yu (2015, PSY) has shown great power in detecting financial bubbles.

PWY approach

A highlight of the Phillips, Wu, and Yu's (2011) approach is the ability to capture the characteristics of periodically collapsing bubbles by Evans (1991), and the PWY is effective in detecting a single bubble episode. The testing procedure is known as the SADF test and is implemented as follows. For each time series x_t, we apply the Augmented Dickey-Fuller (ADF) test for a unit root against the alternative of an explosive root (right-tailed). The following autoregressive specification for x_t is estimated by least squares:

$$x_t = \mu_x + \delta x_{t-1} + \sum_{j=1}^{J} \phi_j \Delta x_{t-j} + \varepsilon_{x,t}, \qquad \varepsilon_{x,t} \sim \text{NID}(0, \sigma_x^2), \tag{2.23}$$

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 $H_1 : \delta > 1$. The above equation is estimated repeatedly using subsets of the sample data incremented by one additional observation at each pass in the forward recursive regression. Thus the SADF test is constructed by repeatedly estimating the ADF test. Let r_w be the window size of the regression. The window size $r_w (r_w = r_2 - r_1)$ expands from r_0 to 1, where r_0 is the smallest sample window width fraction and 1 is the largest window fraction (the full sample). 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 ADF statistic for a sample that runs from 0 to r_2 is therefore denoted by $ADF_0^{r_2}$. The SADF statistic is defined as the sup value of the ADF statistic sequence:

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

The SADF test statistic cannot locate the ordination and collapse dates of a bubble. To identify the origin and the collapse dates, we can compare the recursive test statistic ADF_r against the relevant right-tailed critical values. If r_e is the origination date and r_f is the collapse date, we can construct estimates of these dates as follows:

$$\hat{r}_e = \inf_{r_2 \in [r_0, 1]} \left\{ r_2 : ADF_{r_2} > cv_{r_2}^{adf} \right\},$$

$$\hat{r}_f = \inf_{r_2 \in [\hat{r}_e, 1]} \left\{ r_2 : ADF_{r_2} < cv_{r_2}^{adf} \right\}.$$

The ADF statistic and its corresponding critical value are used for dating the origination and termination dates of a bubble.

Rolling-window approach

Phillips, Wu, and Yu (2011) also mention an alternative testing method to detect the presence of of bubbles-the rolling window approach. The rolling window approach is used as a robustness check compared with the PWY approach in Phillips, Wu, and Yu (2011). The feature of this new testing method is that a rolling window of fixed length is chosen for the whole sample. The ADF test statistic is calculated over a fixed size $r_w = r_0$ for all estimations. The stating point r_1 and ending point r_2 are incremented by one observation at each time, yielding an ADF statistic denoted as $ADF_{r_1}^{r_2}$. The corresponding rolling ADF (RADF) statistic can be defined as the supremum $ADF_{r_1}^{r_2}$ statistic cross all possible windows.

Bubble origination dates (r_e) and collapse dates (r_f) estimated under the rolling-window procedure are constructed as:

$$\hat{r}_e = \inf_{s \geq w} \left\{ s : DF_s > cv_{\beta_n}^{adf} \right\},$$

$$\hat{r}_f = \inf_{s \geq \hat{r}_e + L_T} \left\{ s : DF_s < cv_{\beta_n}^{adf} \right\}.$$

PSY approach

The Phillips, Shi, and Yu (2015, PSY) is also extended from an early bubble detection test of Phillips, Wu, and Yu (2011). The PSY test has great power in detecting the presence of multiple bubbles. The martingale null with an asymptotic drift is specified as:

$$H_0 : y_t = dT^{-\eta} + y_{t-1} + \varepsilon_t, \qquad \varepsilon_t \sim \text{NID}(0, \sigma^2), \tag{2.24}$$

where d is a constant, the localizing coefficient η is greater than $1/2$ and T is the sample size. The alternative hypothesis is a mildly explosive process:

$$H_1 : y_t = \delta_T y_{t-1} + \varepsilon_t,$$

where $\delta_T = 1 + cT^{-\theta}$ with $c > 0$ and $\theta \in (0, 1)$.

The following regression model is estimated:

$$\Delta y_t = \hat{\alpha} + \hat{\beta} y_{t-1} + \sum_{i=1}^{k} \hat{\gamma}_i \Delta y_{t-i} + \hat{\varepsilon}_t, \tag{2.25}$$

where $\hat{\alpha}$ is an intercept and k is optimum lag length.

The generalized sup ADF (GSADF) test relies on repeated estimation of the ADF test regression model on subsamples of the data in a recursive fashion. The window size r_w expands from r_0 to 1, where r_0 is the minimum window size. The end point r_2 varies from r_0 to 1 and the starting point r_1 varies from 0 to $r_2 - r_0$. The GSADF statistic is the largest ADF statistic over the range of r_1 and r_2:

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

The GSADF test is used for assessing explosive behavior for the entire sample period, however, it does not provide the origination and termination dates of identified bubble episodes. In order to provide a real-time monitoring of market exuberance, we will use the backward SADF (BSADF) test. The BSADF statistic is defined as the sup value of the ADF statistic sequence:

$$BSADF_{r_2}(r_0) = \sup_{r_1 \in [0, r_2 - r_0]} ADF_{r_1}^{r_2}$$

.

The BSADF statistic and its corresponding critical value are used for dating the origination and termination dates of a bubble as follows:

$$\hat{r}_e = \inf_{r_2 \in [r_0, 1]} \left\{ r_2 : BSADF_{r_2}(r_0) > cv_{r_2}^{\beta_T} \right\},$$

$$\hat{r}_f = \inf_{r_2 \in [\hat{r}_e, 1]} \left\{ r_2 : BSADF_{r_2}(r_0) < cv_{r_0}^{\beta_T} \right\}.$$

2.4 Previous Empirical Studies of Bubbles

Section 2.4 overviews the existing empirical studies testing for bubbles in several key financial areas including stock, housing and exchange rate markets.

2.4.1 Stock Price Bubbles

There has been increasing interests in exploring the presence of bubbles in stock markets with the majority of studies concentrating on US evidence.

The Mississippi Bubble and the closely related South Sea Bubble in 1719-20 are the two most famous and earliest example of stock price bubbles in history. Empirical studies have tested the Mississippi Bubble or the South Sea Bubble as these episodes have generated considerable interest in the literature. Neal (1990) carries out a statistical analysis using the method of Blanchard and Watson (1982) and concludes that the Mississippi share price contains a rational bubble from mid-July 1719 to the end of November 1719 and the South Sea share price contains a rational bubble between 23 February 1720 and 15 June 1720 only. Carlos et al. (2002) examine the Royal African Company share prices during the South Sea episode and find no significant evidence to support the existence of a bubble. They call into question the arguments by Chancellor (2000) that the South Sea Bubble was the result of mania and speculative excesses. However, Garber (2001) claims that he provides market fundamental explanations for the three most famous bubbles: Tulipmania, the Mississippi Bubble and the South Sea Bubble, which seems to provide no evidence of bubbles for these historical episodes. Velde (2009) concludes that the Mississippi Company is overvalued.

Japan's stock market bubble during the 1980-90s is an interesting recent example. One of the earliest studies on Japan's stock market is by Hardouvelis (1988), who uncovers evidence of rational bubbles in stock markets of Japan and the United States before the October 1987 crash. Ueda (1990) also reports that the sharp rise in the Japan's stock market in the 1980s is due to one of the following three causes: bubbles, declines in risk premium, and expectation of land price inflation. An early study of French and Poterba (1991) suggests that accounting differences cannot explain the doubling of the Japanese price-earnings ratio in 1986 and its decline in 1990. An empirical model is also suggested by Lee (1995) which decomposes asset prices into fundamental and non-fundamental bubble components, which finds significant deviations of the Japanese stock and land prices from the fundamentals (Japan's GNP) between 1955 and 1992. Sato (1995) argues that the low nominal interest rate caused by loose monetary policy contributed to the 1980s stock bubble in Japan. Additionally, Ito and Iwaisako (1996) explore the behaviour of stock and land prices during Japan's bubble economy. They argue that asset price increases from mid-1987 to mid-1989 are in line with the changes of fundamentals, and the remarkable rise in stock prices in the latter 1989 and land prices in 1990 are not explained by either fundamentals or rational bubbles. Chan et al. (1998) show no overwhelming evidence of bubbles in Japan's stock market using the duration dependence and conditional skewness tests of McQueen and Thorley (1994) based on weekly and monthly stock market index from January 1975 to April 1994. The findings from Chung and Lee (1998) also suggest that the deviations from Japanese stock market fundamentals are quite substantial in the late 1980s, suggesting the existence of a bubble. An influential paper on Japan's asset price bubble by Chirinko and Schaller (2001) also provides evidence to support the existence of a bubble in Japan's stock market based on orthogonality tests. Binswanger (2004) also argues that real activity shocks explain a small proportion of the variation in Japan's stock price in the 1983-1999. A Bayesian approach proposed by Asako and Liu (2013) identifies speculative bubbles in Japan's stock market from the period May 1979 to January 2010 and more interestingly, concluding that Japan's stock market around 1990 was truly a bubble.

There has been increasing interests in exploring the presence of bubbles in stock market with the majority of studies focusing on the US stock market. Campbell and Shiller (1987) test for cointegration relationships between stock price and dividend using US S&P stock index from 1871 to 1986, given that both variables are stationary in first difference. The proposed cointegration tests used in their paper are ambiguous as results are mixed. They also find that the choice of discount rate may have an influence on the cointegration test results, which violates the assumption of a constant discount rate.

Diba and Grossman (1988) find that there is insufficient evidence to support the existence of rational bubbles in stock prices. Based on their empirical results, they conclude that the stock price and dividends are cointegrated and they rule out the rational bubble in US stock prices.

Craine (1993) applies a standard Augmented Dickey-Fuller (ADF) test to the log dividend-price ratio of S&P 500 data ranging from 1876 to 1988. He found a unit root in the log dividend-price ratio, which violates the restriction of no rational bubble.

Han (1996) applies the Canonical Cointegration Regression (CCR) procedure and Johansen's maximum likelihood method to investigate the long-run relationship between the stock price and dividends of US market by testing the present value model. The results obtained from the CCR procedure are in line with the Johansen's ML results and in favour of the presence of rational bubbles.

Lamont (1998) finds evidence in favor of the existence of bubbles of the log dividend-price ratio in US data using Dickey-Fuller tests. However, the Horvath and Watson (1995) cointegration tests produce strong cointegration relationship between prices and dividends. The results obtained from the unit root tests are not consistent with the results from the cointegration tests.

Koustas and Serletis (2005) utilise an autoregressive fractionally integrated moving average model (ARFIMA) with volatility modelling (ARCH, GARCH) to investigate the degree of persistence of the log dividend yield. Instead of using integer orders of integration, they use fractional integration techniques, which allow more flexible modelling. A good feature of this approach is that it is not sensitive to data frequency. They find strong evidence of rejections of a unit root in the log dividend yield and this result supports the non-existence of bubbles of null hypothesis.

Cuñado et al. (2005) analysis the NASDAQ stock market index between 1994 and 2003 by utilizing fractional integration techniques based tests of Robinson (1994) and Robinson (1995). They obtain evidence to suggest that the sampling frequency adopted in the analysis is crucial to determine the existence of bubbles. In their paper, they use monthly, weekly, and daily data of the NASDAQ index to examine the presence of bubbles. Based on empirical evidence, they are in favor of the presence of bubbles only when using the monthly data on the price-dividend ratios.

Chang et al. (2007) have re-investigated the presence of rational bubbles in the US stock market by applying both the Johansen cointegration test and the Bierens nonparametric cointegration test for the period from 1871 to 2002. Most of empirical studies in the literature have employed cointegration tests as a measure to check the existence of bubbles. The results from the Johansen cointegration test support the presence of rational bubbles while the results from Bierens nonparametric cointegration test favour the absence of rational bubbles in the US stock market.

Time series based analysis has been widely adopted to investigate the evidence of bubbles phenomena, but the empirical results remain mixed. Cerqueti and Costantini (2011) utilise a panel data approach that combines panel unit root and panel cointegration tests with the flexibility of allowing multiple structural breaks in each series to examine the phenomena of rational bubbles. They present empirical evidence of a rational bubble hypothesis in international stock markets for 18 OECD countries using the model proposed by Campbell (2000). The key feature of their approach is that the panel data approach gives a global analysis of rational bubble in international stock markets. Perron (1989) finds that the existence of structural breaks could strongly influence the ability of rejecting the unit root null hypothesis. The intuition behind this is that many empirical studies could lead to biased results by simply ignoring the structure breaks in the data. Empirical results from Cerqueti and Costantini (2011) are in favour of the existence of bubbles in all 18 OECD countries.

Similarly, Frömmel and Kruse (2012) also test for rational bubbles for the S&P 500 stock index using the fractional integration technique by Sibbertsen and Kruse (2009) that allows both long memory and changing persistence and the unit root tests of Demetrescu et al. (2008). They find a changing long memory parameter of the dividend-price ratio during the whole sample period. Their results confirm that there is a structure break in July 1991 in the log dividend-price ratio. The log dividend-price ratio has significant long memory before July 1991 while it contains a rational bubble after that as the unit root test suggests non-stationarity.

Many studies utilise the PSY tests to stock markets and some selected studies are summarized here. Chen et al. (2015) investigated the health care sector of the US, UK, and German stock markets. The identified bubbles corresponded to the Dotcom and pre-subprime crisis periods. Bohl et al. (2013) tested for bubbles in German renewable energy stocks between 2004 and 2011. Almudhaf (2017) also explored explosive behaviour for a group of African stock markets. In addition, Escobari and Jafarinejad (2016) adopted the PWY and PSY tests to test for the existence of bubbles in various Real Estate Investment Trust (REIT) indices including Equity REITs, Mortgage REITs and Hybrid REITs. Similarly, Chang et al. (2016) make use of the PSY bubble detection test to the BRICS stock markets during the 1990-2013. Homm and Breitung (2012) applied the PWY tests to the Nasdaq composite, the S&P 500, the Nikkei225, the Hang Seng and the Shanghai stock indices.

2.4.2 Housing Bubbles

There is a large literature examining the existence of house price bubbles especially for the US. For example, Malpezzi (1999) find no evidence of bubbles in US

Metropolitan areas as house prices and income are cointegrated. Case and Shiller (2003) argue that fundamentals play a crucial role in explaining much of the rapid increase in the housing market. In particular, income growth explains the house price appreciation in most areas and a low interest rate also contributes to the recent increase in house prices. Similarly, McCarthy and Peach (2004) and Himmelberg et al. (2005) conclude that economic fundamentals such as low interest rates and high income growth could explain house prices growth in the early 2000s. Therefore, Case and Shiller (2003), McCarthy and Peach (2004) and Himmelberg et al. (2005) deny the existence of a bubble in the US housing market.

However, there is other empirical evidence that supports the existence of a housing bubble in the US in the literature. Gallin (2006) finds no evidence of a long-run cointegration relationship between house prices and income in US using 27 years of US national-level data. As standard cointegration tests are known to have low power in small samples, Gallin (2006) applied several panel cointegration tests to a panel of 95 US Metropolitan areas over a 23 year period at the city-level and also found no evidence of cointegration. In both cases, a finding of no cointegration relationship indicates the presence of bubbles. Zhou and Sornette (2006) investigated the existence of US housing bubbles at the regional and national levels using quarterly data between 1993 and 2005 and concluded that 22 States exhibited evidence of a bubble based on a definition as a 'faster-than-exponential price growth'. Mikhed and Zemčík (2009) also applied a panel test for the price-rent ratio in 23 US Metropolitan areas from 1978 to 2006 and concluded that there was evidence of a bubble.

Holly et al. (2010) used Moon and Perron (2004) and Pesaran (2007) panel tests to consider the relationship between real house prices and real per capita disposable incomes in the US at the State level using annual data 1975-2003 and found little evidence of house price bubbles with a few exceptions (e.g., California, New York, Massachusetts, Connecticut, Rhode Island, Oregon and Washington). Empirical results from Kivedal (2013) suggest that there was a bubble in the US housing market prior to the 2007 subprime financial crisis. Nneji et al. (2013) examined the presence of intrinsic bubbles of a Froot and Obstfeld (1991) type or rational speculative bubbles of a Blanchard and Watson (1982) type in the residential property market in the US between 1960 and 2011. They split the data into two periods (1960-1999) and (2000-2011) and found an intrinsic bubble for the first period and a rational speculative bubble for the second period only. In this paper, they tested to see if house prices are driven solely by changes in fundamentals (rents) for the intrinsic bubble model. Escobari et al. (2015) proposed a new test to identify house price bubbles, which explored a specific feature of the market such that low tier house prices should appreciate more during the upswing of a growing boom and fall faster during the bust, and found evidence of bubbles in 15 US Metropolitan Statistical Areas.

A number of recent studies apply the popular PSY tests to housing markets. Yiu et al. (2013) applied the multiple bubble detection test to identify real estate bubbles in the Hong Kong residential property market including the overall market, the mass segment and the luxury segment. In particular, they found the existence of the well-known real estate bubble in 1997 for all three markets. Pavlidis et al. (2016) examined the explosive behaviour of three housing market indicators (real house prices, price-to-income ratios, and price-to-rent ratios) for 22 countries. They also proposed a panel version of the PSY test to exploit large cross-sectional dimension of the dataset. Engsted et al. (2016) also conducted an econometric analysis of housing bubbles for 18 OECD countries from 1970 to 2013. Greenaway-McGrevy and Phillips (2016) and Shi et al. (2016) investigated the presence of bubbles across various regional markets in New Zealand and Australia, respectively.

2.4.3 Exchange Rate Bubbles

A number of studies have tested for bubbles in the exchange rates using different econometric methodologies. Of the existing studies, most apply the variance bounds, unit root or cointegration tests. Using a monetary model approach, Huang (1981) applied the variance bounds test of Shiller (1980) and Leroy and Porter (1981) to test for the presence of bubbles in the US Dollar-Sterling, US Dollar-Mark, and Sterling-Mark exchange rates from March 1973 to March 1979 and rejected the no bubble hypothesis. Evans (1986) found evidence to support the presence of bubbles in the Sterling-dollar exchange rate between 1981 and 1984. Similarly, Meese (1986) provided evidence of bubbles for the Dollar-Deutsche Mark and Sterling-Dollar exchange rate using the monthly data between 1973 and 1982. However, the analysis from West (1987b) seemed to provide no evidence of bubbles in the US Dollar-Mark exchange rate during 1974 to 1984. Woo (1987) studied the bilateral exchange rate of the US Dollar with the currencies of Germany, France, and Japan using a portfolio balance approach. Kearney and MacDonald (1990) adopted a similar approach with Huang (1981) and concluded no bubbles in Australian/US dollar exchange rate over the period January 1984 to December 1986. Kaskarelis (1995) examined the presence of bubbles in the franc-mark exchange rate covering the period January 1981 to December 1992 and the obtained results which seemed to support the existence of bubbles. Wu (1995) applied the Kalman filter technique to estimate and tested for exchange rate bubbles between the US Dollar, the British Pound, the Japanese Yen and the Deutsche Mark using the monthly data over 1974-1988. However, Wu (1995) found no significant evidence of bubbles in these exchange rates. The empirical analysis from Charemza (1996) seemed to suggest the evidence of a rational bubble in the foreign exchange rate in Poland during the period 1988 to 1990. Elwood et al. (1999) made use of state-space models and Monte Carlo experiments to explore the presence of a stochastic rational bubble in the Japanese and German exchange rates over the period December 1984 to November 1998. According to the theory of uncovered interest parity, a series under rational expectation is supposed to be white noise. Elwood et al. (1999) therefore inspected this condition for evidence of bubbles. A finding of a deviation from white noise implies the existence of a stochastic rational bubble. Their results suggest a bubble burst between the end of March and the end of April of 1990, which coincides with economic turmoil in Japan and Germany.

Many studies have tested for evidence of price bubbles during hyperinflation using Cagan's (1956) model. These studies access the validity of the Cagan model and interpret deviations from the fundamental as evidence of bubbles. However, those deviations may be caused by model misspecification and the possibility of model misspecification is frequently ignored. Chan et al. (2003) suggested and extended a testing procedure of Durlauf and Hooker (1994) and Hooker (2000) to differentiate between model misspecification and the presence of bubbles. They examined the presence of price and exchange rate bubbles during the interwar European hyperinflation of Germany (September 1920-May 1923), Hungary (September 1921 - February 1924) and Poland (September 1920 - December 1923), and concluded that no evidence of price and exchange rate bubbles can be found in these three countries. Jarrow and Protter (2011) developed a new model for exchange rate bubbles using the martingale based bubble approach of Jarrow et al. (2010). This new model provides new insights into exchange rate bubbles and can be used for testing for bubbles.

In theory, a finding of a cointegrating relationship between exchange rate and fundamentals may be interpreted as the evidence of no bubbles. MacDonald and Taylor (1991) used cointegration tests to investigate the validity of the monetary model. If both the no bubbles hypothesis and the monetary model are valid, fundamentals and exchange rates must be cointegrated. Using this fact MacDonald and Taylor (1991) find evidence which supports the joint hypothesis. The empirical results from MacDonald and Marsh (1997) lent support to a long run cointegration relationship between exchange rate (the UK pound, German mark and Japanese yen against the US dollar) and fundamentals based on a modified version of PPP. Based on a panel cointegration test, Mark and Sul (2001) found a cointegration relationship between the nominal exchange rate and monetary fundamentals in 19 countries during January 1973 to January 1997. Jirasakuldech et al. (2006) examined the existence of rational speculative bubbles in the bilateral exchange rates of the British pound, the Canadian dollar, the Danish krone, the Japanese yen and the South African rand against the US Dollar covering the period from January 1989 to December 2004. Three different bubble detection procedures have been used: the augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) unit root tests, the Johansen's multivariate cointegration test and the duration dependence test of McQueen and Thorley (1994). All three tests provide firm evidence of no rational speculative bubbles in theses currency pairs. Chang and Su (2014) explored the long run relationship for several Pacific Rim countries using the Johansen (1988) cointegration test and cointegration tests with structural breaks of Hansen and Seo (2002). The Johansen cointegration test fails to reject the null hypothesis of no cointegration except Taiwan-US while the cointegration test with structure breaks rejects the null hypothesis of cointegration for Canada, Japan, Korea, and Thailand, with respect to the US.

A number of studies attempted to explore the evidence of nonlinear cointegration relationship between nominal/real exchange rate and economic fundamentals using the Alternating Conditional Expectation (ACE) algorithm (e.g., Chinn (1991), Meese and Rose (1991), Ma and Kanas (2000), Torres (2007) and Tang and Zhou (2013). Granger and Hallman (1991) suggested and developed a two stage testing procedure to test for a nonlinear cointegration relationship using the ACE algorithm and the ACE was firstly introduced by Breiman and Friedman (1985). Both Chinn (1991) and Meese and Rose (1991) found a nonlinear cointegration relationship between the nominal exchange rate and fundamentals and the results may be interpreted as non-existence of bubbles. Similarly, Ma and Kanas (2000) employed a nonparametric nonlinear cointegration approach and a nonlinear Granger causality methodology to test for a nonlinear cointegration relationship between fundamentals and exchange rates for Netherlands-Germany and France-Germany. For Netherlands-Germany, they find a nonlinear cointegration relationship which indicates no bubbles. For France-Germany, they find no evidence of a nonlinear cointegration but do find a nonlinear Granger causality from French money to the exchange rate. Results from a fractional ARIMA model suggest that these nonlinearities are not due to bubbles. Torres (2007) also studied the potential nonlinear cointegration relationship between three ERM exchange rate (French Franc/Deutsche Mark, Belgian Franc/Deutsche Mark, Danish Krona/Deutsche) and fundamentals. Instead of using nominal exchange rate, Tang and Zhou (2013) assessed nonlinear aspects in the determination of the real exchange rate. They investigated the nonlinear relationship between the real exchange rates of Chinese Yuan and Korean Won, and economic fundamentals over the period 1980Q1-2009Q4 and they find no evidence to indicate the presence of bubbles.

A series of papers have adopted a regime switching approach to examine the presence of bubbles in exchange rates. Van Norden (1996) investigated the existence of speculative bubbles in exchange rates of the Japanese yen, the German mark and the Canadian dollar from 1977 to 1991 by applying a new regime switching test. The presence of bubbles displays a particular kind of regime-switching behavior by implying some coefficient restrictions on a simple switching-regression model of the exchange rate. Empirical results are sensitive to the choice of exchange rate fundamentals and measurement of exchange rate innovations. Ferreira (2006) applied the Markov-switching ADF test of Hall et al. (1999) to explore the presence of bubbles in exchange rates for Canada, France, Germany and the UK and rejected the null hypothesis of bubbles. Maldonado et al. (2012) proposed a new bubble model to the exchange rate of the Brazilian real to the US Dollar during March 1999 to February 2011. The new model is an extension of the Van Norden (1996) approach which allows for a non-linear specification between the innovation and the bubble size in the survival regime. This allows them to test for rational expectations. Using the same model, Maldonado et al. (2014) further examined the occurrence of bubbles in the exchange rate of Brazil, Russia, India, China and South Africa (“BRICS” countries) against the US Dollar. Panopoulou and Pantelidis (2015) also developed two and three-state regime switching models based on the work of Van Norden (1996) and provided evidence of periodically collapsing bubbles in the Sterling-Dollar exchange rate from January 1973 to January 2011.

More recently, several studies apply the PSY tests to exchange rates. For example, Bettendorf and Chen (2013) and Jiang et al. (2015) empirically examined the explosive behaviour in the Sterling-dollar and Chinese RMB-dollar exchange rates at the monthly frequency, respectively. Steenkamp (2018) tested for explosiveness in G11 currencies at a daily frequency. The volatility of daily exchange rates tend to vary over time. Non-stationary volatility could cause a size distortion in unit root tests. Therefore Steenkamp (2018) utilised a wild-bootstrap approach of Harvey et al. (2016) to take into account the possibility of non-stationary volatility to a range of highly traded exchange rates.

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Chapter 3

Are There Bubbles in Exchange Rates? Some New Evidence from G10 and Emerging Market Economies.

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Addendum

In this paper, 'the intercept in the null hypothesis' should be interpreted as 'the intercept in the regression model'. We apologize for this mistake.

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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:

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

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:

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

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:

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

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%).

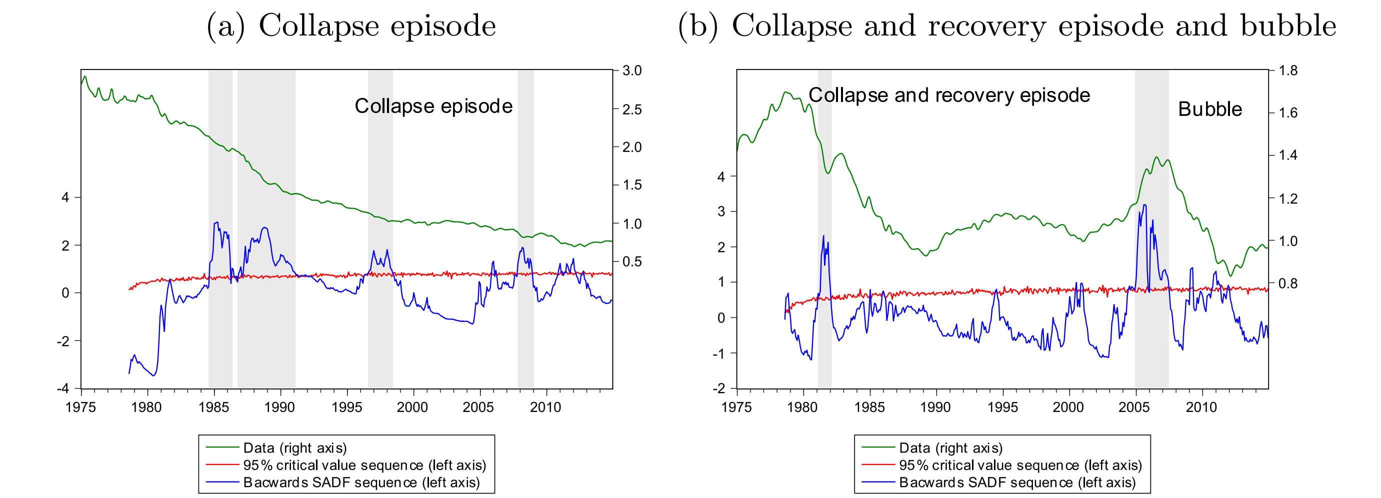

Fig. 1. Examples of the collapse episode, the collapse and recovery episode and the bubble. (a) Collapse episode. (b) Collapse and recovery episode and bubble. (For interpretation of the references to color in this figure, the reader is referred to the web version of this article.)

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.

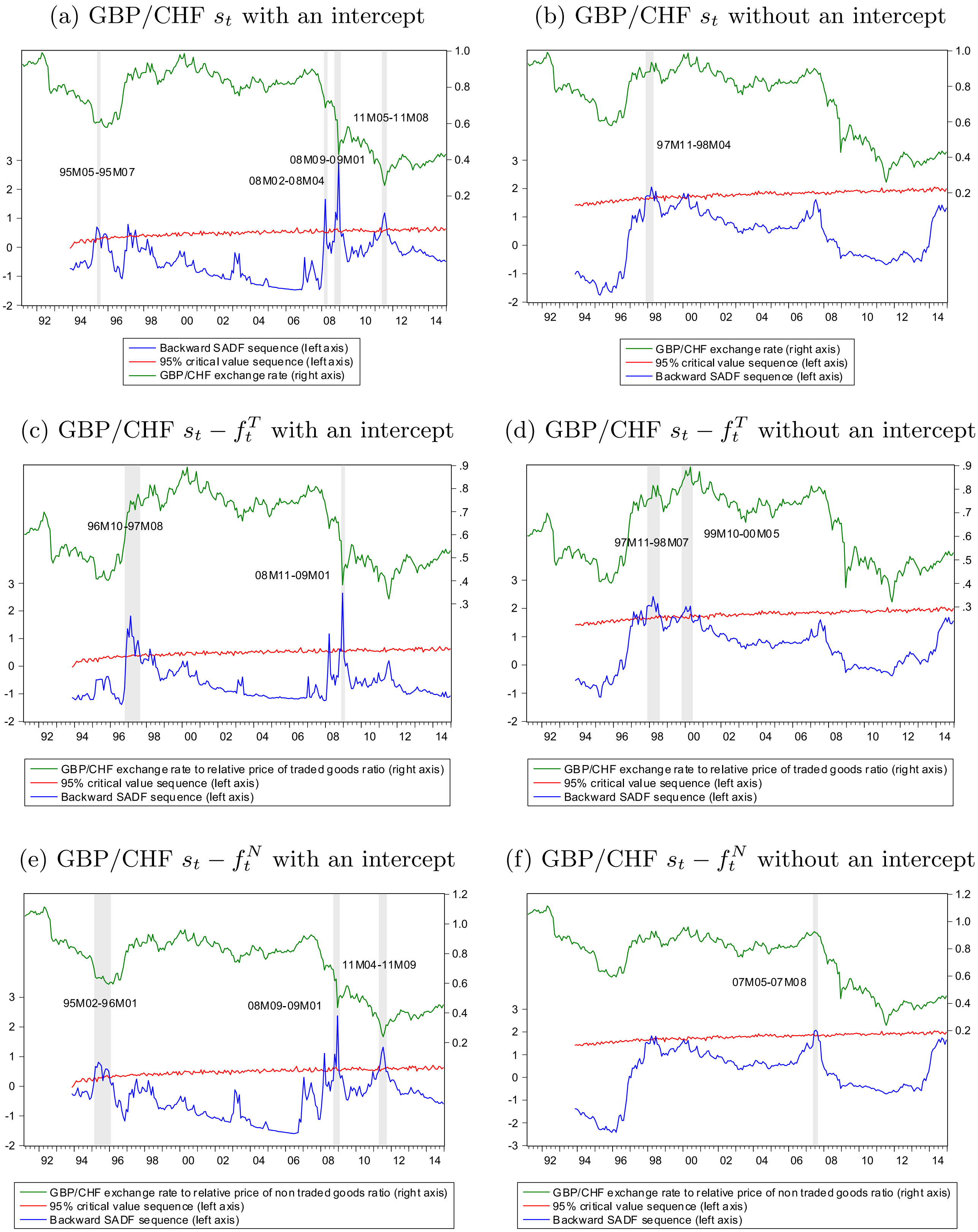

Fig. 2. Dating strategy for GBP/CHF nominal exchange rate s_t, the relative ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$ and the relative ratio of the exchange rate to the non-traded goods fundamental $s_t - f_t^N$. (a) GBP/CHF s_t with an intercept. (b) GBP/CHF s_t without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

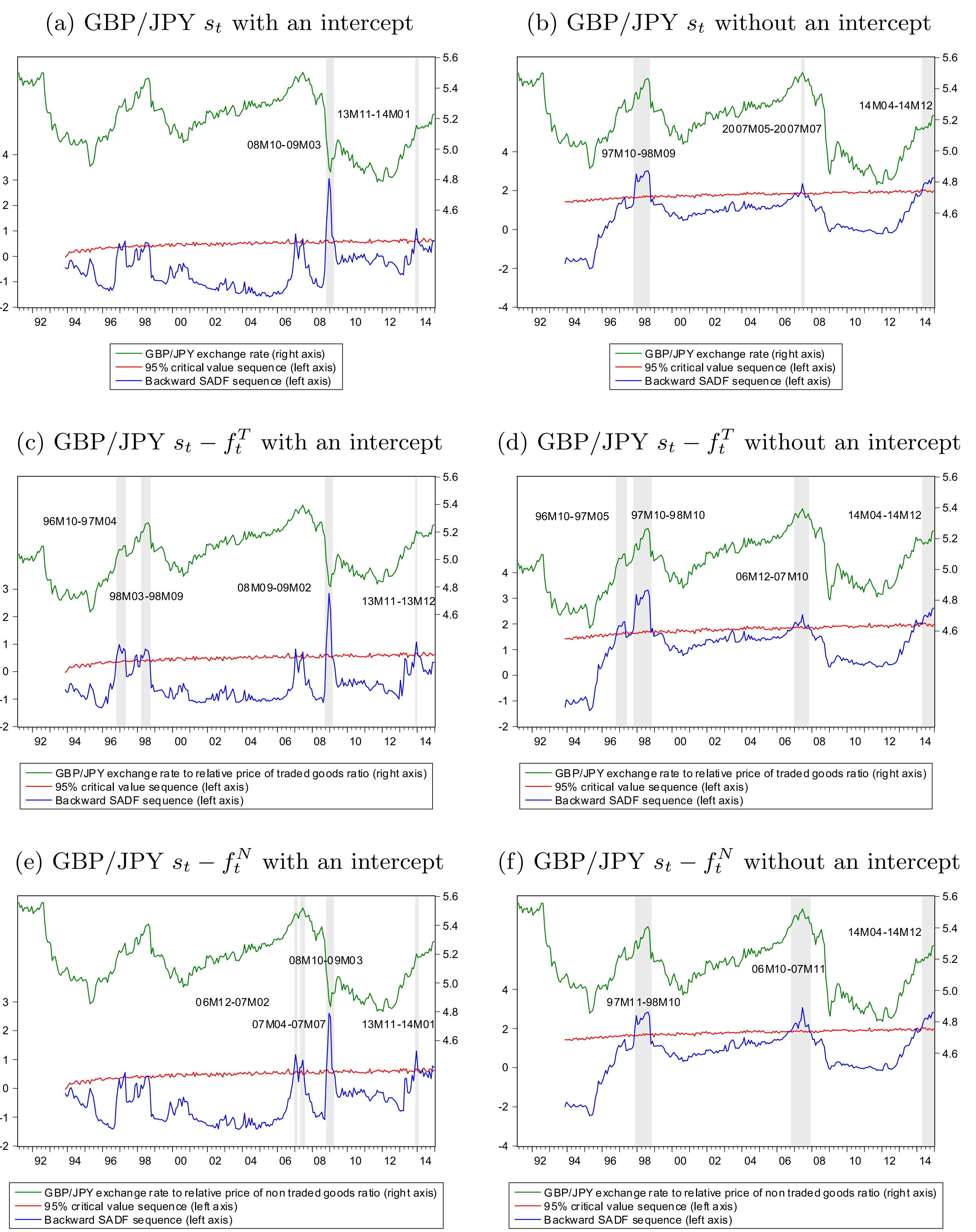

Fig. 3. Dating strategy for GBP/JPY nominal exchange rate s_t, the relative ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$ and the relative ratio of the exchange rate to the non-traded goods fundamental $s_t - f_t^N$. (a) GBP/JPY s_t with an intercept. (b) GBP/JPY s_t without 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 08M01-08M05, 10M08-10M12	2.4169	97M10-98M02
USD/IDR				
s_t	9.1720***	94M08-96M08, 96M11-98M09 13M07-14M02	15.7484***	93M11-98M02, 98M05-98M08 13M08-14M12
$s_t - f_t^N$	11.0643***	95M04-98M09, 13M08-14M02	4.6668***	94M06-98M02, 98M05-98M08 13M09-14M12
$s_t - f_t^T$	8.6602***	97M07-98M02, 08M03-08M08 13M08-13M09	2.0424	97M10-98M01
USD/KRW				
s_t	9.9778***	95M03-95M08, 96M12-98M02 08M08-08M11, 09M01-09M02	4.5216***	93M11-95M04, 96M05-98M02
$s_t - f_t^N$	9.5177***	95M02-95M08, 97M01-98M03 04M11-05M05, 05M12-06M06 08M08-08M11, 09M01-09M02	2.3598	93M11-94M05, 97M02-98M02
$s_t - f_t^T$	9.9778***	95M03-95M08, 97M09-98M02 08M08-08M11	2.9672	93M11-94M11, 97M08-97M12 08M09-08M11

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

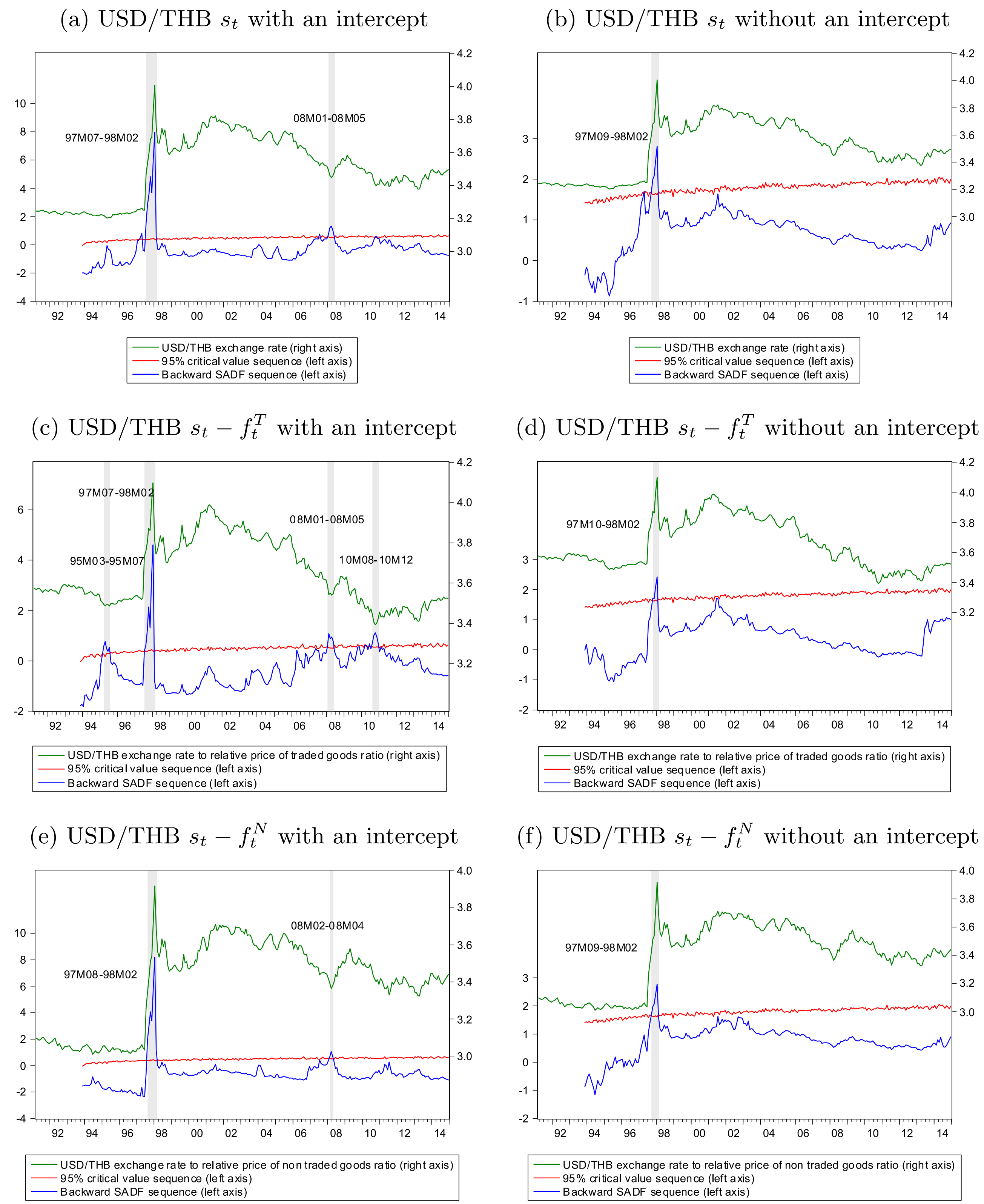

Fig. 4. Dating strategy for USD/THB nominal exchange rate s_t, the relative ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$ and the relative ratio of the exchange rate to the non-traded goods fundamental $s_t - f_t^N$. (a) USD/THB s_t with an intercept. (b) USD/THB s_t without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

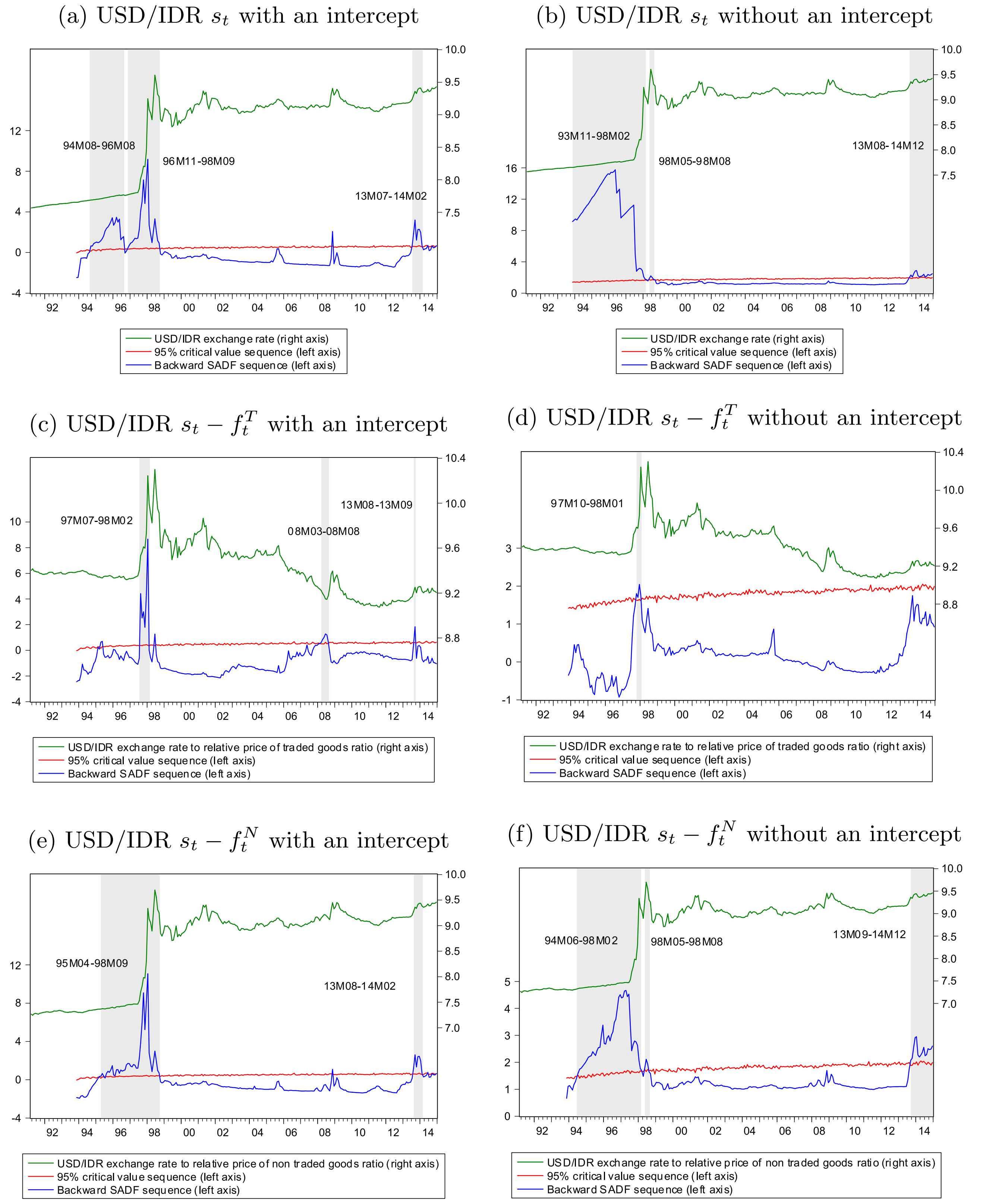

Fig. 5. Dating strategy for USD/IDR nominal exchange rate s_t, the relative ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$ and the relative ratio of the exchange rate to the non-traded goods fundamental $s_t - f_t^N$. (a) USD/IDR s_t with an intercept. (b) USD/IDR s_t without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

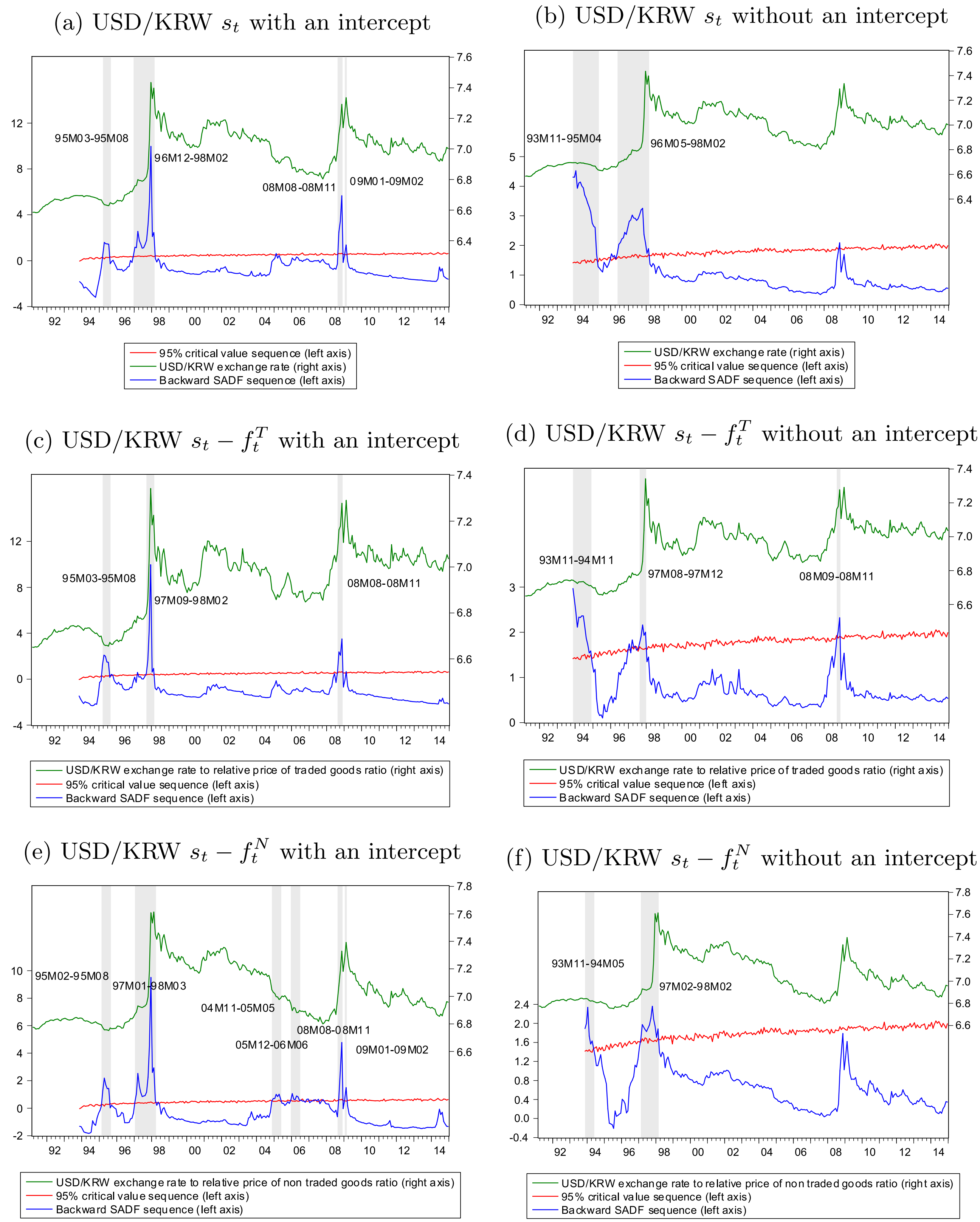

Fig. 6. Dating strategy for USD/KRW nominal exchange rate s_t, the relative ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$ and the relative ratio of the exchange rate to the non-traded goods fundamental $s_t - f_t^N$. (a) USD/KRW s_t with an intercept. (b) USD/KRW s_t without 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 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 07M10-08M07, 11M03-11M09 12M07-13M06	3.5298**	97M09-98M03, 98M05-98M10 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 07M09-08M08, 11M01-11M09	3.1190	97M11-98M02
$s_t - f_t^N$	3.7030***	94M07-95M11, 97M10-98M01 08M02-08M04, 08M11-09M01 10M08-11M09	2.5260	97M12-98M02
$s_t - f_t^T$	3.0141***	97M07-98M02, 98M05-98M09	3.2448[c*]	97M08-98M02, 98M05-98M09 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.

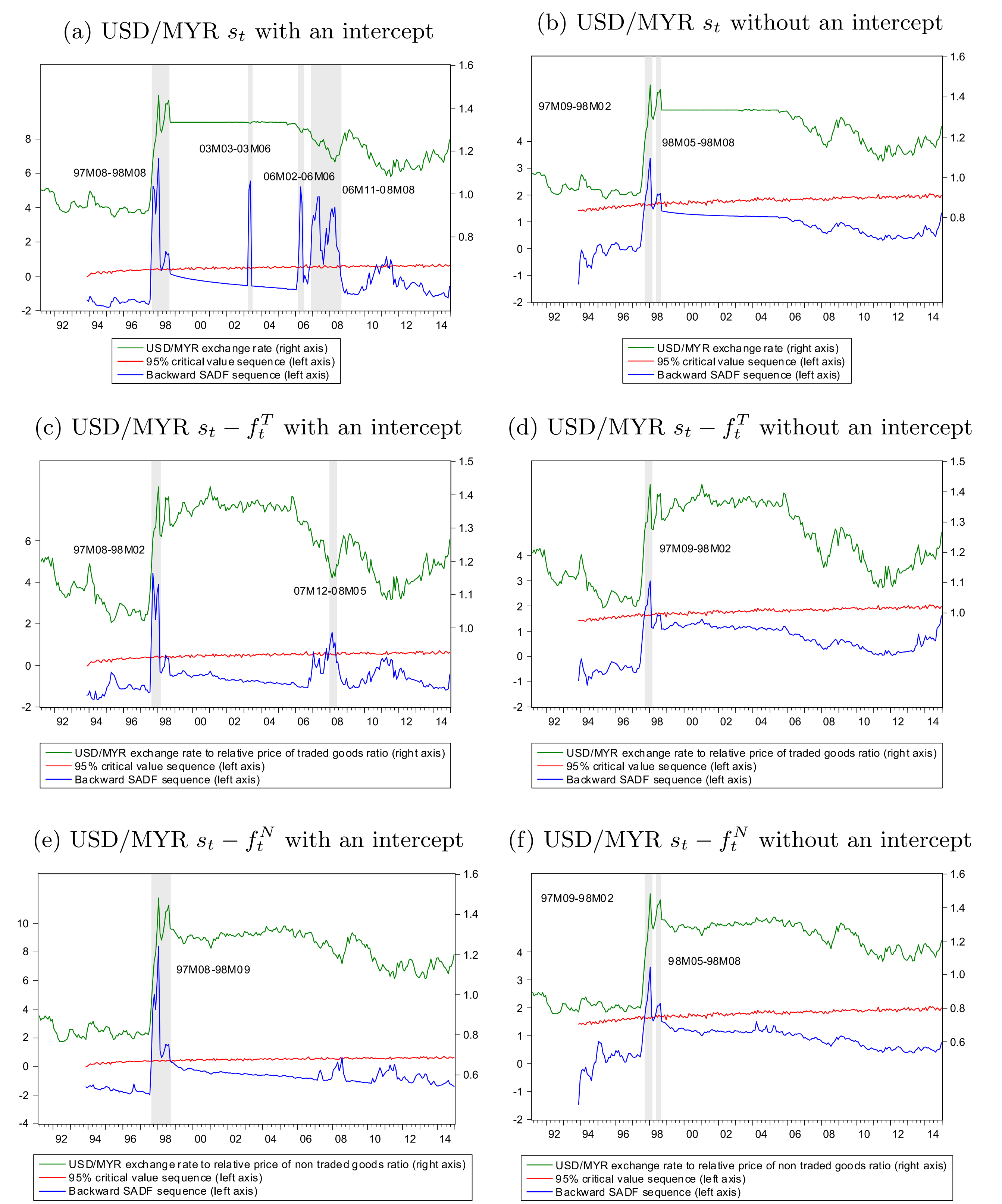

Fig. 7. Dating strategy for USD/MYR nominal exchange rate s_t, the relative ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$ and the relative ratio of the exchange rate to the non-traded goods fundamental $s_t - f_t^N$. (a) USD/MYR s_t with an intercept. (b) USD/MYR s_t without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

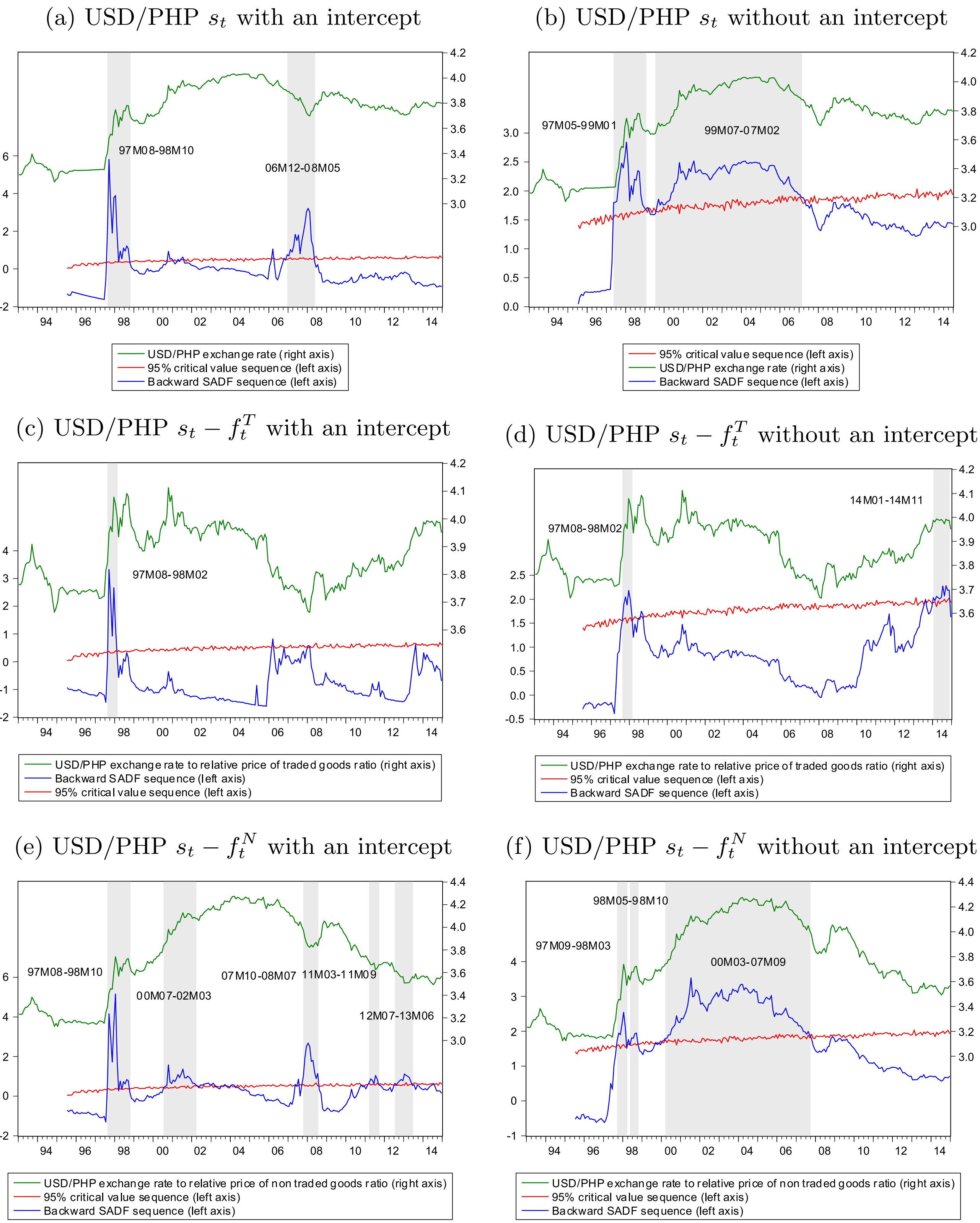

Fig. 8. Dating strategy for USD/PHP nominal exchange rate s_t, the relative ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$ and the relative ratio of the exchange rate to the non-traded goods fundamental $s_t - f_t^N$. (a) USD/PHP s_t with an intercept. (b) USD/PHP s_t without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

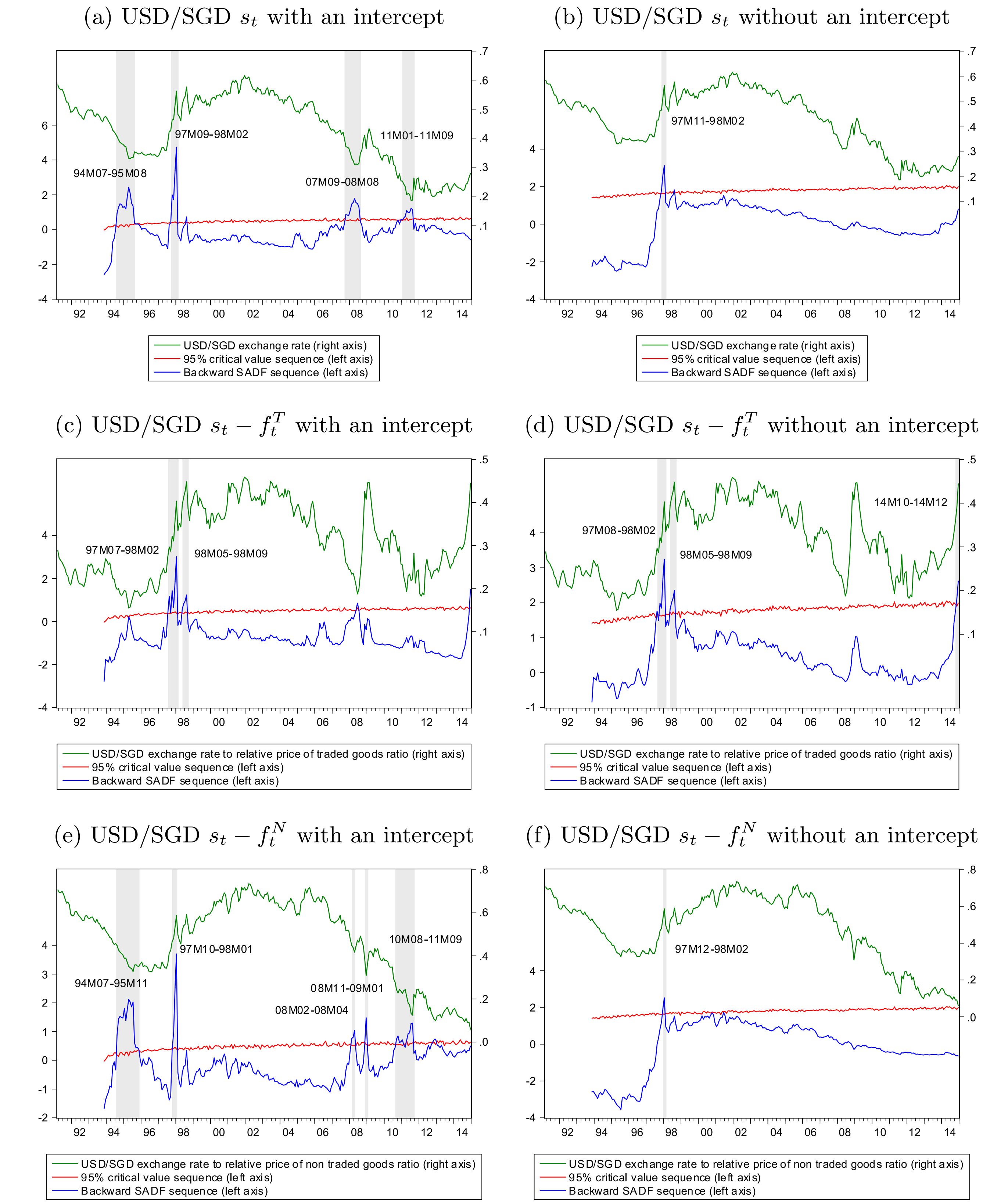

Fig. 9. Dating strategy for USD/SGD nominal exchange rate s_t, the relative ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$ and the relative ratio of the exchange rate to the non-traded goods fundamental $s_t - f_t^N$. (a) USD/SGD s_t with an intercept. (b) USD/SGD s_t without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

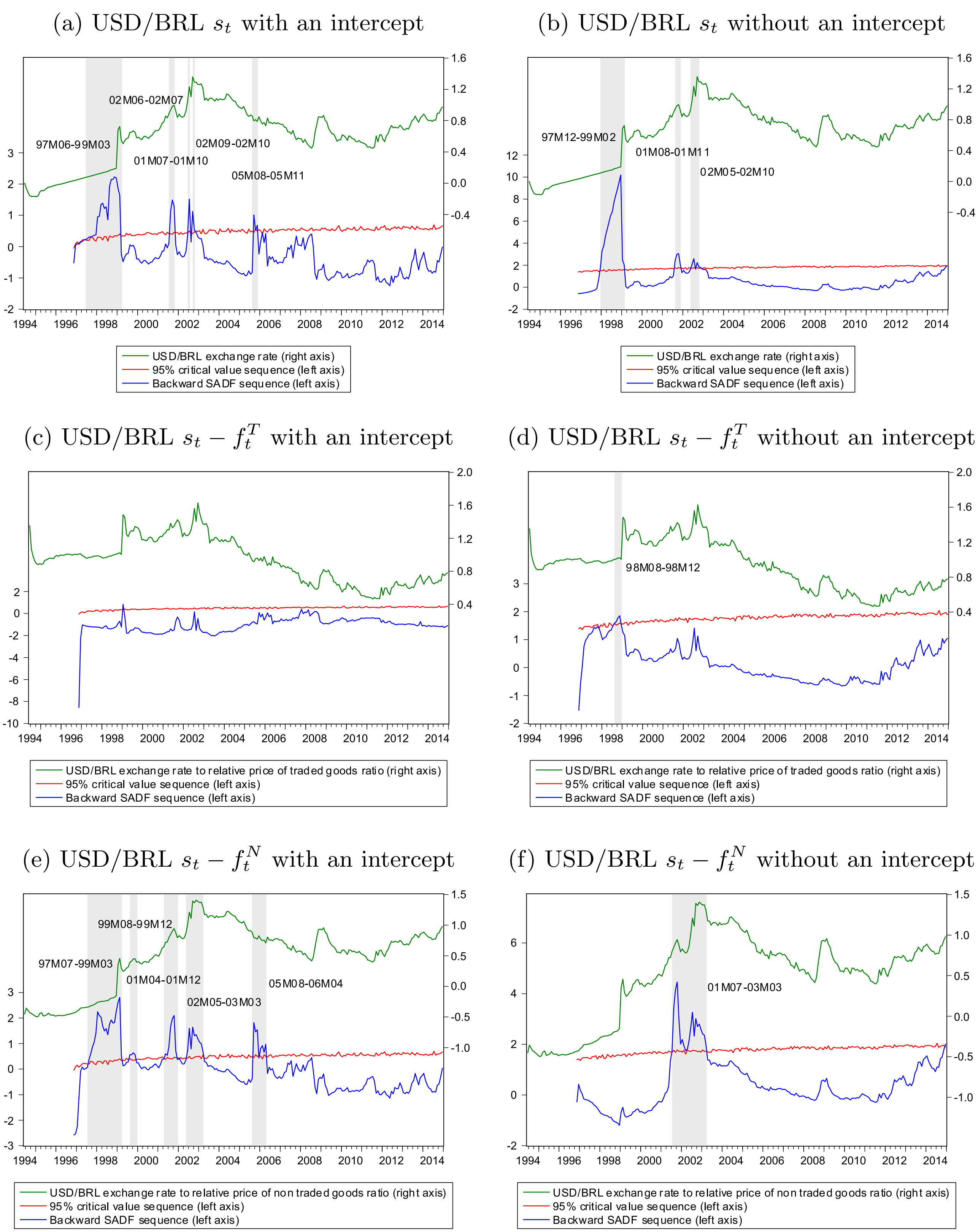

Fig. 10. Dating strategy for USD/BRL nominal exchange rate s_t, the relative ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$ and the relative ratio of the exchange rate to the non-traded goods fundamental $s_t - f_t^N$. (a) USD/BRL s_t with an intercept. (b) USD/BRL s_t without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

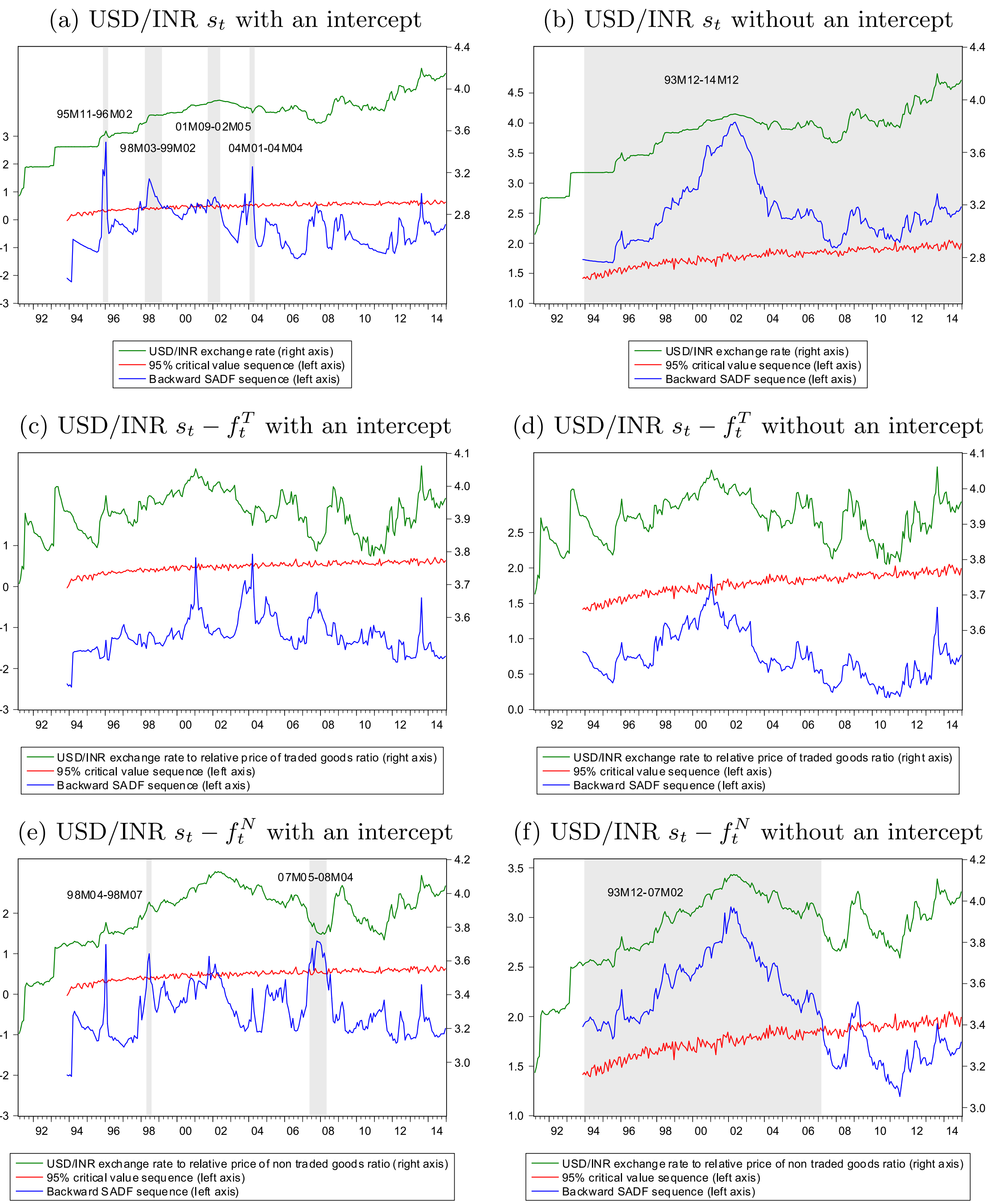

Fig. 11. Dating strategy for USD/INR nominal exchange rate s_t, the relative ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$ and the relative ratio of the exchange rate to the non-traded goods fundamental $s_t - f_t^N$. (a) USD/INR s_t with an intercept. (b) USD/INR s_t without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

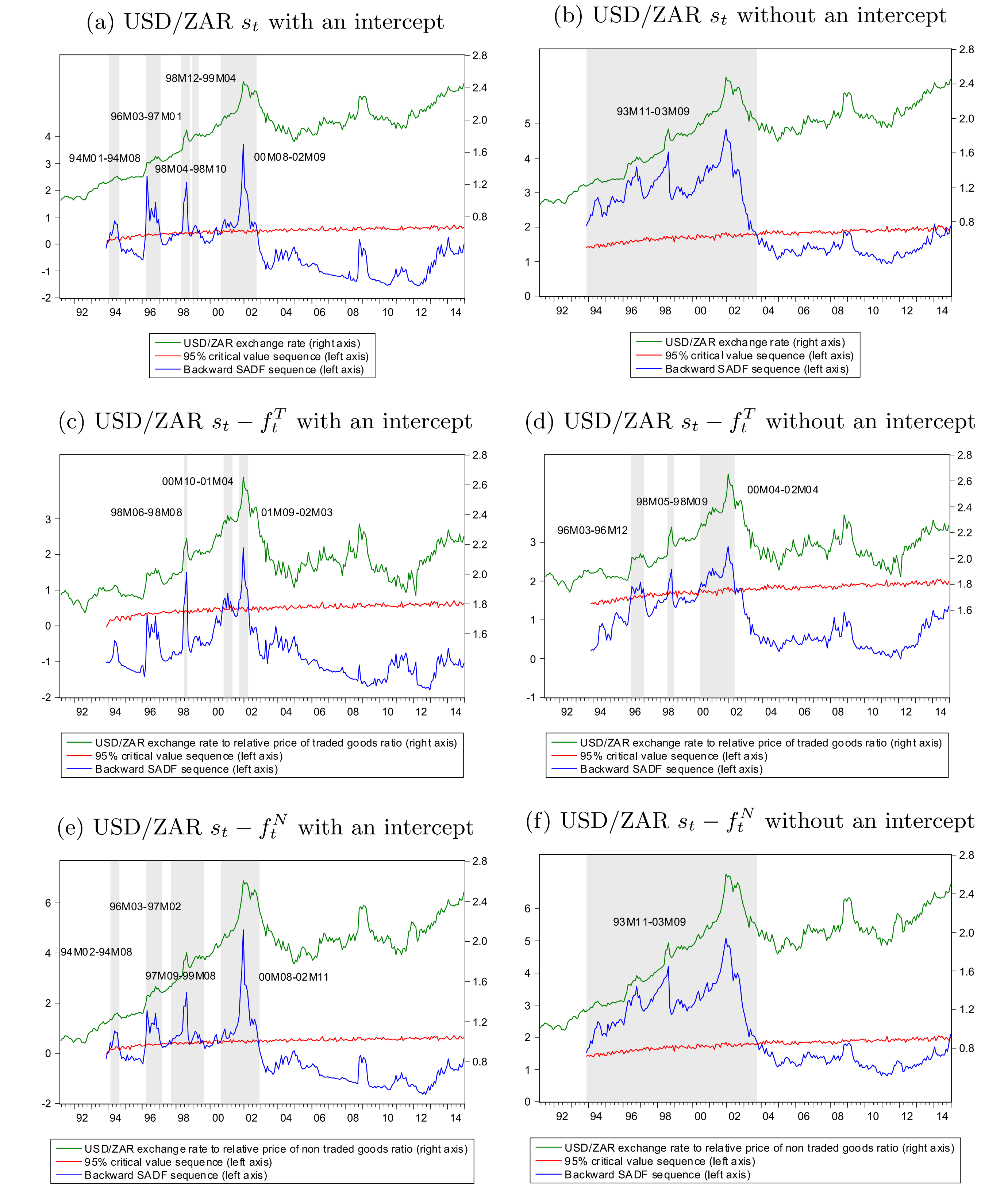

Fig. 12. Dating strategy for USD/ZAR nominal exchange rate s_t, the relative ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$ and the relative ratio of the exchange rate to the non-traded goods fundamental $s_t - f_t^N$. (a) USD/ZAR s_t with an intercept. (b) USD/ZAR s_t without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

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.

[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.

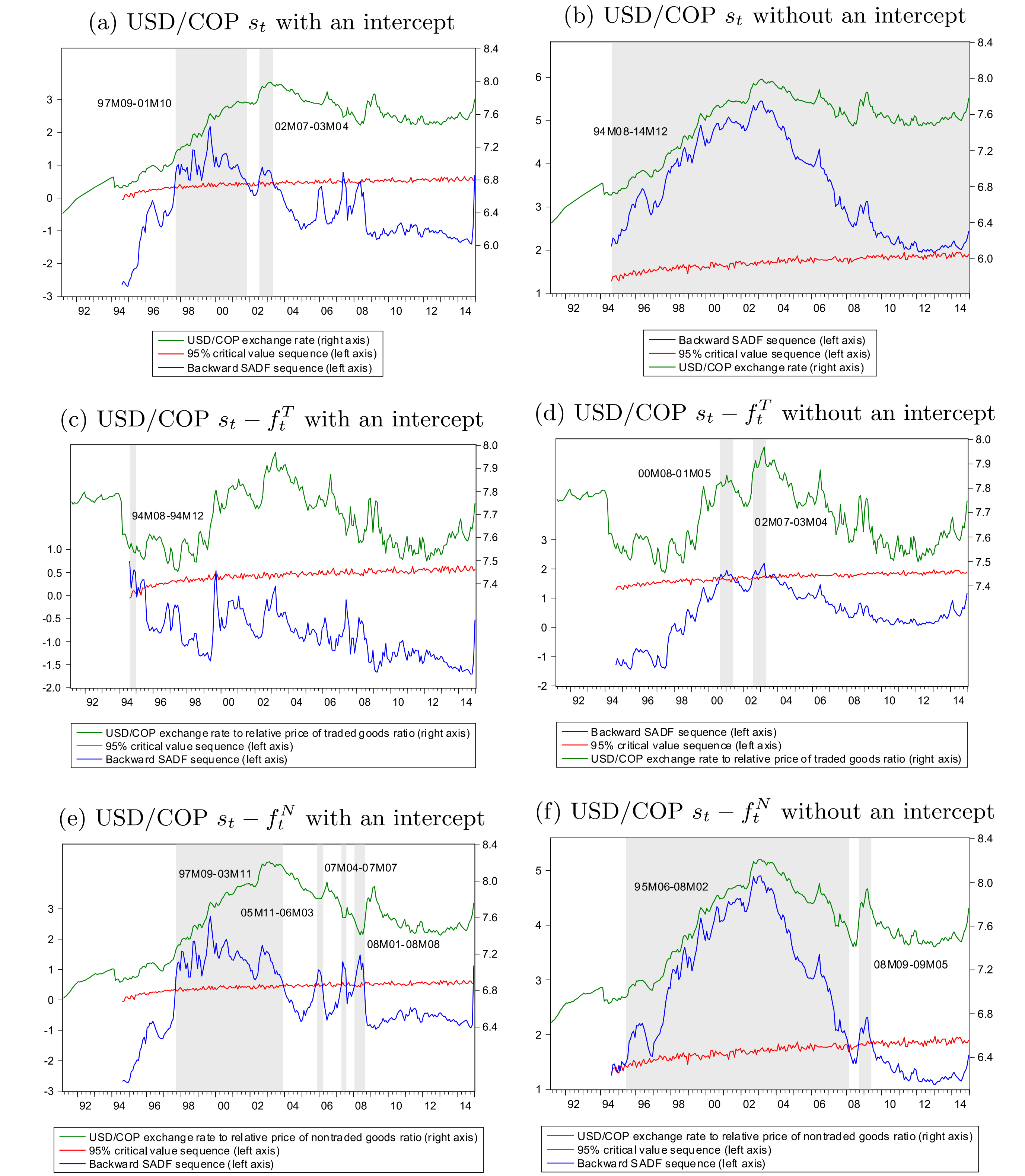

Fig. 13. Dating strategy for USD/COP nominal exchange rate s_t, the relative ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$ and the relative ratio of the exchange rate to the non-traded goods fundamental $s_t - f_t^N$. (a) USD/COP s_t with an intercept. (b) USD/COP s_t without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

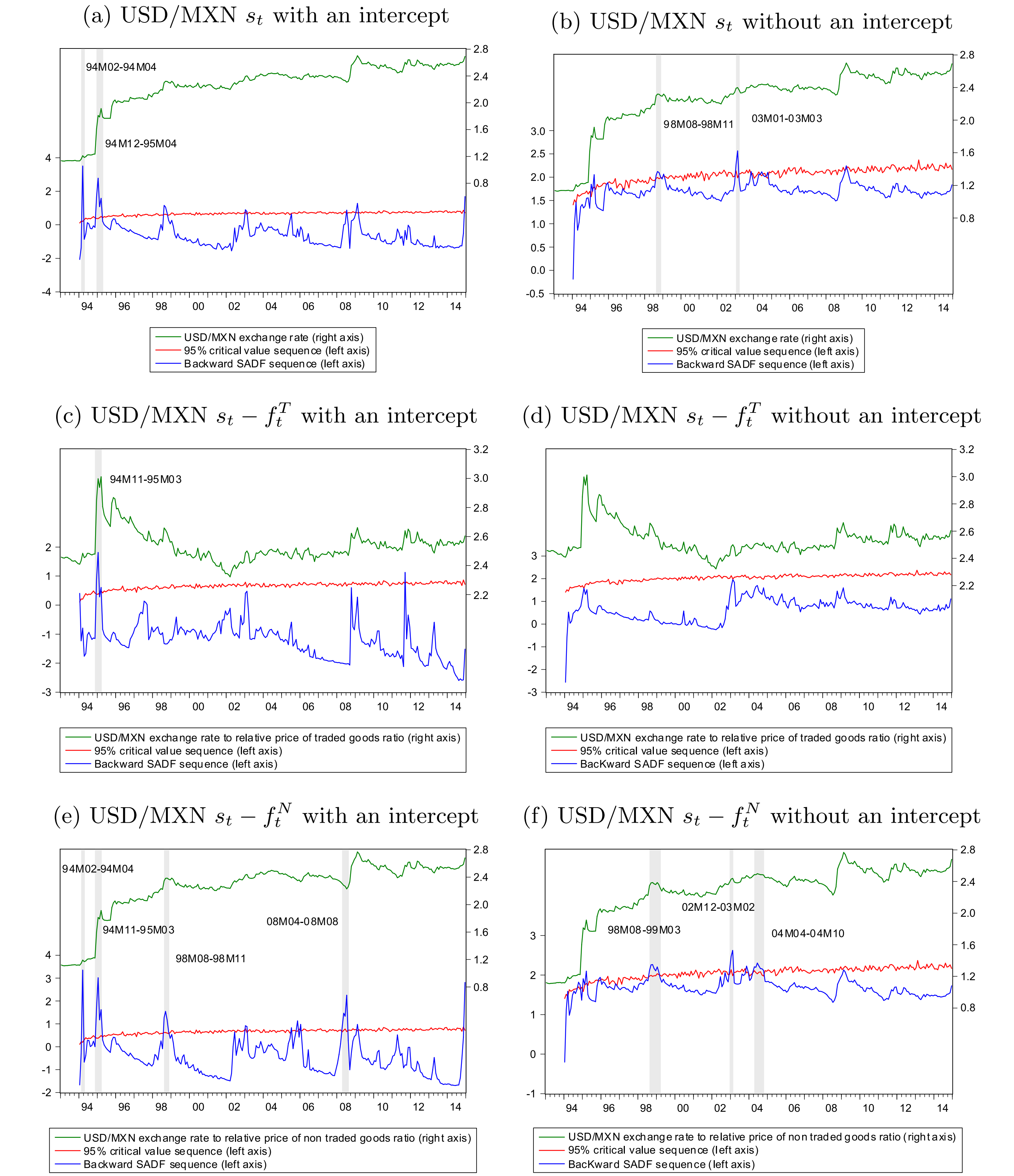

Fig. 14. Dating strategy for USD/MXN nominal exchange rate s_t, the relative ratio of the exchange rate to the traded goods fundamental $s_t - f_t^T$ and the relative ratio of the exchange rate to the non-traded goods fundamental $s_t - f_t^N$. (a) USD/MXN s_t with an intercept. (b) USD/MXN s_t without 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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Chapter 4

Exuberance, Bubbles or Froth? Some Results using Very Long Run Historical House Price Data for Amsterdam, Norway and Paris.

Abstract

The idea that house prices can experience periods of 'bubbles' has recently gained support with some suggesting they were either a cause or effect of the Global Financial Crisis. In this research, we econometrically test whether historical house prices exhibit bubbles and, if they do, whether contagion from other historical financial crises are mirrored in these housing markets. We apply the generalized sup ADF (GSADF) test of Phillips, Shi, and Yu (2015a, PSY) and long memory estimators of Geweke and Porter-Hudak (1983), Robinson (1995) and Shimotsu (2010) to search for evidence of exuberance or bubbles in several historical housing price indices, in particular, the Herengracht index for Amsterdam (1649-2010), Norway (1819-2014) and Paris (1650-2012). We find, firstly, there is little evidence of exuberance in the real Herengracht index or of bubbles in the house price-rent ratio for Amsterdam. Secondly, our results provide evidence of exuberance in Norwegian house prices, where exuberant episodes coincide with the major financial crises in Norwegian history. Thirdly, evidence of exuberance is found in the historical house price series of Paris only for a particular model specification. House price exuberance appears to be a very recent phenomenon as no explosive behaviour is detected in these house price series before 1850. Fourthly, our results from PSY are generally in line with those obtained from long memory models based on the mean value of d but not the revelent 95% confidence intervals.

Keyword

Bubbles; Generalized sup ADF test; Exuberance; House prices; Herengracht Index; Norway; Paris

JEL classifications: *G01; N2; N90; R30; R39*

4.1 Introduction

The Global Financial Crisis (GFC) and its aftermath, the Great Recession, has popularized the idea that house prices can exhibit "bubbles" and although some economists continue to deny their existence, several central banks have decided to try and cool their hot domestic housing markets[1]. Perhaps the most famous deniers, Case and Shiller (2003), consider that the early 2000s were not in fact the beginning of a house price bubble in the US, but that income growth driven "fundamentals" could explain why houses were more affordable than they had been in 1995. Similarly, McCarthy and Peach (2004) and Himmelberg et al. (2005) conclude that, based on data from 1980-2004, real estate prices could be

[1] In 2014 the Reserve Bank of New Zealand introduced minimum deposit requirements to try and deflate what they saw as a bubbly Auckland housing market. The paper by Greenaway-McGrevy and Phillips (2016) supports this conclusion and that a similar, but more pervasive bubble occurred in New Zealand house prices in 2003.

explained by fundamentals alone. Recent work by Hu and Oxley (2018a) leads to conclusions that dispute these traditional views.

Identifying bubbles, or periods of aberrant, exuberant, price changes empirically, however, is a difficult and often controversial exercise, one made harder by the relatively short sample periods used by some recent authors see for example, Del Negro and Otrok (2007), Goodman and Thibodeau (2008), Wheaton and Nechayev (2008) and Mayer (2011). It seems particularly hard to identify such episodes when bubbles grow, but deflate slowly rather than 'bursting with a bang', see Ambrose et al. (2013).

It is well-documented that the presence of a unit root in log price-dividend ratio implies rational bubbles, see Diba and Grossman (1988), Froot and Obstfeld (1991) and Craine (1993). When standard unit root tests are employed in such studies, they report mixed results, perhaps because such tests find it hard to distinguish between a unit root and a fractional-order root (a near unit root). Several studies, therefore, have applied long memory models to test for the presence of bubbles, for example, see Cuñado et al. (2005), Cuñado et al. (2012), Koustas and Serletis (2005), Frömmel and Kruse (2012). Long memory models have also been applied to housing markets, see Barros et al. (2012), although such applications are rare. The presence of a long memory process is inconsistent with rational bubbles. Following a recent study of Ramanan (2016) on US housing markets using long memory models, we will consider a range of estimators for long memory focusing, in particular, on the semi-parametric two-step exact local Whittle procedures of Shimotsu (2010) to estimate the value of the long memory parameter d for historical house price series. This particular estimator is consistent for $d \in (\frac{1}{2}, 1)$ and robust to non-normality and heteroskedastic errors.

In terms of recent econometric testing, however, Phillips and co-authors (Phillips and Yu (2011), Phillips, Wu, and Yu (2011)[2], Phillips, Shi, and Yu (2015a, PSY), Phillips, Shi, and Yu (2015b)) have developed right-tailed only unit root tests to

[2]Harvey, Leybourne, Sollis, and Taylor (2016) investigate the non-stationary volatility effect

"date-stamp" bubbles, even (different) bubble episodes that might occur multiple times in a long time series. Their empirical examples include analysis of 150 years of data for the S&P 500, the NASDAQ during the 1990s, and recently an investigation of the housing market in Auckland, New Zealand see Greenaway-McGrevy and Phillips (2016). Hu and Oxley (2018b) also investigate the famous South Sea and Mississippi Bubbles during the period 1719-1720.

A crucial starting point for the discussion and identification of bubbles relates to how they are defined. We follow the standard approach that defines price movements in relation to some 'fundamental'. In terms of house prices this might be a price-to-rental cost ratio; a price-to-average income ratio etc, for example, the PSY approach is often applied to a price-fundamental ratio to assess the explosive behaviour. Because this is a right-tail only test, if we reject the null hypothesis of a unit root, we conclude in favor of explosive behavior for x_t. If the time series x_t involves an economic fundamental[3], we may conclude that a finding of explosive behavior denotes the presence of *a bubble*. Alternatively, if the time series x_t doesn't involve an economic fundamental, but simply a proxy like price-to-income ratio, we may only conclude that a finding of explosive behavior is evidence of *an exuberant episode*. Without a fundamental series to compare the time path of (e.g, house) prices to, econometric tests, including those of PSY, can only identify periods of exuberant growth (or decline) in the series, which may be a necessary property of a bubble, but is not sufficient (without a model for the fundamental). If the measured series is neither a bubble nor a period of exuberance (or alternatively collapse in prices), then we refer to changes over time in the market as 'frothy'.

In this paper, apart from long memory models, we will utilize the econometric approaches of PSY and variants thereof, to consider the extent to which house

on the reliability of the Phillips, Wu, and Yu (2011) test and propose wild bootstrap implementation of this test to overcome the spurious indications of explosive bubbles.

[3]The time series x_t is commonly expressed as a ratio (e.g., house prices/ rents or income, stock prices/ dividends).

price bubbles or periods of exuberant behavior are exclusively recent behavior or whether history provides similar time series episodes. In particular, we will consider the time series properties (long memory) and results of PSY-type tests applied to three long-term series; 355 years of the Herengracht, Amsterdam house price series (see, Eichholtz (1997), Ambrose et al. (2013), and Eichholtz et al. (2015)); house prices in four of the five main Norwegian cities (Oslo, Bergen, Trondheim and Kristiansand) between 1819-1989 (see Eitrheim and Erlandsen (2004)) and Friggit (2001)'s house price series for Paris, 1650-2012[4].

We investigate the explosiveness of these historical house price indices for several reasons. First, long-term house price indices are extremely scarce. Most house price indices in the literature exist for a few decades or even less. However, there are few exceptions for example, Eichholtz (1997) has constructed a house price index of an area of Amsterdam since 1628, which is widely known as the Herengracht index. The Herengracht index reports house prices over a 350 year period in one of the most prestigious locations in Amsterdam in the Netherlands. Friggit (2001) has constructed an annual house price index for Paris (1200-2012), which is regarded as the longest historical house price series in the literature. Eitrheim and Erlandsen (2004) have also developed annual house price indices for four main Norwegian cities since 1819. In addition, an aggregate Norwegian index has also been created.

Second, these historical series represent different types of real estate markets in terms of their coverage. For example, the Herengracht index covers a small area in Amsterdam while the Paris and the city-level Norwegian indices represent the whole city. The aggregate Norwegian index simply measures the house price inflation for the whole country. It would seem to be of interest to investigate

[4] Recently, Blöndal (2015) investigated the possibility of speculative bubbles in the Stockholm housing market 1875-1935, using a newly created series and tests based upon the existence of common trends in a cointegrating regression framework. He concluded that there was no indication of a speculative bubble in the Stockholm housing market. His sample period is relatively short for our purposes.

explosive behavior in different types of real estate markets and this is what we do here.

Third, being very long data, these series appear to be ideal candidates to apply the PSY approach in the search for house price bubbles and potential contagion as it allows for the possibility of multiple bubble episodes. For example, Amsterdam experienced several crisis episodes in its history including the Mississippi Bubble (1719-1720), the South Sea Bubble (1720), the Amsterdam Banking Crisis of 1763 and the Credit Crisis of 1772. Similarly, several severe banking crises have occurred in Norway especially during the 19th century (1899-1905, 1920-1928 and 1988-1993), see Gerdrup (2004) and Vale (2004). Grytten and Hunnes (2010) also discuss nine major financial crises in Norway's history since independence in 1814. There are also several major historical bubbles/crises in French history including the Mississippi Bubble (1719-1720), the Crisis of 1763, the Crisis of 1772 and the French stock market crisis of 1882. Fourth, there exist a clearly recorded historiography and sets of complementary data, which allow possible bubble/exuberant periods, identified by the time series approaches of PSY, to be scrutinized against contemporary events. We utilize this extant historiography to consider the extent to which events identified by the tests could likely have been driven by events of the time.

There are several contributions from this research. First, the majority of existing studies investigate the presence of housing bubbles using a short series with no more than a few decades. Those short house price series tend to be associated with a fewer (possible/likely) bubble episodes and are typically pre-screened in that they are chosen to test whether a particular episode was bubbly. To this extent any published evidence of recent bubbles could be pre-test biased. Our paper attempts to overcome some of these problems by considering very long periods of data where one might expect to identify a range of possible episodes, calm, froth, exuberance, and bubbles. Our approach does not attempt to 'cherry pick' a particular period with an expected high probability of a bubble, but to trace the time series behaviour over a very long period of time. Second, prior to the results presented here, no housing bubbles or house price exuberance has been reported or investigated using these powerful PSY-type bubble detection tests using historical data prior to 1900. Unlike the famous stock price bubbles of the Mississippi, the South Sea and the British Railway Mania in the eighteenth and nineteenth centuries, house price bubbles seem to be a very recent phenomenon. Among many prominent bubble episodes, there are only two well-documented housing bubbles in thoroughly researched in recent history: Japan's real estate bubble in the 1980-90s and the US housing bubble during the 2000s. Our paper therefore contributes to the literature by investigating house price exuberance/bubbles prior to 1900. Third, although some authors (e.g., Friggit (2008), Ambrose et al. (2013) and Eichholtz et al. (2015)) have undertaken some statistical analysis of these historical house price series, none have formally tested for the existence of bubbles or price exuberance. Moreover, their statistical tests cannot locate the origination and termination of bubbles. However, an econometric test that formally tests for evidence of bubble or price exuberance (and allows for multiple episodes of such) is now available see Phillips, Wu, and Yu (2011), Phillips, Shi, and Yu (2014) and Phillips, Shi, and Yu (2015a, PSY). The more recent right-tailed unit root test of Phillips, Shi, and Yu (2015a, PSY) in particular, is ideally suited to the search for multiple bubbles/exuberant episodes in the house price series of Amsterdam, Norway and Paris. In this study, we present empirical results based on different model specifications and lag order selection.

The remainder of the paper is organized as follows. Section 4.2 provides an overview of the data used[5]. Section 4.3 provides a brief description of the long memory estimators of Geweke and Porter-Hudak (1983), Robinson (1995) and Shimotsu (2010) and the GSADF test of Phillips, Shi, and Yu (2015a). Section 4.4 provides results for the long run house price series and Section 4.5 concludes.

[5] An excellent description of the data including sampling, coverage and sources can be found in Friggit (2008) and the references therein.

4.2 Data

This section describes the three historical housing price indices used in the paper and the time series plots of these indices are shown in Figure 4.1. We also construct the house price-rent ratio for Amsterdam, which is presented as Figure 4.2.

4.2.1 The Herengracht Index (1649-2010) and the Biennial Herengracht Index (1628-1973)

The Herengracht index was sourced from Piet Eichholtz's website for the period 1649-2010. We also obtain the biennial Herengracht index between 1628 and 1973 from Eichholtz (1997). These data represent an update on the biennial series (1628-1973) first presented and discussed extensively in Eichholtz (1997), updated to 2005 in Ambrose et al. (2013) and Eichholtz et al. (2015). Figure 4.1a and Figure 4.1b (below) provide the time series plots of the real Herengracht index and biennial Herengracht index (both on a log scale)[6].

The series cover, on a biennial basis, 487[7] properties located along the banks of the Herengracht, a canal in Amsterdam. In the Golden Age of this city, the 17th century, this area was the most fashionable place in the Netherlands. It was urbanized very early: by 1680 nearly all the lots along the canal had been developed. The index was created using a repeat sales method by comparing the successive sale price of buildings. 4252 transaction prices for the 1628-1973 period were collected, which means on average around 12.3 prices per year.

The only quality effect taken into account is the use of the buildings. Beginning in the 19th century, but especially in the 1920s and 1930s, many buildings along the Herengracht were changed into offices, which increased their value. Buildings used as offices have been excluded from the calculation. All other quality effects have

[6]The real Herengracht index is deflated by the CPI from van Zanden (2005).

[7]487 presently, as opposed to 614 originally, this decrease stemming from the combination of lots to allow for the construction of bigger buildings.

not been filtered out, thus if for example central heating was installed between two successive transactions, the resulting effect of this amenity on the price is ignored.

4.2.1.1 House Price-rent Ratio for the Herengracht (1628-1850)

Ambrose et al. (2013) created an Amsterdam house price-rent ratio for the period 1650-2005, however, this ratio is not publicly available. Thus, we construct the house price-rent ratio (1628-1850) to assess the evidence of housing bubbles using data obtained from Eichholtz (1997) and Eichholtz et al. (2012). The nominal biennial Herengracht index 1628 and 1850 was obtained from Eichholtz (1997) and is shown in Figure 4.2a. We utilise the temporal disaggregation method of Dagum and Cholette (2006) to disaggregate the nominal biennial Herengracht index series (1628-1850) to the annual Herengracht index series (1628-1850) and is displayed as Figure 4.2b. The rent index obtained from Eichholtz et al. (2012) is presented as Figure 4.2c and we select the sub-period 1628 and 1850. A house price-rent ratio was calculated and presented as Figure 4.2d.

4.2.2 House prices in Norway (1819-2014)

The house price indices for Norway are from Eitrheim and Erlandsen (2004) and can be downloaded from Norges Bank[8]. Figure 4.1c provides the real house price index[9] in Norway (1912=100, both in log scale) between 1819 and 2014 with 195 observations. The construction of this annual house price index, 1819-2014, is described in some detail in (Eitrheim and Erlandsen, 2004, p.357) and is calculated by a hedonic-weighted repeat sales (hybrid) approach based upon transaction prices in the property registers to 1985 and per square meter from 1986 (based upon data from the Norwegian Association of Real Estate Agents). In particular, the indices are constructed on the basis of nominal transaction prices of property,

[8]The house price index can be accessed from http://www.norges-bank.no/en/Statistics/Historical-monetary-statistics/.

[9]The real house price index is deflated by the CPI from Grytten (2004).

compiled from the archives of real property registers of the four cities (Eitrheim and Erlandsen, 2005, p.8). Only buildings located in the centre of the respective town are used and the types of buildings vary from rental apartment blocks to single family homes.

In addition to the national level series, data are available separately for four of the five main Norwegian cities, Oslo (1841-2014), Bergen (1819-2014), Trondheim (1897-2014) and Kristiansand (1867-2014). Due to data availability, the samples for these four cities cover different years in the 1800s. A weighted repeat sales method is used to construct the city house price indices, which are described in (Eitrheim and Erlandsen, 2005, p.15).

4.2.3 House prices in Paris (1200-2012)

The house price index for Paris was constructed by Friggit (2008) using several different sources including, d'Avenel (1894), Duon (1946), Friggit (2001) and covers the period 1200 to 2012 (2000=100). The time series plot of the real house price index is given as Figure 4.1d. The real series are deflated by a consumer price index from several sources. Friggit (2008) presents an extensive discussion of four related price indices for Paris and from this constructs his own Friggit (2001):

(i) d'Avenel (1894) who provides a series of average home prices (averaged over 25 years) for Paris, 1200-1800. No adjustments are made for house quality changes;

(ii) Duon (1943a,b) creates two variants of a repeat sales home price index for Paris. For the period 1790-1850 the calculations are based upon 10 year periods 1790-1850 and for 1840-1944 a yearly home price index; Notaries databases. This index applies to apartments sold by the unit and is based upon a record of transactions (both current and previous transactions for the unit are recorded). Quality changes are not recorded or recognised. Some

34,594 transaction pairs were used to construct the index for the period 1944-1999;

(iii) Notaires-INSEE index. For the years after 1999 an annualised value for quarterly Notaries-INSEE indices can be constructed based upon hedonic indices (see here Gouriéroux and Laferrère (2009));

(iv) Based upon the properties of these indices, Friggit (2001, 2008) constructs a Paris house price 1840-2006 (updated to 2012) which has been adjusted for obsolescence prior to 1914 (see Friggit (2008) for details).

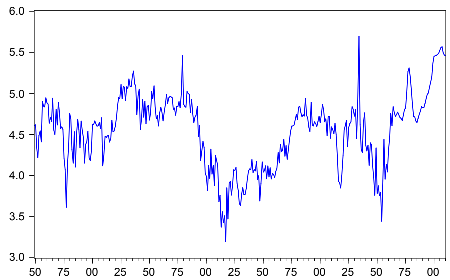
4.3 Method

Testing for the existence of bubbles has followed three basic routes: i) cointegration-based tests for example, Campbell and Shiller (1987); ii) the long-memory/fractional integration approach; iii) the right tailed only, date-stamping approach of PSY. In this paper, we will concentrate on ii) and iii) as the cointegration based tests are indirect tests which do not reject the no-bubble hypothesis if the price is cointegrated with the fundamental value. For example if the house price to rent ratio is stationary or house price is cointegrated with the fundamental price then the no bubble hypothesis cannot be rejected. The approach is not well suited for tests of multiple bubbles (something we are keen to test for) and is now rarely used in the literature as it has been superseded by ii) and iii).

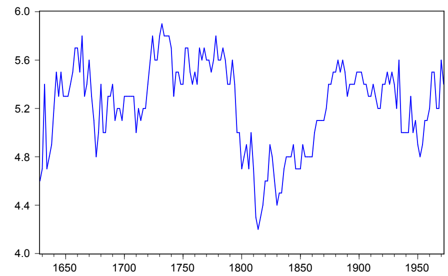
4.3.1 Long Memory Estimation

A number of studies for example, Gil-Alana and Hualde (2009) and Barros et al. (2012) utilize the notion of persistence as exhibited by long-memory processes to define testable hypotheses about the existence or otherwise of bubbles using estimates (d) of a fractional integration (long memory) process; $0 < d < 1$.

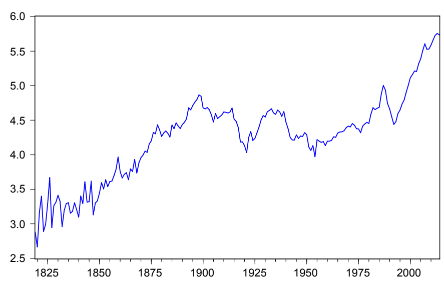
(a) Herengracht index in Amsterdam (1649-2010)



(b) Biennial Herengracht index in Amsterdam (1628-1973)



(c) House price index in Norway (1819-2014)



(d) House price index in Paris (1650-2012)

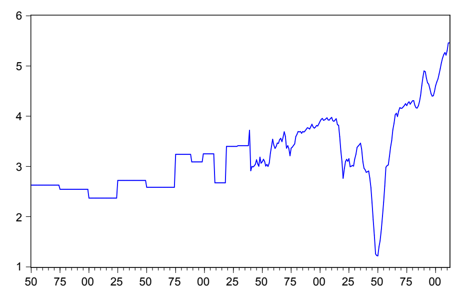


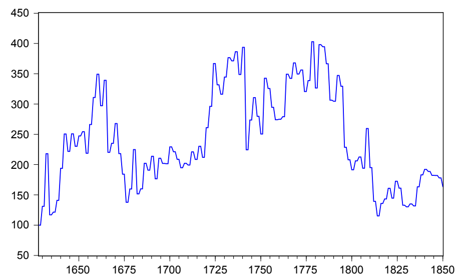
Figure 4.1 – (a) The real Herengracht index (1649=100, both in log scale) between 1649 and 2010; (b) the real biennial Herengracht index (1628=100, both in log scale) between 1628 and 1973; (c) the real house price index in Norway (1912=100, both in log scale) between 1819 and 2014; (d) the real house price index in Paris between 1650 and 2012 (2000=100, in log scale).

In particular, for the following values (ranges) for d, this approach implies the following[10]:

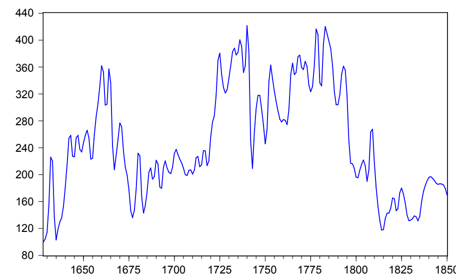
- $-\frac{1}{2} < d < 0$ stationary or overdifferenced, mean reverting, no rational bubble;
- $d = 0$ stationary, no rational bubble;

[10]See Ramanan (2016) for an extended discussion of the rational bubble.

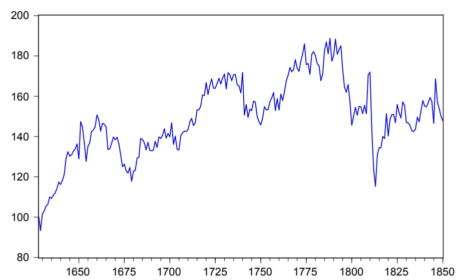
(a) Nominal biennial Herengracht index (1628-1850)



(b) Nominal Herengracht index (1628-1850)



(c) Nominal rent index (1628-1850)



(d) House price-rent ratio (1628-1850)

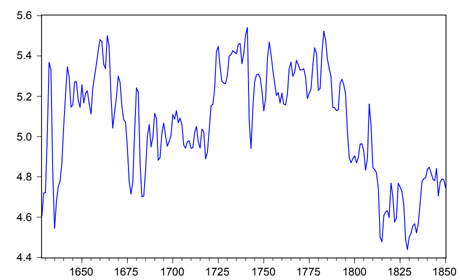


Figure 4.2 – (a) The nominal biennial Herengracht index (1628-1973, 1628=100) is obtained from Eichholtz (1997); (b) The nominal annual Herengracht index (1628-1973, 1628=100) is obtained using the temporal disaggregation method of Dagum and Cholette (2006); (c) The nominal rent index (1628-1850, 1628=100) is obtained from Eichholtz et al. (2012); (d) A house price-rent ratio for this period is therefore calculated (1628=100, in log scale).

- $0 < d < \frac{1}{2}$ possesses long memory and mean reversion, no rational bubble;
- $\frac{1}{2} < d < 1$ non-stationary ARFIMA process. Long memory but mean reverting, no rational bubble;
- $d \geq 1$ non-stationary and explosive process, no mean reversion, rational bubble.

As was demonstrated in Rea et al. (2013), ‘not all estimators (of d) are born equal’ and it is important to be aware of the various properties of alternative estimation methods for d given how crucial it is in this approach to determine whether there is prima facie evidence of bubbles or not. The sample sizes available in the data used here cover long historical periods, but the number of observations is relatively small compared to the typical high frequency financial data used to estimate long memory processes. As a result, we consider three methods to estimate the value of the long memory parameter, d, i) GPH estimator of Geweke and Porter-Hudak (1983); ii)the local Whittle estimator of Kuensch (1987) and Robinson (1995); iii) and the two-step exact local Whittle estimator of Shimotsu (2010). As will be seen in the reported results one of the major issues with very popular estimators like GPH is its lack of efficiency in small samples. Unlike other studies estimating d we will report not only the mean estimate but also its standard error (se).

4.3.1.1 GPH Estimator

Geweke and Porter-Hudak (1983) develop the so-called GPH estimator to estimate d of the ARFIMA model using log-periodogram regression. The periodogram of time series data x_t at frequency ω with the sample size T is described as:

$$I_x(\omega_j) = \frac{1}{2\pi}\Big|w(\omega_j)\Big|^2 = \frac{1}{2\pi T}\left|\sum_{i=1}^{T} x_t e^{-it\omega_j}\right|^2, \tag{4.1}$$

$$\omega_j = \frac{2\pi}{T}\left(-\frac{T}{2}+j\right), \qquad \text{for} \qquad j = 1, \ldots, T. \tag{4.2}$$

The spectrum density function of ARFIMA (0,d,0) is given as:

$$f_x(\omega) = \frac{\sigma^2}{(4\sin^2\pi\omega)^d}. \tag{4.3}$$

Regressing the log of the periodogram $I_x(\omega_j)$ on the log of the spectrum density $f_x(\omega)$ gives,

$$logI_x(\omega_k) = c - dlog(4\sin^2\pi\omega) + \varepsilon_k, \tag{4.4}$$

where c is a constant, $\omega_k = k/T$ and $k = 1, \ldots, n < T/2$. The OLS estimate of d_{GPH} is then calculated.

4.3.1.2 Local Whittle Estimator

The local Whittle estimator is proposed by Kuensch (1987) and Robinson (1995). The following notation is taken from Robinson (1995) and Ramanan (2016). The discrete Fourier transform and periodogram of x_t are defined as:

$$\omega(\lambda) = (2\pi n)^{-1/2} \sum_{t=1}^{n} x_t e^{it\lambda}. \tag{4.5}$$

$$I_x(\lambda) = \left| \omega(\lambda) \right|^2, \tag{4.6}$$

The objective function is defined as:

$$Q_m(G, d) = \frac{1}{m} \sum_{j+1}^{m} \left[\log(G\lambda_j^{-2d}) + \frac{\lambda_j^{2d}}{G} I_x(\lambda_j) \right], \tag{4.7}$$

where $I_x(\lambda_j)$ is the periodogram of X_t evaluated at the fundamental frequencies and X_t is a fractional process with order d.

The local Whittle estimates G and d by minimising $Q_m(G, d)$,

$$\left(\hat{G}, \hat{d}\right) = \underset{G \in (0,\infty), d \in [\Delta_1, \Delta_2]}{\operatorname{argmin}} Q_m(G, d) \tag{4.8}$$

where Δ_1 and Δ_2 are numbers such that $0 < \Delta_1 < \Delta_2 < \infty$. The local Whittle estimator of $\hat{d}$ is denoted as $\hat{d}_{LWE}$.

4.3.1.3 Two-step Exact Local Whittle Estimator

Shimotsu (2010) proposed a 2-step ELW estimator that allows for an unknown mean and a trend. The data X_t can be generated by

$$X_t = \mu_0 + X_t^0; \quad X_t^0 = (1 - L)^{-d_0} u_t \mathbf{I} \{t \geq 1\} \tag{4.9}$$

where μ_0 is a non-random unknown finite number.

Shimotsu (2010) estimates the unknown mean μ_0 as a linear combination of the sample mean $\bar{X}$ and the first observation X_1:

$$\widetilde{u}(d) = \omega(d)\bar{X} + (1 - \omega(d))X_1, \tag{4.10}$$

where $\omega(d)$ is a twice continuously differentiable weight function such that $\omega(d) = 1$ for $d \neq 1/2$ and $\omega(d) = 0$ for $d \geq 3/4$.

The modified ELW objective function is given by

$$R_F(d) = \log \hat{G}_F(d) - 2d\frac{1}{m}\sum_{j=1}^{m} \log \lambda_j; \quad \hat{G}(d) = \frac{1}{m}\sum_{j=1}^{m} \mathbf{I}_{\Delta^d(x-\hat{\mu})}(\lambda_j) \tag{4.11}$$

The estimation of the long memory parameter d is achieved by a 2-step procedure. The 2-step exact local Whittle estimator $\hat{d}_{2ELW}$ is given by

$$\hat{d}_{2ELW} = \hat{d}_T - R'_F(\hat{d}_T)/R''_F(\hat{d}_T), \tag{4.12}$$

where $\hat{d}_T$ is the first step estimator and $R_F(d)$ is the modified objective function given in Equation (4.11). The two-step ELW estimator $\hat{d}_{2ELW}$ can be extended to the cases where the data have a polynomial time trend with an unknown mean:

$$X_t = \mu_0 + \beta_{10}t + \beta_{20}t^2 + \cdots + \beta_{k0}t^k + X_t^0; \quad X_t^0 = (1-L)^{-d_0}u_t\mathbf{I}\{t \geq 1\} \tag{4.13}$$

The long memory parameter d can be estimated by applying the two-step estimation to the residuals $\hat{X}_t$. We will denote the two-step ELW estimator with detrending as $\hat{d}_{2ELWdetrend}$.

4.3.2 PSY Approach

Phillips, Wu, and Yu (2011) develop a sup ADF (SADF) procedure that can test for evidence of price exuberance and date stamp its origination and collapse. Such a test procedure makes use of a right-tailed unit root and a sup test in a recursive way. The SADF test is recursively applied to the sample data and is implemented as follows. For each time series x_t, we apply the Augmented Dickey-Fuller (ADF) test for a unit root against the alternative of an explosive root (right-tailed). The following autoregressive specification for x_t is estimated by least squares:

$$x_t = \mu_x + \delta x_{t-1} + \sum_{j=1}^{J} \phi_j \Delta x_{t-j} + \varepsilon_{x,t}, \qquad \varepsilon_{x,t} \sim \mathrm{NID}(0, \sigma_x^2), \tag{4.14}$$

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 $H_1 : \delta > 1$. Equation (4.14) is estimated repeatedly using subsets of the sample data incremented by one additional observation at each pass in the forward recursive regression. Thus the SADF test is constructed by repeatedly estimating the ADF test. Let r_w be the window size of the regression. The window size r_w expands from r_0 to 1, where r_0 is the smallest sample window width fraction and 1 is the largest window fraction (the full sample). 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 ADF statistic for a sample that runs from 0 to r_2 is therefore denoted by $ADF_0^{r_2}$. The SADF statistic is defined as the sup value of the ADF statistic sequence:

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

Unlike the SADF test, the GSADF test is extended by using a more flexible window size. 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:

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

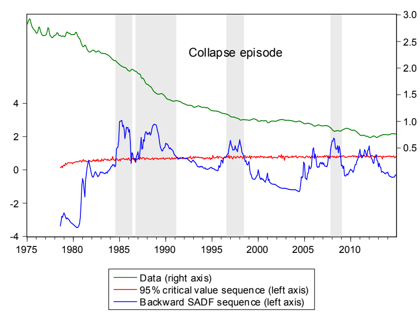
According to Phillips et al. (2015a), the minimum window size r_0 needs to be large enough to allow initial estimation, but it should not too large to miss the chance of detecting an early bubble period. We therefore follow Phillips et al. (2015a) and let $r_0 = 0.01 + 1.8/\sqrt{T}$, where T is number of observations. A small fixed lag order approach is used in this study as suggested by Phillips, Shi, and Yu (2015a). Hence, the lag order is chosen at 0 and 3 for the following analysis. The finite critical values are obtained from Monte Carlo simulation with 2000 replications.

Many studies have followed Phillips, Shi, and Yu's (2014) suggestion to include an intercept in the regression model. As a result, many empirical papers have reported rejections of the null suggesting periods of rapid increases in, for example, prices associated with a growing bubble, when in fact the data identifies a 'collapse' or a 'collapse and recovery' phase and not a bubble. An example of 'collapse episode' and 'collapse and recovery episode' can be seen in Figure 4.3 below. The backward SADF statistic (blue line) and its 95% critical value (red line) for Figure 4.3a suggest 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). Figure 4.3b presents data and test results consistent with a 'collapse and recovery' episode and a genuine 'bubble' or an 'exuberant' episode. The plot of the actual data makes the classification of these different episodes clear and highlights why the actual data and the test statistic (and relevant critical values) need to be presented on the same graph. It should be pointed out that both specifications inflate the the probability of an incorrect rejection of the null hypothesis. The inclusion of an intercept could increase the probability of an incorrect rejection of the null hypothesis (false positive) by detecting a collapse episode or a collapse and recovery episode. Phillips, Shi, and Yu (2014) also argue that omitting an intercept from the regression model inflates the probability of a false positive. In this paper, we consider two different regression model specifications (a model without an intercept and a model with an intercept) to explore evidence of bubbles and compare the results obtained from both formulations. This paper compares the bubble detection results using the aforementioned model specifications, see Hu and Oxley (2018a).

4.4 Results

Here we investigate the evidence of exuberant episodes or bubbles (where relevant, i.e. when prices can be compared to a 'fundamental') in historical housing price indices for the Herengracht index of Amsterdam, Norway and Paris based upon

(a) Collapse episode



(b) Collapse and recovery episode and bubble

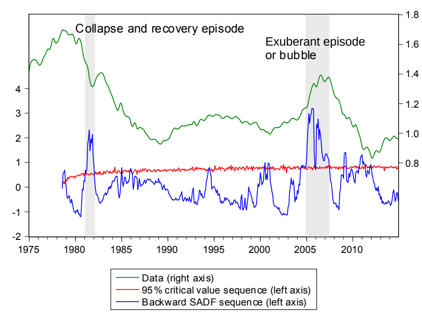


Figure 4.3 – Examples of collapse episode, collapse and recovery episode and bubble/ exuberant episode.

first the long memory estimates to be followed by the right-tailed unit root tests, where in the latter case we report results with or without an intercept term in the regression model. We apply the PSY to both the Herengracht index (1649-2010) and biennial Herengracht index (1628-1973) to check the robustness of the test on data with different frequencies. In addition, we also apply three semi-parametric long memory procedures to estimate long memory parameter d for assessing the presence of bubbles or explosive behavior.

4.4.1 Long Memory Results

We follow the discussion of Ramanan (2016) using the estimated values of the fractional integrating parameter d to assess explosive behavior in housing markets. If d lies between zero and one, the process is mean reverting and has no rational bubbles in housing prices. If d is greater than 1, house prices exhibit bubble-like behavior. Three long memory parameter estimates are used for estimating the long memory parameter d including the GPH estimator $\hat{d}_{GPH}$, local Whittle estimator $\hat{d}_{LW}$, the 2-step exact local Whittle estimator with detrending $\hat{d}_{2ELWdetrend}$. The corresponding long memory results are presented in Table 4.1. Section 4.4.1.1 discusses the long memory parameter estimates and Section 4.4.1.2

presents the standard errors for different estimators.

4.4.1.1 Different Estimators

We first focus on the Herengracht index in Amsterdam. As presented in Table 4.1, the estimated values of the long memory parameter for the Herengracht index (1649-2010) estimated by three semi-parametric estimators are 0.7328 ($\hat{d}_{GPH}$), 0.6252 ($\hat{d}_{LW}$) and 0.7947 ($\hat{d}_{2ELWdetrend}$), respectively. Overall, it is reasonable to conclude that the long memory parameter (d) is less than 1, indicating no explosive behavior. There is no exuberance in the Herengracht index for this period.

Results from Table 4.1 also suggest no evidence of explosive behavior in the biennial Herengracht index (1628-1973) as the three estimators are all less than 1, suggesting no explosive behavior (e.g., $\hat{d}_{GPH}$: 0.7891, and $\hat{d}_{LW}$: 0.6544 and $\hat{d}_{2ELWdetrend}$: 0.9749). A similar conclusion may be drawn for the Amsterdam house price-rent ratio between 1628 and 1850 as two estimates are less than 1 (e.g., $\hat{d}_{LW}$: 0.7434 and $\hat{d}_{2ELWdetrend}$: 0.6074). The exception is $\hat{d}_{GPH}$ (1.0007). Both the log periodogram regression of Geweke and Porter-Hudak (1983) and local Whittle estimation approaches of Robinson (1995) have nonstandard behavior for $d \in \left[\frac{3}{4}, 1\right]$ and these estimators are inconsistent for $d > 1$, see Phillips and Shimotsu (2004) and Phillips (2007). Hence, $\hat{d}_{2ELWdetrend}$ provides a consistent and unbiased estimate for d. We, therefore, conclude no explosive behavior in the biennial Herengracht index.

If we consider the long memory results for the aggregate Norwegian house price index (1819-2014), the $\hat{d}_{2ELWdetrend}$ is 1.0660, indicating explosive behavior. In contrast, $\hat{d}_{GPH}$ and $\hat{d}_{LW}$ are 0.7952 and 0.8191, respectively. Overall, there is evidence of explosive behavior in the aggregate Norwegian house price index based on long memory estimator $\hat{d}_{2ELWdetrend}$, but not the other two methods. If we now consider long memory results for Bergen (1819-2014), $\hat{d}_{2ELWdetrend}$ is 1.0241 while $\hat{d}_{GPH}$ and $\hat{d}_{LW}$ are less than 1 (e.g., 0.8313 and 0.8216). Given the properties of the estimators, we conclude there is evidence of explosive behavior in the house price index in Bergen. Interestingly, all three estimators are less than 1 for Kristiansand (1867-2014) and, therefore, suggest no sign of explosive behavior (e.g., $\hat{d}_{GPH}$: 0.7719, $\hat{d}_{LW}$: 0.9514 and $\hat{d}_{2ELWdetrend}$: 0.8674). Oslo also exhibits no explosive behavior with all three estimators are less than 1 (e.g., $\hat{d}_{GPH}$: 0.2144, $\hat{d}_{LW}$: 0.5216 and $\hat{d}_{2ELWdetrend}$: 0.9602) from 1841 to 2014. We do, however, find evidence of explosive behavior in the house price index in Trondheim (1897-2014) as $\hat{d}_{GPH}$ and $\hat{d}_{2ELWdetrend}$ are greater than 1 although $\hat{d}_{LW}$ is 0.7956.

If we consider the results for Paris (1650-2012), we find mixed evidence of explosive behavior. The two-step exact local Whittle estimator $\hat{d}_{2ELWdetrend}$ is greater than 1 while $\hat{d}_{GPH}$ and $\hat{d}_{LW}$ are 0.7701 and 0.6522, respectively.

4.4.1.2 Estimate Standard Errors

As shown in Table 4.1, confidence intervals of $\hat{d}_{LW}$ and $\hat{d}_{2ELWdetrend}$ for the Herengracht index do not contain 1, suggesting no explosive behavior. The confidence interval for $\hat{d}_{GPH}$ includes 1 with a standard error of 0.1681. We know from Rea et al. (2013) that the GPH estimator is very inefficient in small samples and thus $\hat{d}_{GPH}$ is perhaps not a reliable estimator in this case. On a close inspection of Table 4.1, $\hat{d}_{GPH}$ has a relatively large standard error compared with those of other estimators. We also take into account variability of long memory parameter for the biennial Herengracht index by constructing confidence intervals of three estimators. There is no significant evidence to rule out explosive behavior as the upper bound of confidence interval for $\hat{d}_{2ELWdetrend}$ exceed 1. The confidence interval for Herengracht's price rent ratio also do not contain 1 for $\hat{d}_{LW}$ and $\hat{d}_{2ELWdetrend}$, indicating no sign of explosive behavior.

The evidence of explosive behavior for Norway is mixed based on the three confidence intervals. Confidence intervals for $\hat{d}_{GPH}$ and $\hat{d}_{LW}$ do not exceed 1 while the confidence interval for $\hat{d}_{2ELWdetrend}$ ranges from 0.9 to 1.2. Hence the evidence of explosive behavior based on $\hat{d}_{2ELWdetrend}$ is not very strong despite the fact that

$\hat{d}_{2ELWdetrend}$ is greater than 1. Similarly, the evidence of explosive behavior for Bergen is also not very strong as the lower confidence interval for $\hat{d}_{2ELWdetrend}$ is less than 1. This phenomenon can be observed for Kristiansand and Oslo as well, where the lower bound of confidence interval for $\hat{d}_{2ELWdetrend}$ is less than 1 and the upper bound is above 1. No significant evidence of explosive behavior for Paris can be found based on the confidence interval of $\hat{d}_{2ELWdetrend}$.

Based on results from Section 4.4.1.1, Amsterdam data shows little evidence of explosive behavior. Norwegian data exhibits explosive behavior at the national and city levels and Parisian data also seems to support exuberance in house prices based on long memory estimate of $\hat{d}_{2ELWdetrend}$. Section 4.4.1.2 discussed the variability of long memory estimators by constructing their confidence intervals. Due to the small sample size of the housing series, when we consider the often wide confidence intervals for the three long memory estimators, we would now conclude that there is no overwhelming evidence to support explosive behavior in general, which is somewhat different to the findings based on only the mean value of the long memory estimates in Section 4.4.1.1.

4.4.2 PSY Results

4.4.2.1 Herengracht Index in Amsterdam (1649-2010)

The date-stamping outcomes for the real Herengracht index in Amsterdam are provided as Figure 4.4. We firstly present results based on the lag order of 0 under the assumption 'with an intercept' and 'without an intercept' in Figure 4.4a and Figure 4.4b, respectively. According to these two figures, there is no evidence of exuberant episodes in the real Herengracht index between 1649 and 2010.

Consider now results based on a lag order of 3 (in the PSY test) for the real Herengracht index under both assumptions as Figure 4.4c and Figure 4.4d, respectively. Using the model specification 'with an intercept', the null hypothesis of no explosive behavior is not rejected at the 10% significance level. The test statistic is

Table 4.1 – Long memory parameter estimates for different housing markets (standard errors in parentheses).

Housing Market	Long memory estimators		
	$\hat{d}_{GPH}$ (S.E.)	$\hat{d}_{LW}$ (S.E.)	$\hat{d}_{2ELW detrend}$ (S.E.)
Amsterdam			
Herengracht index (1649-2010)	**0.7328** (0.1681)	**0.6252** (0.0613)	**0.7947** (0.0636)
95% C.I.	[0.4033, 1.0623]	[0.5051, 0.7453]	[0.6700, 0.9194]
Biennial Herengracht index (1628-1973)	**0.7891** (0.4289)	**0.6544** (0.0916)	**0.9749** (0.0823)
95% C.I.	[-0.0515, 1.6297]	[0.4749, 0.8339]	[0.8136, 1.1362]
Herengracht's price-rent ratio (1628-1850)	**1.0007** (0.1366)	**0.7434** (0.0802)	**0.6074** (0.0753)
95% C.I.	[0.7330, 1.2684]	[0.5862, 0.9006]	[0.4598, 0.7550]
Norway			
House price in Norway (1819-2014)	**0.7952** (0.1018)	**0.8191** (0.0788)	**1.0660** (0.0788)
95% C.I.	[0.5957, 0.9947]	[0.6647, 0.9735]	[0.9116, 1.2204]
House price in Bergen (1819-2014)	**0.8313** (0.1123)	**0.8267** (0.0796)	**1.0241** (0.0788)
95% C.I.	[0.6112, 1.0514]	[0.6707, 0.9827]	[0.8697, 1.1785]
House price in Kristiansand (1867-2014)	**0.7719** (0.0796)	**0.8393** (0.0885)	**0.8674** (0.0870)
95% C.I.	[0.6159, 0.9279]	[0.6658, 1.0128]	[0.6969, 1.0379]
House price in Oslo (1841-2014)	**0.2144** (0.3357)	**0.5216** (0.0678)	**0.9602** (0.0822)
95% C.I.	[-0.4436, 0.8724]	[0.3887, 0.6545]	[0.7991, 1.1213]
House price in Trondheim (1897-2014)	**1.0600** (0.2121)	**0.7956** (0.1064)	**1.1939** (0.0941)
95% C.I.	[0.6443, 1.4757]	[0.5871, 1.0041]	[1.0095, 1.3783]
Paris			
House price in Paris (1650-2012)	**0.7701** (0.1358)	**0.6522** (0.0561)	**1.0277** (0.0635)
95% C.I.	[0.5039, 1.0363]	[0.5422, 0.7622]	[0.9032, 1.1522]

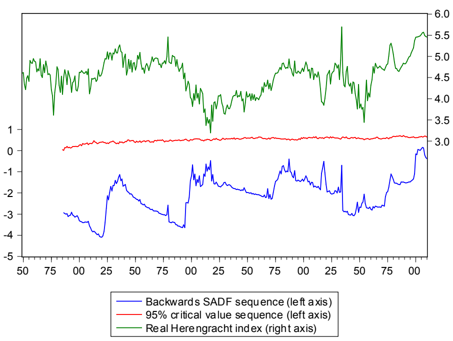
1.4978 which is smaller than the 10% right-tail critical value of 1.9504. As can be seen from Figure 4.4c, two episodes are identified: 1799-1804 (a 'collapse and recovery' episode) and 1812-1817 (a 'collapse' episode). Under the regression model 'without an intercept term', the null hypothesis of no explosive behavior is not rejected at the 10% significance level as the test statistic is lower than the 10% critical value (e.g., 2.4577<3.1855) even if an exuberant episode originated and collapsed during 1879-1889 is identified in Figure 4.4d. Under different formulations and lag order selection in the regression model, the above results suggest no evidence of exuberant episodes in the real Herengracht index and thus the Herengracht index alone is not explosive or 'exuberant'.

4.4.2.2 Biennial Herengracht Index (1628-1973)

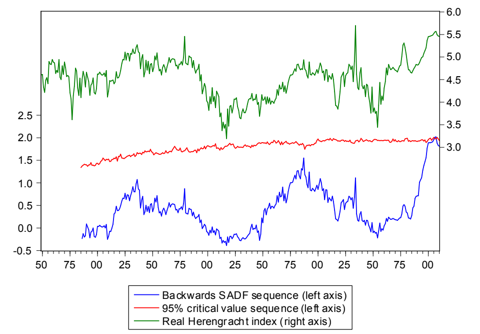
Empirical results for the real biennial Herengracht index are presented as Figure 4.5. Date-stamping outcomes based on the lag order of 0 for the real biennial Herengracht index under the assumption 'with an intercept' and 'without an intercept' are presented as Figure 4.5a and Figure 4.5b, respectively. Under the assumption 'with an intercept', the null hypothesis of no explosive behavior is not rejected at the 10% significance level as the test statistics (0.3134) is much smaller than the 10% right-tail critical value. As shown in Figure 4.5a, we observe no evidence of exuberance. Similarly, under the assumption 'without an intercept', the null hypothesis of no explosive behavior is not rejected at the 10% significance level (e.g., 2.4524<3.1219) although the presence of exuberance during 1878-1888 is detected in Figure 4.5b.

Date-stamping outcomes based on the lag order of 3 for the real biennial Herengracht index under the assumption 'with an intercept' and 'without an intercept' are presented in Figure 4.5c and Figure 4.5d, respectively. Under the assumption 'with an intercept', the test statistics is 1.4245. As indicated in Figure 4.5c, we observe three episodes in the real Herengracht index: 1800-1802 (a 'collapse' episode), 1812-1818 (a 'collapse and recovery' episode) and 1874-1888 (an 'exuberant' episode). On the other hand, under the assumption 'without an in-

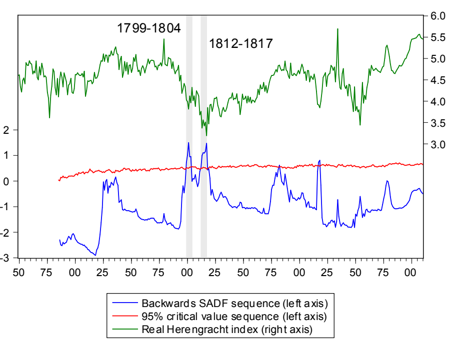
(a) real Herengracht index (1649-2010) with an intercept at the lag order of 0



(b) real Herengracht index (1649-2010) without an intercept at the lag order of 0



(c) real Herengracht index (1649-2010) with an intercept at the lag order of 3



(d) real Herengracht index (1649-2010) without an intercept at the lag order of 3

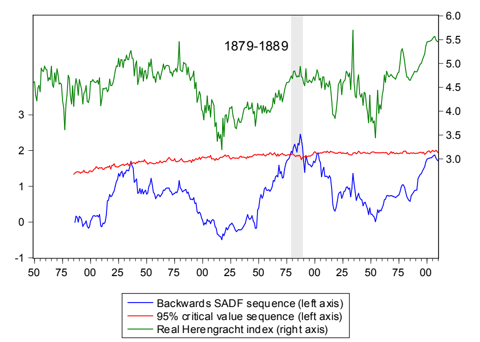
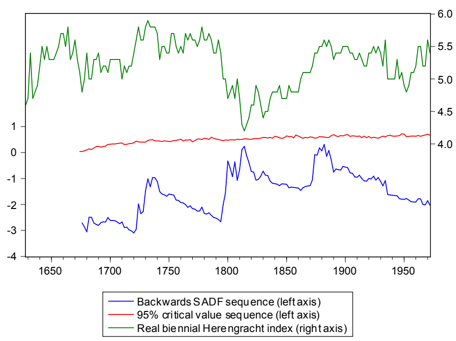


Figure 4.4 – Date-stamping strategy for the real Herengracht index (1649=100, log scale) between 1649 and 2010 using the GSADF test under the assumption with/without an intercept in the regression model with a lag order of 0 and 3.

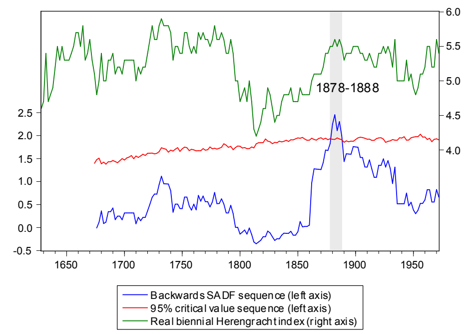
tercept', the test statistics is smaller than the 10% right-tail critical value (e.g., 2.7283<3.1219). We detect an exuberant episode between 1878 and 1888 as presented in Figure 4.5d.

We find some short-lived episodes between the mid-1870s and 1880s, which are clearly demonstrated in the real Herengracht index of Figure 4.4d (1880-1889)

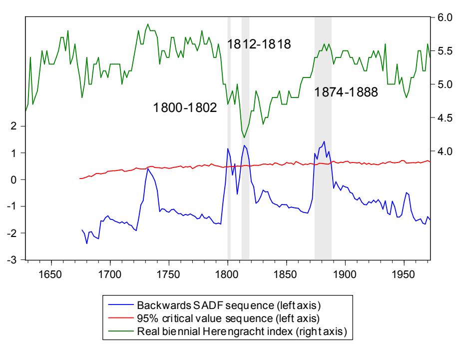
(a) real biennial Herengracht index (1628-1973) with an intercept at the lag order of 0



(b) real biennial Herengracht index (1628-1973) without an intercept at the lag order of 0



(c) real biennial Herengracht index (1628-1973) with an intercept at the lag order of 3



(d) real biennial Herengracht index (1628-1973) without an intercept at the lag order of 3

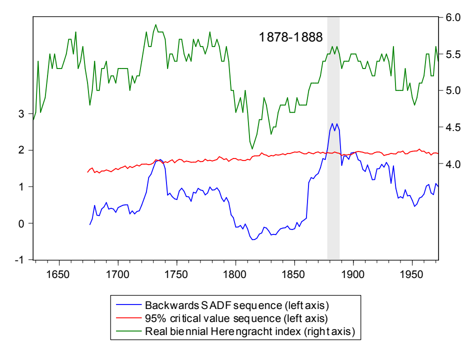


Figure 4.5 – Date-stamping strategy for the real biennial Herengracht index (1628=100, log scale) between 1628 and 1973 using the GSADF test under the assumption with/without an intercept in the regression model with a lag order of 0 and 3.

and the biennial Herengracht index of Figure 4.5b (1878-1888), Figure 4.5c (1874-1888), and Figure 4.5d (1878-1888). Overall, our results seem to suggest the biennial Herengracht index (1628-1973) is not exuberant. Such a finding is perhaps not surprising. The nominal index increases more than tenfold from 1628-1973 but the real value of the index at the end of period is a little more than double the value in 1628. This indicates that Herengracht index doesn't offer very positive real returns as it took nearly 350 years double the real biennial Herengracht index, suggesting an annual average price increase of only 0.2% a year (Shiller, 2006).

4.4.2.3 House Price-rent Ratio in Herengracht (1628-1850)

We also test for evidence of house price bubbles in the Amsterdam *house price-rent ratio* during the period 1628-1850. Date-stamping outcomes obtained from the GSADF test for the Amsterdam house price-rent ratio under the assumption 'with or without an intercept' in the regression model with the lag order of 0 and 3 are presented as Figure 4.6. We observe a short-lived bubble (1663-1665) in the price-rent ratio in Figure 4.6b. Overall, as can be seen from Figure 4.6 under different assumptions in the regression model, we find no evidence of bubbles in the *house price-rent ratio* for Amsterdam. The test statistic is not significant at the 10% level for all model specifications, suggesting no evidence of bubbles. The historical house price-rent ratio during this period in Amsterdam is not explosive.

However, our results are not in line with those in Ambrose et al. (2013). Using the Amsterdam rent-price ratio from 1650 to 2005, Ambrose et al. (2013) conclude that the rent-price ratio deviates from its long run average for substantial periods of time. They find that the deviation of house prices from rents can be persistent and long-lasting and the correction will take decades to get back to an equilibrium. Ambrose et al. (2013) also conclude that "our analysis of the rent-price ratio reveals sustained periods of "bubble" and "crisis" conditions [in Amsterdam] that can continue without a corresponding correction (or crash)".

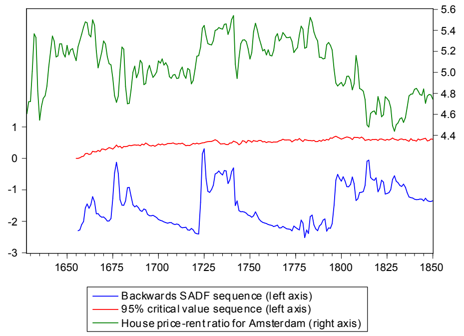
4.4.2.4 Discussion

Eichholtz (1997) argues that the Tulipmania episode had a very limited effect on the Amsterdam housing market as in 1632, the real biennial Herengracht house price index was 212.7. However, the real biennial Herengracht house price index fell to 113.5, which was almost half its value in 1632. The Tulipmania bubble burst in 1637. However, at the same time, the real biennial Herengracht house price index started to recover reaching 136.8 in 1639. Eichholtz (1997) argues that a pest epidemic likely had more impact on the Amsterdam housing market as Amsterdam lost 14% of its population due to the epidemic in 1636 alone.

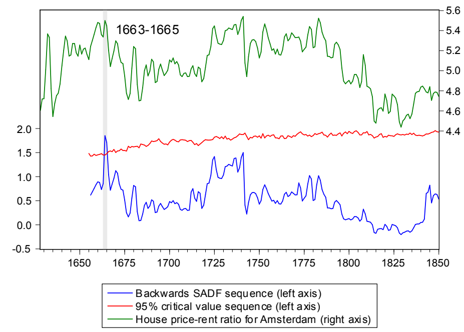
Our findings of an absence of housing bubbles in the Amsterdam house price-rent ratio (1628-1850) is not surprising as the two house price series (the Herengracht index and biennial Herengracht index) are not exuberant during this period. A finding of 'exuberance' in house prices seems to be a necessary property of a bubble, but not sufficient (without considering fundamentals). As Amsterdam house prices experience no exuberant episodes, we would not expect the presence of housing bubbles in the price-rent ratio during the same period.

There were four other famous bubbles in the eighteenth century: the Mississippi Bubble (1719-1720), the South Sea Bubble (1720), the Amsterdam Banking Crisis of 1763 and the Credit Crisis of 1772 (see Sheridan (1960), Garber (2001), Schnabel and Shin (2004), Kindleberger and Aliber (2011), Brunnermeier and Schnabel (2016)). Among these financial bubbles, the Mississippi Bubble and the closely related South Sea Bubble are the most well-known examples in the literature. The Amsterdam Banking Crisis of 1763 originated in Amsterdam and spread to Hamburg, Berlin, London. The Crisis of 1772 was more widespread as it spread to the Continent of Europe including Amsterdam. Most importantly, empirical results from Figure 4.4 and Figure 4.5 suggest these four crises do not have a great impact on real house prices in Amsterdam as we could not find evidence of either collapse episodes or exuberant episodes in the two Herengracht series. Hence, these historical bubbles/crises do not seem to spill-over to the Amsterdam housing markets, which is a very interesting finding and until now an issue that has not been addressed in the literature. Overall, it seems that there is little independent evidence for us to expect to observe the rent-price ratio reveals sustained periods of "bubble" and "crisis" (Ambrose et al., 2013) and our test results do not overturn this expectation.

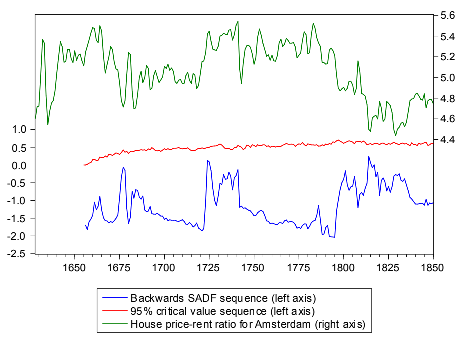
(a) House price-rent ratio (1628-1850) with an intercept at the lag order of 0



(b) House price-rent ratio (1628-1850) without an intercept at the lag order of 0



(c) House price-rent ratio (1628-1850) with an intercept at the lag order of 3



(d) House price-rent ratio (1628-1850) without an intercept at the lag order of 3

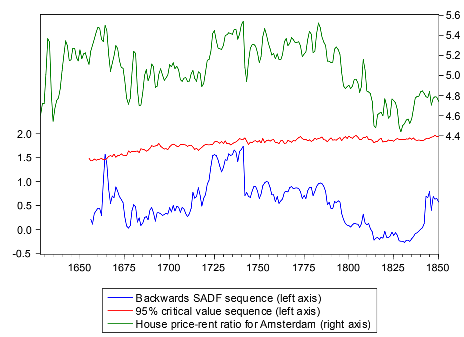


Figure 4.6 – Date-stamping outcomes obtained from the GSADF test for the house price-rent ratio in the Herengracht index (1628=100, log scale) from 1628 to 1850 under the assumption with or without an intercept in the regression model with a lag order of 0 and 3.

4.4.2.5 House Price Index in Norway (1819-2014)

Here, we investigate evidence of explosive behavior in the real Norwegian house price index between 1819 and 2014. Empirical results based on the PSY procedure for the real Norwegian house price index are presented as Figure 4.7 under various regression model specifications (e.g., with/without the intercept and the lag order choice). As previously, we firstly present results based on the lag order of 0 for two model specifications. Considering an intercept in the regression model, the test statistics is larger than the 5% right-tail critical value (i.e., 2.1488>2.1261), providing evidence of exuberance in the real Norwegian house price index. The corresponding date-stamping outcomes are shown in Figure 4.7a and two explosive episodes are identified (e.g., 1985-1988 and 2004-2014). Without including the intercept in the regression model, the test statistic is larger than the 10% right-tail critical value (i.e., 3.1718>3.1670), indicating some evidence of house price exuberance. As suggested in Figure 4.7b, we identify additional exuberant episodes in the late nineteenth century (1875-1879 and 1890-1903) along with more recent evidence of exuberance during 1981-1989 and 2000-2014.

Results under the assumption of with/without an intercept at the lag order of 3 are presented in Figure 4.7c and Figure 4.7d, respectively. The null hypothesis of no explosive behavior is rejected at the 5% significance level (i.e., 2.3836>2.1261), suggesting evidence of exuberance in Figure 4.7c. We can observe several exuberant episodes from Figure 4.7c: 1859-1860, 1873-1901, and 1982-1989. during the 19th-20th centuries. On the other hand, without considering the intercept in the regression model, the test statistic is also greater than the 5% significance level (e.g., 3.5987>3.5445). As can be seen from Figure 4.7d, there are two exuberant episodes (i.e., 1869-1918 and 1980-1990).

Overall, under various regression model specifications, our results seem to suggest evidence of exuberance. Several exuberant episodes have been identified in Figure 4.7 and a summary of these episodes can be found in Table 4.2. Our identified episodes generally coincide or overlap with several major financial crises in Norway as discussed in Grytten and Hunnes (2010), who conclude in favour of nine 'devastating financial crises' in almost 200 years of Norwegian history (e.g., 1814-1839, 1847-1850, 1856-1861, 1875-1888, 1899-1905, 1920-1928, 1930-1933, 1987-1993, and 2007-2010). Therefore, our identified episodes are highlighted in bold as shown in Table 4.2 where our empirically identified episodes coincide or overlap with these historical financial crises.

Among these identified episodes (including both collapse episodes and exuberant episodes), three exuberant episodes are of particular interest and worth discussing in detail. The first two exuberant episodes coincide with two banking crises in Norway, see Gerdrup (2004) and Brunnermeier and Schnabel (2016). Gerdrup (2004) discusses three major banking crises in Norwegian history during the 20th century (e.g., 1899-1905, 1920-1928 and 1988-1993). Brunnermeier and Schnabel (2016) review 23 of the most important bubble episodes from the past 400 years including the Norwegian crisis (1895-1900) and the Scandinavian crisis in Norway (1984-1992). The third episode is related to the recent run-up in house prices during the 2000s. Those three identified exuberant episodes from the long-term aggregate house price series coincide with the history of financial developments in Norway, which should be of considerable interest.

The first exuberant episode that should be highlighted is between 1890 and 1903 in Figure 4.7b under the model without an intercept at the lag order of 0, which is related to the Kristiania Crisis or known as the first banking crisis in Norway. In particular, as discussed in Grytten and Hunnes (2010), the Kristiania Crisis (1899-1905) was one of nine major financial crises in Norway. However, such an exuberant episode is not identified under the model 'with an intercept' in Figure 4.7a. Two long lasting exuberant episodes (i.e., 1873-1901 and 1869-1918) are identified in Figure 4.7c and Figure 4.7d, respectively.

Gerdrup (2004) studies three booms and busts involving banking crises in Norway (1899-1905, 1920-1928 and 1988-1993) and finds that strong bank expansion and asset price inflation are common features during these crises. In particular, as discussed in Grytten and Hunnes (2010), the Kristiania Crisis (1899-1905) was one of nine major financial crises in Norway. The economic development of the late 1980s-early 1900s was linked to a spectacular real estate boom and bust in Oslo and other large Norwegian cities (Gerdrup, 2004, p.149). As shown in the Figure on p.161 of Gerdrup (2004), house prices in Oslo rose more than 20% per year between 1896 and 1899. The failure of a highly leveraged non-financial firm accelerated the collapse in asset markets. As credit markets were affected, a real estate crash took place in several Norwegian cities due to bank failures at the end of the boom. Hence the identified exuberant episode between the late 19th century and the early 20th century (e.g., 1890-1903 in Figure 4.7b) coincides with the first major banking crisis in Norway during the 19th century that triggered by a real estate crash.

Under all four model specifications in Figure 4.7, PSY also successfully identifies another two exuberant episodes in the aggregate house price index: one in late 1980s, and the other one in the late 2000s. The crisis between the 1980s and early 1990s had its roots in the structural imbalances that developed in the 1970s and 1980s, which stemmed from a heavily regulated financial system after World War II to the market-based system in the mid-1980s (Gerdrup, 2004). The Norwegian housing market was heavily regulated on quantity and prices after World War II, and these regulations ended in July 1982 (Anundsen and Jansen, 2013). The Norwegian credit market regulations were also gradually lifted in the 1980s. Anundsen and Jansen (2013) conclude that the combined effect of liberalization processes (the deregulation on housing markets and credit markets) lead to a boom in the real estate market. Krogh (2010) overviews the changes of regulations in Norwegian credit markets in 1970-2008 especially during the strict credit market regulations in the 1970s and the gradual deregulation of credit markets in the 1980s. Financial deregulations tend to increase the availability of credit and therefore freed banks and other financial institutions to promote a lending boom. During the period of credit liberalization in the 1980s the house price indices also increase sharply in real terms (Eitrheim and Erlandsen, 2004). The financial deregulations on banks' lending resulted in a bank lending boom, which was accompanied by a boom in both residential and non-residential real estate (Vale, 2004).

According to Gerdrup (2004), house prices declined as much as 1/3 in real terms from 1988 and 1992. Allen and Gale (1999) argue that the collapse of an asset price bubble and the banking crisis were due to the decline in oil prices. The collapse of house prices in the late 1980s possibly related to the Norwegian banking crisis. The Norwegian banks were exposed to little credit risk during the heavily regulated period 1945-1984 as banks could not easily expand their lending. However, banks expanded their lending rapidly when the quantitative regulation was lifted. According to Reinhart and Rogoff (2008), the Norwegian Banking Crisis (1988-1993) is one of the five major bank-centered financial crises in history. The crisis peaked in 1991 with the second and fourth largest banks losing all their capital with a total market share of 24% and the largest bank getting into serious trouble (Vale, 2004). Grytten and Hunnes (2010) conclude that the crisis in 1980-90s was the most severe financial crisis in Norway since the 1930s and the worst banking crisis since the 1920s. Hence, our identified episode during the 1980s is in line with one of the major banking/financial crises.

An exuberant episode during the 2000s is identified in Figures 4.7a, 4.7b, 4.7c and 4.7d, suggesting the presence of exuberance in the real aggregate house prices. There has been a rapid increase in house prices since 2000 in Norway. In a report from the International Monetary Fund (2012), house prices in Norway were shown to grow at an annual rate of 11% during the period 2004-07, which was much higher than the OECD average of 5.5%. The price-to-rent ratio shows price overvaluation as Norway has the highest price-to-rent ratio at around 70% above its historical average, the highest level of almost all OECD countries. Similarly, the price-to-income ratio also indicates the presence of overvaluation as it is 28% above its historical average. More importantly, the report suggests that the Norwegian house prices are overvalued by $15\% - 20\%$. Another report from International Monetary Fund (2013) suggests that house prices in Norway may be overvalued by 40% and both price-to-rent and price-to-income ratios still indicate the signs of overvaluation. The authorities adopted tightening measures to cool down the housing market with Robert Shiller warning of the existence of bubbles in Norway's housing market in 2013. Jurgilas and Lansing (2013) also conclude

Table 4.2 – Episodes identified from the real house price index in Norway (1819-2014), where ** and * indicate the 5% and 10% significance level. Episodes are highlighted in bold if they are overlapped with major financial crises in Norway as discussed in Grytten and Hunnes (2010).

Country/Model Specification	**Test Stat**	**Episode(s)**
Lag order=0		
Norway (with an intercept)	2.149**	**1985-1988, 2004-2014**
Norway (without an intercept)	3.172*	**1874-1883, 1885-1904,**
		1982-1989, 2000-2014
Lag order=3		
Norway (with an intercept)	2.384	**1875-1882, 1894-1900**
		1986-1989
Norway (without an intercept)	3.599*	**1870-1917, 1978-1990**
		2010-2014

real house prices in Norway have risen by 30% since 2006. They also notice that the price-to-rent ratio for Norway from 1960 onwards reaches its highest level and the price-to-income ratio for Norway from 1980 onwards is higher than its previous peak. Moreover, the household leverage ratio (ratio of household debt to disposable income) remains historically high and around 210% above its level in 1980. The recent exuberant episode in the 2000s seems to be a global phenomenon see, Girouard et al. (2006), Kim and Renaud (2009) and Knoll et al. (2017). For example, Knoll et al. (2017) show that house prices in most industrial economies stayed constant in real terms from the 19th to the mid-20th century, but rose sharply in recent decades. To sum up, our finding of the recent exuberant episode during the 2000s seems to be in agreement with the conclusion drawn by Robert Shiller in 2013 and with the summary statistical analysis.

Table 4.3 – Episodes identified from the real house price index in Bergen and Kristiansand, where ***, ** and * indicate the 1%, 5%, and 10% significance level. Episodes are highlighted in bold if they are overlapped with major financial crises in Norway as discussed in Grytten and Hunnes (2010).

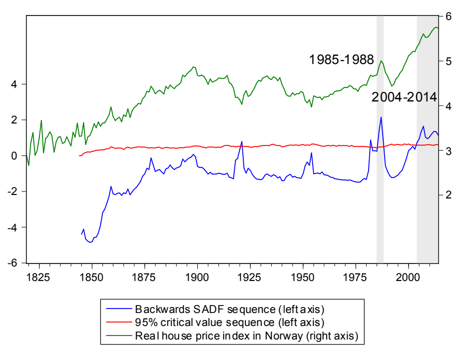
City/Model Specification	**Test Stat**	**Episode(s)**
Lag order=0		
Bergen (with an intercept)	1.9000*	**1986-1988, 2002-2014**
Bergen (without an intercept)	3.1720*	**1986-1988, 2001-2014**
lag order=3		
Bergen (with an intercept)	3.1008***	**1872-1900,** , **1983-1990, 2004-2014**
Bergen (without an intercept)	3.8260**	**1870-1918**, **1981-1990, 2003-2014**
Lag order=0		
Kristiansand (with an intercept)	0.3421	
Kristiansand (without an intercept)	3.0066	**2010-2014**
lag order=3		
Kristiansand (with an intercept)	3.2472***	**1920-1922, 2005-2014**
Kristiansand (without an intercept)	2.6735	1978-1984, **1986-1991** **1997-2014**

Table 4.4 – Episodes identified from the real house price index in Oslo and Trondheim, where ***, ** and * indicate the 1%, 5%, and 10% significance level. Episodes are highlighted in bold if they are overlapped with major financial crises in Norway as discussed in Grytten and Hunnes (2010).

City/Model Specification	Test Stat	Episode(s)
Lag order=0		
Oslo (with an intercept)	0.9853	**2006-2008**, 2012-2014
Oslo (without an intercept)	2.8547	**1898-1900, 2005-2014**
lag order=3		
Oslo (with an intercept)	1.0515	**1898-1900**
Oslo (without an intercept)	2.7117	**1891-1902**
Lag order=0		
Trondheim (with an intercept)	1.0730	**1918-1921, 2006-2008** 2011-2014
Trondheim (without an intercept)	6.3771***	**1985-1989, 2001-2014**
lag order=3		
Trondheim (with an intercept)	1.9726*	**1980-1989**
Trondheim (without an intercept)	4.0636**	**1978-1990**, 2011-2014

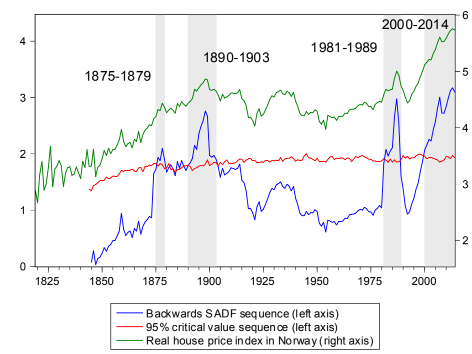
(a) Norway (1819-2014)

with an intercept at the lag order of 0



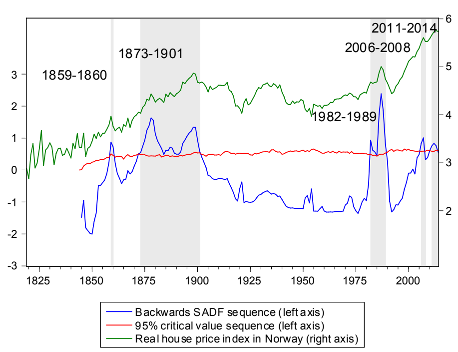
(b) Norway (1819-2014)

without an intercept at the lag order of 0



(c) Norway (1819-2014)

with an intercept at the lag order of 3



(d) Norway (1819-2014)

without an intercept at the lag order of 3

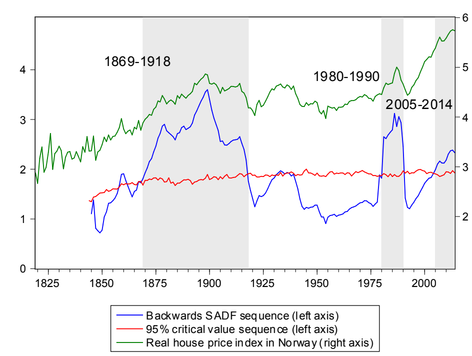


Figure 4.7 – Date-stamping strategy using the GSADF test for Norway under the assumption with/without an intercept at the lag order of 0 or 3.

4.4.2.6 City-level House Price Index

In this section, we consider analysis based on city-level house price indices using the PSY approach. Figures 4.8, 4.9, 4.10 and 4.11 illustrate the date-stamping strategies based on PSY for the real house price indices in Bergen, Kristiansand, Oslo, and Trondheim, respectively.

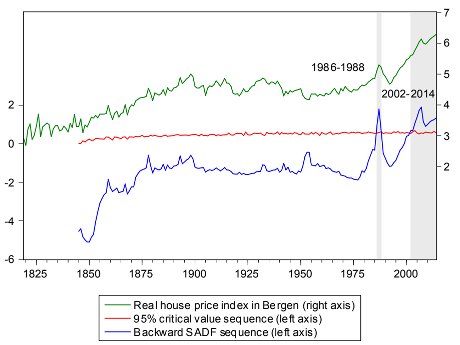
As listed in Table 4.3, the null hypothesis of no explosive behavior for the house price index in Bergen is rejected at least at the 10% significance level under different models. The exuberant episode in the late 1980s is identified in Figure 4.8 under all model specifications. The presence of an exuberant episode between the early or mid-2000s and 2014 is also clearly demonstrated in Figures 4.8a, 4.8b, 4.8c and 4.8d, and this finding seems to be in line with the recent house price boom in Norway. Moreover, the exuberant episode related to the first Norwegian banking crisis of the 19th century is also detected in Figure 4.8c and Figure 4.8d.

However, quite different results are obtained from Kristiansand. Table 4.3 suggests that the null hypothesis of no explosive behavior for Kristiansand can be rejected only at the 1% significance level for including the intercept in the regression model with a lag order of 3. We observe exuberance in Figures 4.9b, 4.9c and 4.9d. A 'collapse and recovery' episode is identified from 1919 to 1922 in Figure 4.9c. Overall, we may conclude no explosive behavior in the house price series.

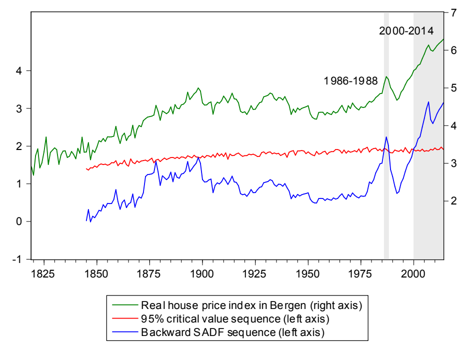
As shown in Table 4.4, we cannot reject the null hypothesis of no exuberance in the real house price for Oslo under all four models, indicating no significant evidence of exuberance. Despite the null hypothesis cannot be rejected, we identify exuberant episodes between the 1890s and early 1900s in Figures 4.10b, 4.10c, and 4.10d. We also find evidence of exuberance in the recent 2000s under the two regression model specifications with a lag order of 0 in Figures 4.10a and 4.10b.

We also find significant evidence of exuberance for Trondheim in Table 4.4 as the test statistic is significant at the 1% level when choosing the lag order of 0 in the regression model without the intercept, at the 10% or 5% level with a lag order of 3 under the assumption of including or excluding the intercept in the regression model. The corresponding date-stamping outcomes suggest exuberance in the 2000s (e.g., Figure 4.11a, Figure 4.11b and Figure 4.11d). An interesting finding from Figure 4.11a is that we identify a 'collapse' episode between 1918-1921. In addition, we find exuberant episodes during the 1980s in Figures 4.11b, 4.11c and 4.11d.

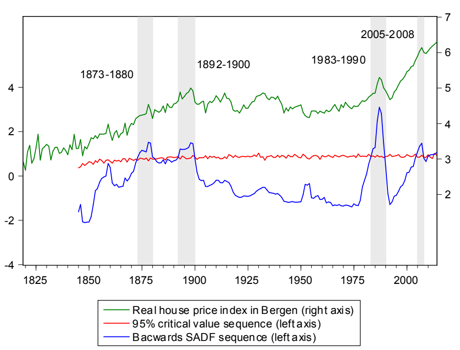
(a) Bergen (1819-2014)
with an intercept at the lag order of 0



(b) Bergen (1819-2014)
without an intercept at the lag order of 0



(c) Bergen (1819-2014)
with an intercept at the lag order of 3



(d) Bergen (1819-2014)
without an intercept at the lag order of 3

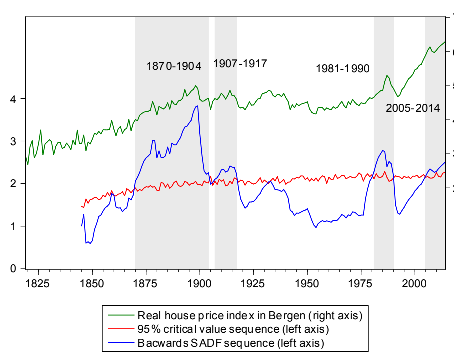
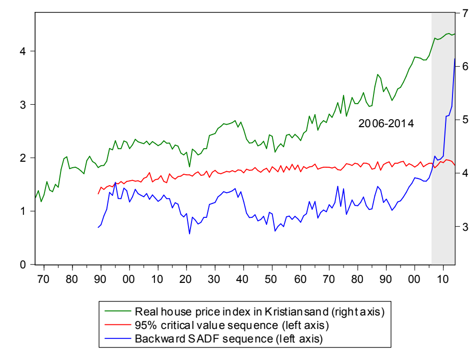


Figure 4.8 – Date-stamping strategy using the GSADF test for Bergen under the assumption with/without an intercept in the regression model with a lag order of 0 and 3.

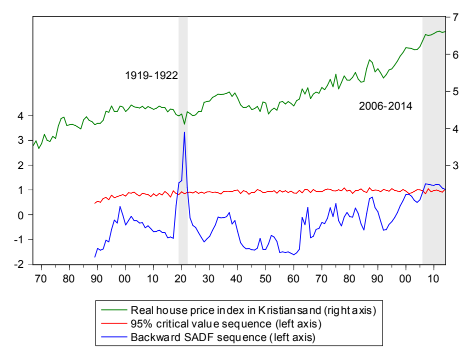
The majority of episodes identified from the city-level indices are consistent with those detected from the real aggregate house price index in Norway as displayed in Figure 4.7, the exceptions being 1918-1921 (Trondheim), and 1920-1922 (Kristiansand). The house price index declined in real terms during WWI; the nominal aggregate house price increased by 72% from 1914 to 1920 while the CPI rose by

(a) Kristiansand (1867-2014)
with an intercept at the lag order of 0

(b) Kristiansand (1867-2014)
without an intercept at the lag order of 0



(c) Kristiansand (1867-2014)
with an intercept at the lag order of 3



(d) Kristiansand (1867-2014)
without an intercept at the lag order of 3

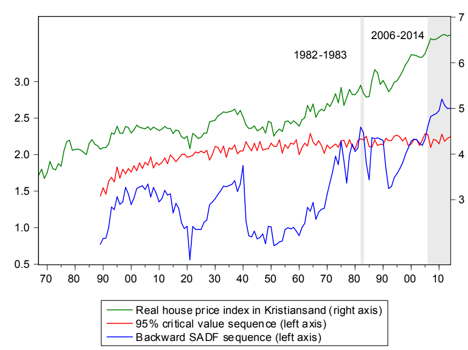
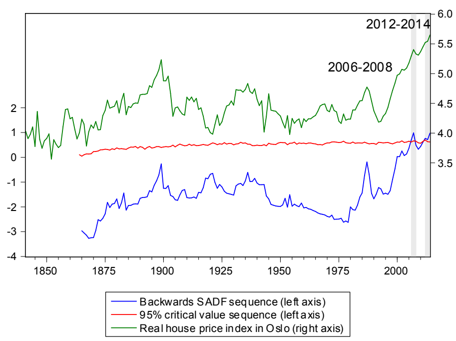


Figure 4.9 – Date-stamping strategy using the GSADF test for Kristiansand under the assumption with/without an intercept in the regression model with a lag order of 0 and 3.

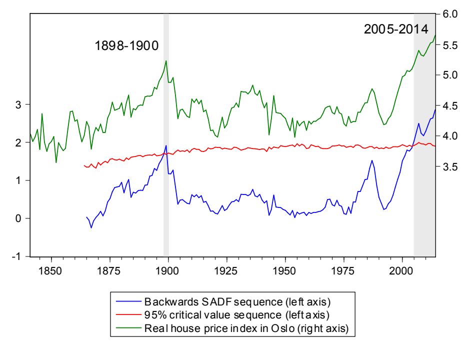
197% during this period (Eitrheim and Erlandsen, 2005). We find a collapse episode in Figure 4.11a for Trondheim and a collapse and recovery episode in Figure 4.9c for Kristiansand. These two 'collapse' episodes are hardly exuberant episodes and these episodes are likely caused by the 1916 introduction of the rent control law as many of the properties in the samples are rentals, see Eitrheim and

Erlandsen (2005).

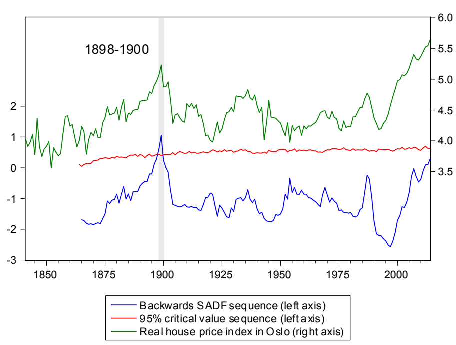
(a) Oslo (1841-2014)
with an intercept at the lag order of 0



(b) Oslo (1841-2014)
without an intercept at the lag order of 0



(c) Oslo (1841-2014)
with an intercept at the lag order of 3



(d) Oslo (1841-2014)
without an intercept at the lag order of 3

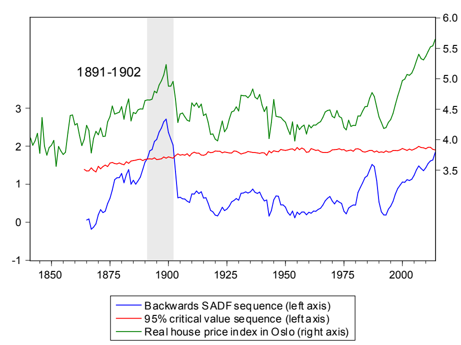
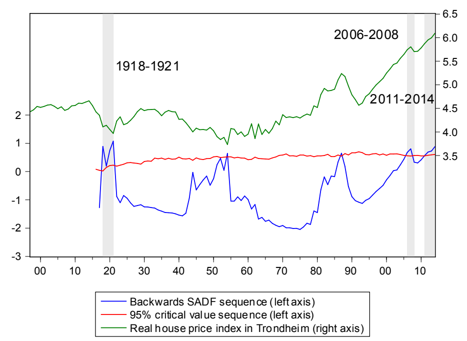


Figure 4.10 – Date-stamping strategy using the GSADF test for Oslo under the assumption with/without an intercept in the regression model with a lag order of 0 and 3.

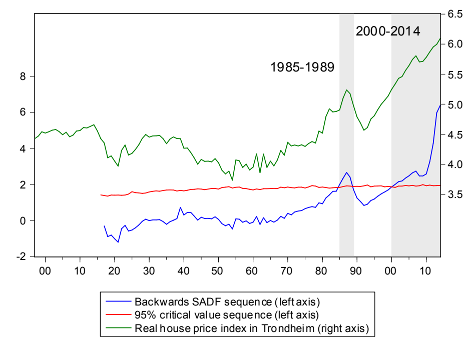
4.4.2.7 House Price in Paris (1650-2012)

We analyse the real house price index in Paris based on the sample period 1650-2012. The date-stamping outcomes under various regression model specifications

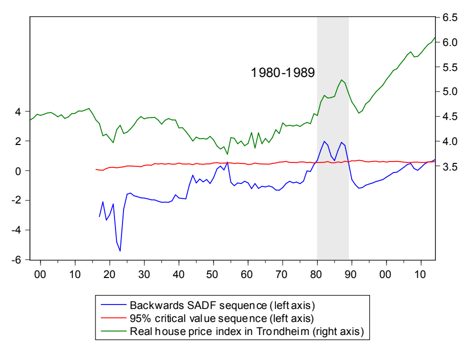
(a) Trondheim (1897-2014)
with an intercept at the lag order of 0



(b) Trondheim (1897-2014)
without an intercept at the lag order of 0



(c) Trondheim (1897-2014)
with an intercept at the lag order of 3



(d) Trondheim (1897-2014)
without an intercept at the lag order of 3

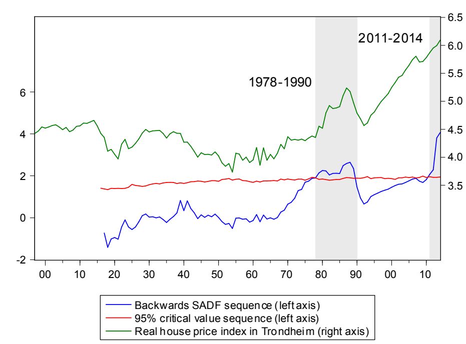


Figure 4.11 – Date-stamping strategy using the GSADF test for Trondheim under the assumption with/without an intercept in the regression model with a lag order of 0 and 3.

based on the PSY procedure are shown as Figure 4.12. Allowing an intercept in the regression model with a lag order of 0, the null hypothesis of no exuberance is rejected at the 1% level (4.4668>2.8755) and we find two 'collapse and recovery' episodes in Figure 4.12a (e.g., 1918-1921 and 1945-1953). Although the test statistic may suggest the rejection of the null hypothesis of no exuberance in the real house price index for Paris, these 'collapse' episodes are hardly believable exuberant episodes as these episodes are detected when the real house price index is continuously declining. This example shows the rejection of the null hypothesis of no exuberance in the PSY could lead to false positive identification of 'genuine' exuberant episodes/ bubbles if the plot of the actual data series is not taken into account. A short-lived exuberant episode between 2010 and 2012 is also detected.

Alternatively, we obtain quite difficult results when the intercept term is excluded from the regression model. There is evidence of house price exuberance in the real house price series for Paris as the null is rejected at the 5% level (3.7572>3.5501). As suggested in Figure 4.12b, we can identify an exuberant episode 1982-2012 and a spurious episode 1911-1915. Thus we find significant evidence of exuberance in the real house price index in Paris only under this model specification. This finding is very interesting as the house price seems to exhibit explosive behaviour only during very recent decades. The recent house price run-up phenomenon has been observed in some existing studies, see Kim and Renaud (2009) and Knoll et al. (2017).

When the lag order of 3 is considered in the regression model, we find little evidence of exuberance in the real house price series. The null of no exuberance cannot be rejected at the 10% level under the assumption with or without an intercept, respectively. For example, we find a short-lived 'collapse' episode in Figure 4.12c between 1946 and 1949 and an episode in Figure 4.12d between 1911 and 1915. The latter episode identified in Figure 4.12d is almost in line with the one detected in Figure 4.12b in 1910s.

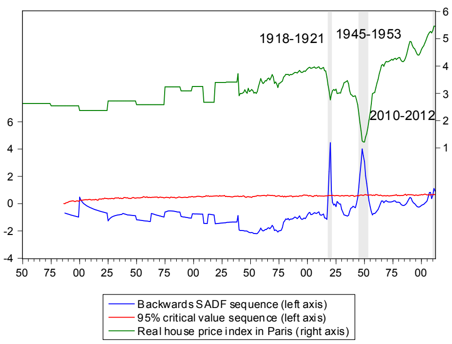
The Parisian data seem to suggest a nonlinear specification. Hence, assuming a linear autoregressive (AR) model could be misspecified. Upon closer inspection of the real house price series in Figure 4.12, the real house price declined significantly from 1942 and reached its lowest level in 1949/50. The decline in the real house price is due to the implementation of rent controls in times of high inflation, caused by World War II (Friggit, 2008, p.26). Because of this special feature, PSY identifies the collapse and recovery episode in Figure 4.12a and Figure 4.12c when the intercept term is considered in the regression model specification. As a result, PSY fails to identify the exuberant episode at the end of sample period under most model specifications. However, under the model without an intercept and the lag order of 0, we find evidence of exuberance in the historical house price index for Paris.

In order to shed insight into these experiences, we also explore whether some historical episodes might have impacted on the real house price index in Paris. First, we could not find evidence of house price exuberance during the Mississippi Bubble (1719-1720). The same conclusions are drawn for the Crisis of 1763 and the Crisis of 1772. Second, the crisis in the French stock market in 1882 was the worst in French history in the nineteenth century and such a crisis led to a deep recession that lasted until the end of the decade (White, 2007). Our empirical results suggest that the crisis of 1882 doesn't cause any explosive behavior in the housing markets. It seems to suggest that these crises have no significant impact on the real house price in Paris as no collapse episodes or exuberant episodes are detected. Third, according to White (2007), the Paris housing price bubbles started around 1884 and bust in 1900-1901. However, our empirical results suggest no evidence of house price exuberance during this period.

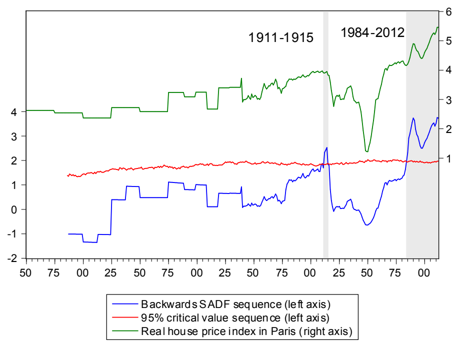
4.4.3 Summary

We also compare results from long memory and PSY approaches in Table 4.5. The Long memory and PSY results for different data are discussed in Section 4.4.1.1 and Section 4.4.2, respectively. As can be seen from Table 4.5, our results from PSY are generally consistent with those obtained from long memory models except the Paris series. Overall, we prefer the PSY approach as it could detect the presence of multiple bubbles.

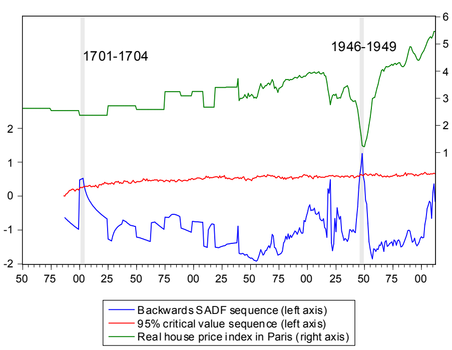
(a) Paris (1650-2012)
with an intercept at the lag order of 0



(b) Paris (1650-2012)
without an intercept at the lag order of 0



(c) Paris (1650-2012)
with an intercept at the lag order of 3



(d) Paris (1650-2012)
without an intercept at the lag order of 3

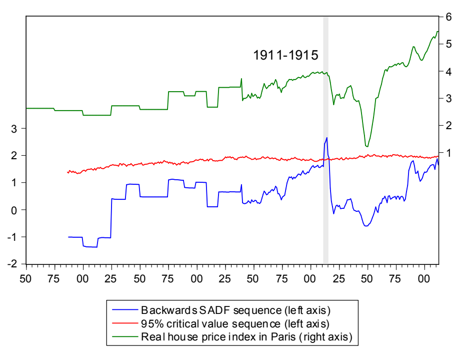


Figure 4.12 – Testing for explosive behaviour in Paris real house price index (2000=100, log scale) between 1650 and 2012 using the GSADF test under two different regression model specifications with a lag order of 0 and 3.

4.5 Conclusion

Recent debates relating to the role of sub-prime mortgage funded housing market booms in the US, has brought the concept of bubbles back into the public and academic arenas. Furthermore, the growing sophistication and integration of financial markets may have led to the rapid contagion of the original financial crisis

Table 4.5 – Summary of the results from the two approaches where N and Y stand for no explosive behavior and explosive behavior.

Data	Long Memory	PSY
Amsterdam		
Herengracht index (1649-2010)	N	N
Biennial Herengracht index (1628-1973)	N	N
Herengracht's price-rent ratio (1628-1850)	N	N
Norway		
House price in Norway (1819-2014)	Y	Y
House price in Bergen (1819-2014)	Y	Y
House price in Kristiansand (1867-2014)	N	N
House price in Oslo (1841-2014)	N	N
House price in Trondheim (1897-2014)	Y	Y
Paris		
House price in Paris (1650-2012)	Y	N

spreading not only spatially but across different markets. In this paper, we have sought to test whether housing market bubbles (or periods of very rapid price rises- 'exuberance') have historical precedents and furthermore, whether there is any evidence that bubbles or crisis in different financial markets (i.e., the Mississippi or South Sea bubbles spilled-over into local or national housing markets). Long memory models of Shimotsu and Phillips (2005) and Shimotsu (2010) and right-tailed unit root tests of Phillips, Shi, and Yu (2015a, PSY) are considered in this paper.

To consider these questions we tested for evidence of exuberance in three long-lasting house price indices, namely the Herengracht index of Amsterdam, Norwegian and Paris house price indices. In addition, we also investigated the presence of housing bubbles for the Amsterdam. Firstly, there is no evidence of exuberance in both the real Herengracht and biennial Herengracht indices. This finding indicates the real Herengracht index on its own is not exuberant. More importantly, our results do not support the existence of housing bubbles in the Amsterdam price-rent ratio during 1628-1850 despite the occurrences of several (other) historical bubbles. A finding of 'exuberance' in house prices seems to be a necessary property of a bubble but not sufficient (without considering fundamentals). As the Amsterdam house price experiences no exuberant episodes, we would not expect the presence of housing bubbles in the price-rent ratio during the same period. Secondly, we find evidence of exuberance in the Norwegian house price index and these episodes either coincide or overlap with several major financial crises in Norway as discussed in Grytten and Hunnes (2010). Our empirical results also suggest that the house price episodes identified from four main cities (Oslo, Bergen, Trondheim and Kristiansand) are generally in line with those episodes obtained from the real house price index at the national level. Of particular interest is that three identified exuberant episodes are related with the banking crises in 1895-1900 and 1988-1993, and the recent house price run-up in the 2000s. Thirdly, the real house price index of Paris exhibits significant evidence of exuberance under a particular model specification. Several crises including the Mississippi Bubble, the Crisis of

1763 and the Crisis of (1772) appear to have no significant impact on the real house price index in Paris. The Parisian series is of great interest as the data seem to suggest a nonlinear specification, which creates considerable difficulty in detecting exuberant episodes for PSY. Future development on bubble detection methodologies could consider this interesting series to check the robustness of new tests. Fourthly, our results seem to draw an important conclusion that housing price exuberance seems to be a very recent phenomenon as no explosive behaviour is detected in housing markets before 1850 based on three very well-known house price series. Fifthly, our results from PSY are generally in line with those obtained from long memory models based on the mean value of d. However, the number of observations in each sample is very small, leading to a wider confidence interval. When we take into account the confidence intervals for long memory models, there is no overwhelming evidence to support explosive behavior. Lastly, there is a need to be careful when interpreting empirical results from these new time series-based methods of PSY, minimally checking that failed test date-stamping have both an empirical (a ‘bubble’ not a ‘collapse and recovery’ episode) and where possible, some historiographically sourced supporting information.

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Chapter 5

Bubbles in US Regional House Prices: Evidence from House Price-Income Ratios at the State Level.

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Bubbles in US regional house prices: evidence from house price–income ratios at the State level

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ABSTRACT
We investigate the presence of bubbles in the US house price-income ratio at the State level by applying the recent time series-based econometric test to data from January 1975 to December 2014. We find evidence of bubbles in several States in the 1980s (i.e. California, Hawaii, Massachusetts, New York, etc.), which coincides with some existing studies that investigate housing bubbles or booms and busts using a range of alternative approaches. Our results show the existence of a housing bubble that originates in the early 2000s and collapses in the mid-2000s in more than 20 States and the District of Columbia concluding that the bubbles of the 2000s were more widespread than the 1980s, which is of special interest and importance. Our results seem to be in agreement with the talk given by Alan Greenspan in 2005, who suggest no sign of a nationwide housing bubble but a lot of local bubbles. We also study the importance of the regression model specification with/without an intercept and the regression model with an intercept could lead to false-positive identification of bubbles.

KEYWORDS
Bubbles; generalized sup ADF test; US regional house prices; house price–income ratio

JEL CLASSIFICATION
C59; R19; R39

I. Introduction

The US seems to have a long history of real estate speculation (Glaeser 2013). The purpose of this article is to investigate evidence of housing bubbles by comparing house price changes with changes in income, and whether the outcomes are generic, localized or create regionally constrained spillovers? To consider such questions we will utilize the recently developed bubble detection and date-stamping approach of Phillips, Shi, and Yu (2015, PSY). The PSY approach has important advantages over the conventional unit root and cointegration tests as this new method not only tests for the empirical existence (or otherwise) of bubbles but is also able to precisely 'date-stamp' their growth and collapse. Furthermore, it can also identify multiple bubble formations and collapses (and their precise dates), making it not only theoretically more interesting but also practically more useful. We will add to this approach by considering evidence based on their right-tailed unit root tests both with and without an intercept in the regression model. By making use of the PSY approach, we are able to identify points (in time and space) related to the origination and termination of any bubble event.

Our empirical study focuses on the 1980s and the early 2000s and has three main aims and contributions where it differs from existing studies in several respects. First, a major focus is to examine the econometric evidence for housing bubble(s) during the 1980s with particular emphasis on the States of California, Hawaii, Massachusetts and New York. Several other studies, for example, Case and Shiller (1988, 1994, 2003) and Wheelock (2006), consider US house price booms and busts or bubbles during the 1980s; however, most of these studies describe or graphically inspect some house price measure (e.g. house price, price–income ratio or price–rent ratio) without applying any econometric tests. Thus, our article fills the gap by providing empirical test-based evidence of housing bubbles during the 1980s. Second, more recently, much attention has been paid to the rapid increase in US house prices in the 2000s, see, for example, Del Negro and Otrok (2007), Shiller (2007), Goodman and Thibodeau (2008), Wheaton and Nechayev (2008), Mayer (2011). A number of States/areas experienced a dramatic boom during the 2000s and thus there is an increasing interest in

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testing for evidence of housing bubbles during this period using a range of approaches. Del Negro and Otrok (2007) see this recent rapid increase as a national phenomenon; however, Greenspan disagreed, stating that 'we don't perceive that there is a national bubble, but it's hard not to see that there are a lot of local bubbles' in 2005. Martin (2011) also concluded that the 2000s US housing bubble was not a national phenomenon. Our second aim will therefore consider whether there is evidence of a national or several local (disconnected) bubbles (or no bubbles at all) at the regional level in the US during the early-mid 2000s. To anticipate, we find evidence of a housing bubble that originates in the early 2000s and collapses in the mid-2000s in more than 20 States and the District of Columbia (DC) in our study. Our results show the bubble of the 2000s is not a national phenomenon, but is more widespread than the 1980s. This is perhaps the first empirical study to make a comparison in terms of their magnitude and coverage between the regional housing bubbles in the 1980s with the recent regional bubbles in the 2000s, which also contributes to the novelty of this article. Third, we also use the US State-level house price–income ratio data to study the importance of regression model formulation highlighted in Phillips, Shi, and Yu (2014) by exploring the role of the intercept in the regression model of Phillips, Shi, and Yu (2015). Based upon the model specification with an intercept, we can identify 'collapse' episodes, 'collapse and recovery' episodes and the potential bubbles. Whereas without an intercept in the regression model leads to identification of no 'collapse' episodes and 'collapse and recovery' episodes. An important finding in the regression model specification without the intercept is that several States do not exhibit any bubble-like behaviours for the whole sample period.

The article is organized as follows. Section II reviews some existing studies on US house price bubbles. Section III provides a brief description of the generalized sup ADF (GSADF) of Phillips, Shi, and Yu (2015) and Section IV describes the data. Section V provides empirical results for all the 50 States and the District of Columbia and Section VI concludes.

II. Literature review

The PSY testing procedure has become one of the most popular new approaches to detect asset price bubbles in the literature. Following the empirical studies of New Zealand and Australian housing markets (e.g. Greenaway-McGrevy and Phillips 2015; Shi et al. 2016) and theoretical developments on bubble detection procedure of Phillips and Shi (forthcoming), Hu and Oxley (2018) subject the famous South Sea and Mississippi Bubbles in 1719–1720, and the British Railway Mania of the 1840s, respectively. Hu and Oxley (2017b) also investigate the well-documented Japan's asset price bubble of the 1980-1990s. They find significant evidence of bubbles in the Japanese stock and housing markets, and signs of bubble contagion from Japan's stock market to its housing market for the first time in the literature. A series of papers also adopt the PSY testing procedure to examine the presence of bubbles in housing markets, see Chen and Funke (2013), Yiu, Yu, and Jin (2013), Jiang, Phillips, and Yu (2015) and Pavlidis et al. (2016).

A large number of studies have been tested for the existence of house price bubbles in the US. A house price bubble is defined as a situation when the growth of the price is not supported by changes in its fundamentals (e.g. Stiglitz 1990). A 'fundamental' in empirical studies of the housing market is often assumed to be either a rental cost–house price ratio, where the logic is that the rent represents the stream of future income from the housing asset, or a personal income–house price ratio, where the idea is that in the long-run house prices cannot exceed the ability to purchase the property or service the debt in the process. These two types of ratio lend themselves to what have become some of the most common forms of analysis of whether house prices deviate from 'fundamentals', which are based on tests for cointegration between the numerator and denominator series. For example, Malpezzi (1999) rejected the null hypothesis of no cointegration between house prices and income in US Metropolitan areas concluding there was no evidence of bubbles. McCarthy and Peach (2004) and Himmelberg, Mayer, and Sinai (2005) find little evidence of a bubble in US home prices. Gallin (2006) provides an excellent example of tests of the relationship between house prices and income in the US where using 27 years of US national-level data and finds no evidence of cointegration. As standard cointegration tests are known to have low power in small samples, Gallin (2006) applied several panel cointegration tests to a panel of 95 US Metropolitan areas over a 23-year period and also found no evidence of cointegration. In both cases, a finding of no cointegration relationship suggests evidence of bubbles. Zhou and Sornette (2006) investigated the existence of US housing bubbles at the regional and State levels using quarterly data, 1993–2005 where instead of tests for 'market fundamentals' they define a bubble as a 'faster-than-exponential price growth'. They concluded that 22 States exhibited evidence of a fast-growing bubble. Mikhed and Zemčík (2009) also used a panel test for the price–rent ratio in 23 US Metropolitan areas for the period 1978–2006 and concluded that there was evidence of a bubble. Holly, Pesaran, and Yamagata (2010) used Moon and Perron (2004) and Pesaran (2007) panel tests to consider the relationship between real house prices and real per capita disposable incomes in the US at the State level using annual data 1975–2003 and found little evidence of house price bubbles with a few exceptions (e.g. California, New York, Massachusetts, Connecticut, Rhode Island, Oregon and Washington). Empirical results from Kivedal (2013) suggest that there was a bubble in the US housing market prior to the 2007 subprime financial crisis. Nneji, Brooks, and Ward (2013) examined the presence of intrinsic bubbles of a Froot and Obstfeld (1991) type [1] or rational speculative bubbles of a Blanchard and Watson (1982) type in the residential property market in the US between 1960 and 2011. They split the data into two periods (1960–1999 and 2000–2011) and found an intrinsic bubble for the first period and a rational speculative bubble for the second period only. Escobari, Damianov, and Bello (2015) proposed a new test to identify house price bubbles, which explored a specific feature of the market such that low-tier house prices should appreciate more during the upswing of a growing boom and fall faster during the bust, and found evidence of bubbles in 15 US Metropolitan Statistical Areas (MSAs). It should be noted, however, that the methodology developed by Escobari, Damianov, and Bello (2015) does not consider 'market fundamentals' in assessing house price bubbles.

Several existing studies attempt to explain the surge in house prices after 2000. Case and Shiller (2003) argue that fundamentals play a crucial role in explaining much of the rapid increase in the housing market. In particular, income growth explains the house price increase in most States and a falling interest rate also contributes to the recent run-up in house prices. McCarthy and Peach (2004) and Himmelberg, Mayer, and Sinai (2005) both argue that the economic fundamentals such as low interest rates and high income growth could explain house prices growth in the early 2000s. Glaeser, Gyourko, and Saks (2005) focus on housing supply constraints and explain that rising housing prices have been accompanied by a declining housing supply due to changing local development regulations in certain areas. Malpezzi and Wachter (2005) also conclude that the impact of speculation is dominated by the effect of the price elasticity of supply. In fact, the large impacts of speculation are only observed when supply is inelastic. Ihlanfeldt (2007) finds that regulation restriction increases house prices and decreases vacant land prices in more than 100 Florida cities. Glaeser, Gyourko, and Saiz (2008) point out that areas with inelastic housing supply are associated with more increase in house prices.

Subprime mortgage lending is a financial innovation in commercial banking, see Frame and White (2014). During the spectacular rise and later collapse of house prices in the early 2000s, the dramatic increase in subprime lending has been blamed for the sharp rise and collapse in house prices. However, Coleman, LaCour-Little, and Vandell (2008) conclude that subprime loan products are not the primary blame for the housing bubble across 20 metropolitan areas over the period 1998–2006. Brueckner, Calem, and Nakamura (2012) also argue that subprime lending is a consequence rather than a cause of the housing bubble based on their theoretical model. Huang and Tang (2012) find that supply restrictions (residential land use regulations and geographic land scarcity) are linked to the booms and busts in the US house prices from 2000 and 2009.

III. Method

We apply the right-tailed unit root test of Phillips, Shi, and Yu (2015) to examine evidence of bubbles in US regional real estate markets. The martingale null with an asymptotic drift is specified as

$$H_0 : y_t = dT^{-\eta} + y_{t-1} + \varepsilon_t, \quad \varepsilon_t \sim \mathrm{NID}(0, \sigma^2), \tag{1}$$

[1] Intrinsic bubbles are driven solely by fundamentals (Froot and Obstfeld 1991).

where d is a constant and the localizing coefficient η is greater than 1/2. The alternative hypothesis is a mildly explosive process:

$$H_1 : y_t = \delta_T y_{t-1} + \varepsilon_t, \tag{2}$$

where $\delta_T = 1 + cT^{-\theta}$ with $c > 0$ and $\theta \in (0, 1)$.

The following regression model is estimated:

$$\Delta y_t = \alpha + \beta y_{t-1} + \sum_{i=1}^{K} \gamma_i \Delta y_{t-i} + \varepsilon_t, \tag{3}$$

where α is an intercept.

The GSADF test relies on repeated estimation of the ADF test regression of Equation (3) on subsamples of the data in a recursive fashion. The window size r_w expands from r_0 to 1, where r_0 is the minimum window size. The end point r_2 varies from r_0 to 1 and the starting point r_1 varies from 0 to $r_2 - r_0$. The GSADF statistic is the largest ADF statistic over the range of r_1 and r_2:

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

The backward SADF (BSADF) statistic is defined as the sup value of the ADF statistic sequence:

$$BSADF_{r_2}(r_0) = \sup_{r_1 \in [0, r_2 - r_0]} ADF_{r_1}^{r_2},$$

where the BSADF statistic and its corresponding critical value are used for dating the origination and termination dates of a bubble. The minimum window size r_0 is equal to $0.01 + 1.8/\sqrt{T}$. Critical values are simulated using 2000 replications.

The PSY approach is often applied to a price–fundamental ratio to assess the explosive behaviour. If the test statistic exceeds the critical value, we may conclude a finding of explosive behaviour. A finding of explosive behaviour in a price–fundamental ratio may be interpreted as evidence of bubbles. Previous applications (e.g. exchange rate market, stock market, housing market and commodity market) of the approach have followed Phillips et al.'s (2014) suggestion to include an intercept in the regression model. As a result, 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 identify 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 in the regression model. An example of 'collapse episode' and 'collapse and recovery episode' can be seen in Figure 1. The backward SADF statistic (blue line) and its 95% critical value (red line) for Figure 1(a) 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). Figure 1(b) presents data and test results consistent that relate to a 'collapse and recovery' episode and a genuine 'bubble'. The plot of the actual data makes the classification of these different episodes clear and highlight why the actual data and the test statistic (and relevant critical values) need to be presented on the same graph. Previous empirical studies have either ignored such cases or if they have mentioned them they have provided no explanation of the possible reason for the false test positives. Some empirical papers even obfuscate this issue by plotting

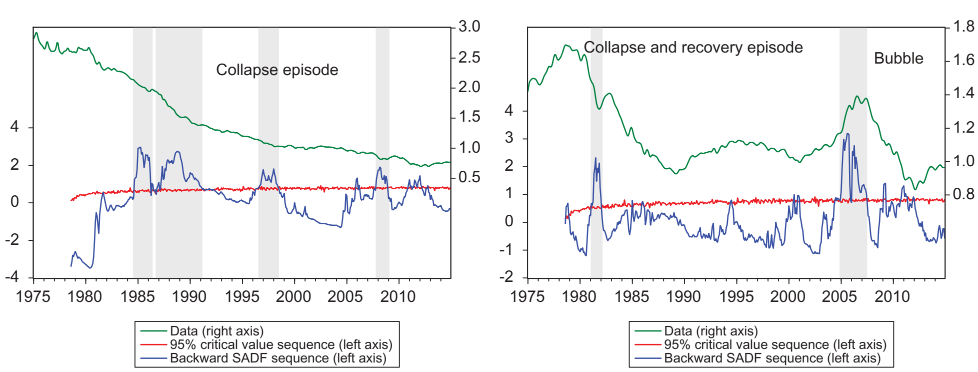

Figure 1. An example of (a) collapse episode and (b) collapse and recovery episode and bubble (For interpretation of the references to colour in this figure legend, the reader is referred to the web version of this article.).

only the backward SADF statistic with the 95% critical value sequences and provide no plot of the actual data series in the date-stamping strategy graph. In this article, we use two different regression model specifications (a model without an intercept and a model with an intercept) to explore the evidence of bubbles and compare the results obtained from both formulations. A small fixed lag order is used in both formulations as suggested by Phillips, Shi, and Yu (2015).

IV. Data

As US regional housing markets could behave quite differently, we focus here on State level data rather than national data. The monthly Freddie Mac State house price index provides a measure of house price inflation for the US in 50 State indexes and the District of Columbia. All series used are from January 1975 to December 2014 (December 2000 = 100). Quarterly State personal income for 50 States and the District of Columbia is obtained from the U.S. Bureau of Economic Analysis. Due to data frequency availability, we use the commonly adopted Chow and Lin's (1971) GLS procedure to interpolate the quarterly personal income series to create a monthly series.[2] All monthly personal income series have been calculated (December 2000 = 100). We are then able to calculate a monthly house price–income ratio for all States as well as the District of Columbia.

V. Results

We divide our results into several parts. The 'General results' section provides an overview of the bubble detection results for all the States and the District of Columbia. The results for several key States – California, District of Columbia, Massachusetts, Hawaii, Nevada and New York – are then discussed in the following subsections. Later, we also discuss results for farm and the 'Rust Belt' States and 'energy-producing' States.

General results

The price–income ratio for all the 50 States and the District of Columbia can be classified into two groups: one is oscillating with peaks and troughs (e.g. California, Florida, Hawaii, etc.) and the other exhibits a declining trend (e.g. Indiana, Mississippi, South Carolina, etc.).[3] Figure 2 displays periods of explosiveness suggested by the GSADF test at the State level based on a regression model with an intercept.[4] Tables A1 and A2 present the corresponding date-stamping outcomes and GSADF test statistics for all the 50 States and the District of Columbia.[5] Overall, we seem to find evidence of explosive behaviour in the house price–income ratio for all the States and the District of Columbia except West Virginia. The GSADF statistics for the house price–income ratio of 34 States and the District of Columbia are significant at the 1% level, which indicates strong evidence of explosive periods. The GSADF statistics of 12 States provide evidence of explosive behaviour at the 5% level (i.e. Alaska, Colorado, Iowa, Idaho, Illinois, Indiana, Kansas, Minnesota, New Hampshire, Pennsylvania, South Dakota and Wyoming). The GSADF statistics of price–income ratio for the remaining 3 States exceeds the 10% right-tail critical values (i.e. Louisiana, Oklahoma and Utah). However, it is necessary to bear in mind that the regression model formulation with an intercept could not distinguish between collapse or collapse and recovery episodes and bubbles. Hence, under the model formulation with an intercept, a finding of explosive behaviour in the price–income ratio does not necessarily lead to identification of bubbles. We need to interpret such results with care.

On the other hand, results obtained from the model formulation without an intercept differ from these findings. Figure 3 displays periods of explosiveness suggested by the GSADF test at the State level based on the model formulation without an intercept in the regression model. Figures 7, 9, 11, A2, A4, A6, A8, A10, A12, A14, A16 and A18 compare the

[2]A number of researchers have applied the Chow and Lin's (1971) GLS procedure that can provide the best linear unbiased interpolations in house prices or property markets (see Kenny (1999), Meese and Wallace (2003), Assenmacher and Gerlach (2008), Goodhart and Hofmann (2008) and Zhou (2010)).

[3]A declining price–income ratio implies that the growth of house prices doesn't outpace the growth in personal income. House prices could drop quite substantially in some States due to unfavourable expectations, low employment and economic recession during the 1980–1990s.

[4]We assume that a housing bubble should last at least for 6 months. Thus, a potential bubble episode with shorter period is ignored. Figures 6, 8, 10, A1, A3, A5, A7, A9, A11, A13, A15 and A17 compare the backward SADF statistic with the 95% critical value sequences for the price–income ratio under the regression model with an intercept.

[5]The critical values under the regression model with an intercept are 2.5343 (90%), 2.7960 (95%), 3.4337 (99%). The critical values under the regression model without an intercept are 3.4989 (90%), 3.8319 (95%), 4.5976 (99%).

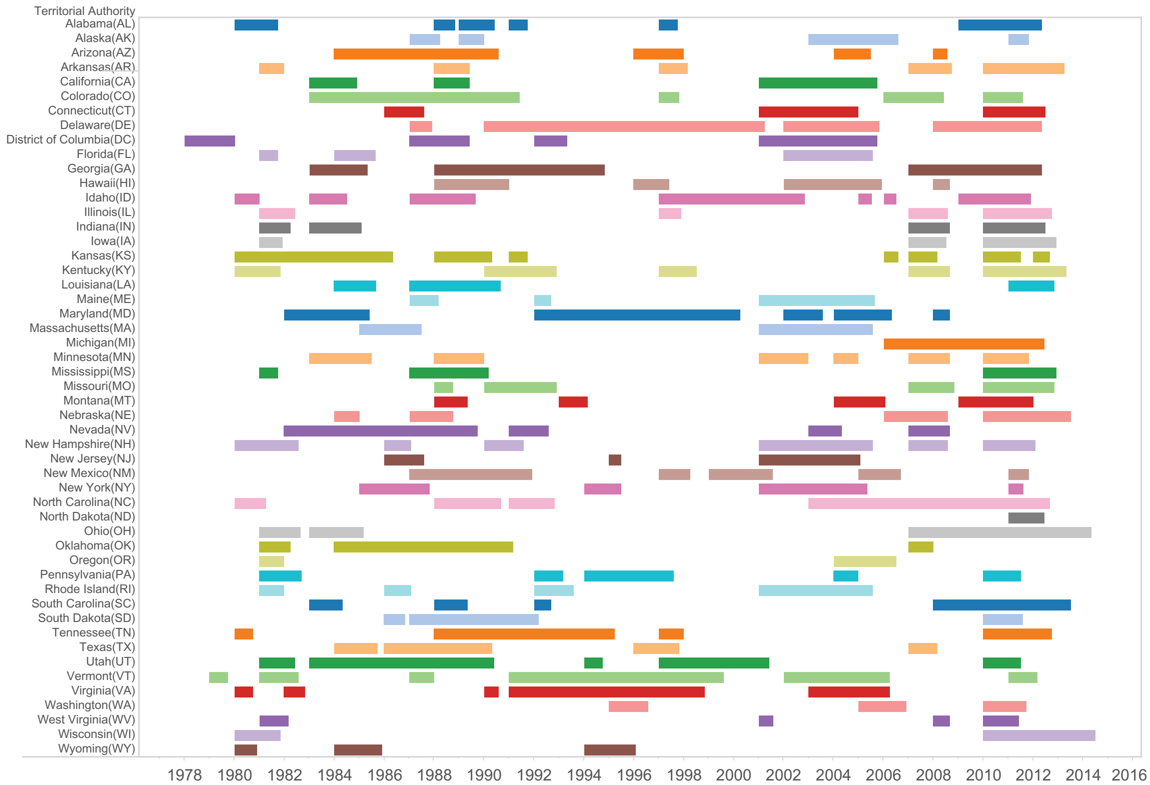

Figure 2. Date-stamping strategies based on the regression model formulation with an intercept.

Figure 3. Date-stamping strategies based on the regression model formulation without an intercept.

backward SADF statistic with the 95% critical value sequences for the price–income ratio under the assumption without an intercept. Tables A3 and A4 present the corresponding date-stamping outcomes and GSADF test statistics for all the States and the District of Columbia. The GSADF statistics of eight States and the District of Columbia are significant at the 1% level, which indicates strong evidence of explosive behaviour (e.g. California, Delaware, Hawaii, Illinois, Maine, New Hampshire, Rhode Island and Virginia). The null hypothesis of no explosive behaviour is rejected at the 5% level for 11 States (e.g. Alaska, Connecticut, Florida, Massachusetts, Maryland, Minnesota, Montana, New Jersey, New York, Pennsylvania and Vermont). The GSADF statistics of house price–income ratio for New Mexico, Oregon, Washington and Wyoming also exceeds the 10% right-tailed critical values. We find no significant evidence of bubbles for the remaining States. Comparing Figures 2 and 3 show that the exclusion of the intercept in the regression model formulation has affected the theory and date-stamping strategy of the PSY approach.

We present results based upon tests of the regression model with and without an intercept. Based on the above results, the intercept term plays a crucial role in identifying the explosive behaviour as it can potentially affect the date-stamping outcomes from the PSY. Our results suggest that we would find the same date-stamping outcomes for with and without intercept in a particular State if the State data does not have any collapse or collapse and recovery events – only the potential bubbles. Examples of collapse episodes can be seen in Figure 10(b,d,e), etc. Similarly, examples of collapse and recovery episodes may be found in Figure 8. When the intercept is included in the regression model, we not only detect the collapse episode or collapse and recovery episode but also the potential bubbles. However, without considering the intercept, the PSY approach detects only the potential bubbles. Special attention should be paid to assessing the evidence of the potential bubbles.

Quite different date-stamping outcomes are obtained from the two model specifications (with or without an intercept) if the price–income ratio of a particular State has a declining trend. Most States with a declining trend exhibit strong evidence of collapse episodes under the model specification with an intercept (e.g. Georgia in Figure A9, Indiana in Figure A5, South Dakota in Figure A7, Tennessee in Figure A13, etc.), which could give the false indication of bubbles. However, when the intercept term is removed in the regression equation, our date-stamping strategies no longer detect the collapse episodes (e.g. Georgia in Figure A10, Indiana in Figure A6, South Dakota in Figure A8, Tennessee in Figure A14, etc.), which suggest no evidence of bubbles.

Our empirical evidence demonstrates the practical importance of regression model specification. Based upon the regression model specification with an intercept, we can identify collapse episode, collapse and recovery episode and the potential bubbles. When the data exhibits a declining trend, the model specification without an intercept seems to provide more promising results as it identifies only the potential bubbles. We do not try to suggest a particular model specification, which always provides the most reliable way in examining the presence of bubbles. One of the take-home messages from our study is that it is useful to try a range of model specifications for assessing evidence of bubbles in right-tailed unit root tests.

Was there a housing bubble in the 1980s?

Existing studies found the existence of a bubble in several States during the 1980s. As shown in Figure 4, based upon the model without an intercept, our empirical results seem to suggest that only 10 States and the District of Columbia experience a housing bubble in the 1980s. The 10 bubbling States include

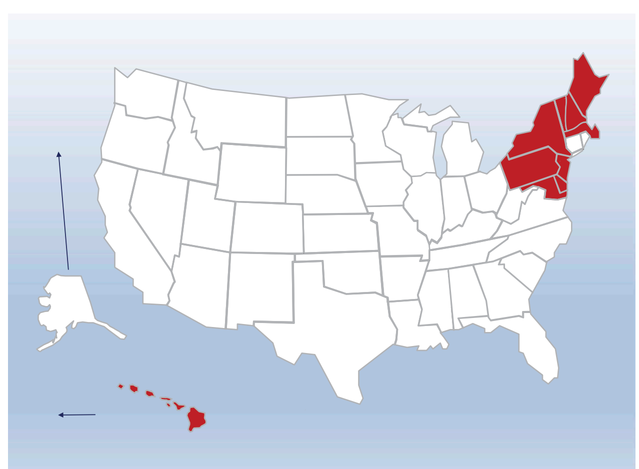

Figure 4. States experiencing a bubble during the 1980s based on the model specification without an intercept.

Delaware, Hawaii, Maine, Massachusetts, Maryland, New Hampshire, New Jersey, New York, Pennsylvania and Vermont. The finding of a housing bubble in the 1980s is in line with several existing studies including Case (1986), Case and Shiller (1988), Case and Shiller (1994), Riddel (1999), Case and Shiller (2003) and Zhou and Sornette (2003). It should be pointed out that California experiences a housing bubble during the 1980s only under the assumption with an intercept.

Was there a housing bubble in the early 2000s?

Several States experienced a dramatic house price boom in the early 2000s followed by a collapse. Glick, Lansing, and Molitor (2015) conclude that the States with the largest house price booms from 2002 to 2006 are Hawaii, Florida, Nevada, California, and Arizona and the States with the smallest house price booms are Mississippi, North Dakota, Oklahoma, South Dakota, Kentucky, Colorado, Idaho, Nebraska, Ohio, Indiana and Michigan. Our results seem to support the view that the States with the largest house price booms are bubbles. The date-stamping strategies of Hawaii in Figure 6(d), Florida in Figure A9, Nevada in Figure 6 (e), California in Figure 6(a) and Arizona in Figure A15 indicate the presence of bubbles between 2001 and 2006 based on the model formulation with an intercept. A similar conclusion can also be drawn based on the model formulation without an intercept except in the case of Arizona and Nevada. On the other hand, the date-stamping strategies of those States with the smallest house price booms suggest no evidence of bubbles between 2001 and 2006 – the exception is Michigan as shown in Figure 9(b).

It should be emphasized that the States with a housing bubble are not limited to California, Florida, Hawaii and Nevada as shown in Figure 5. We also find States with a bubble that grows in the early 2000s and collapses in mid-2000s as presented in Table 1 based on two different model formulations. We can see that the housing bubble is not an isolated phenomenon in particular, 22 or 25 States (depending on the model formulation) and the District of Columbia experience a bubble during this period. Although there are some differences in the origination and collapse dates of the bubbles based on the two model formulations, our results provide similar date-stamping outcomes and also confirm the existence of a bubble in these States. According to Martin (2011), the

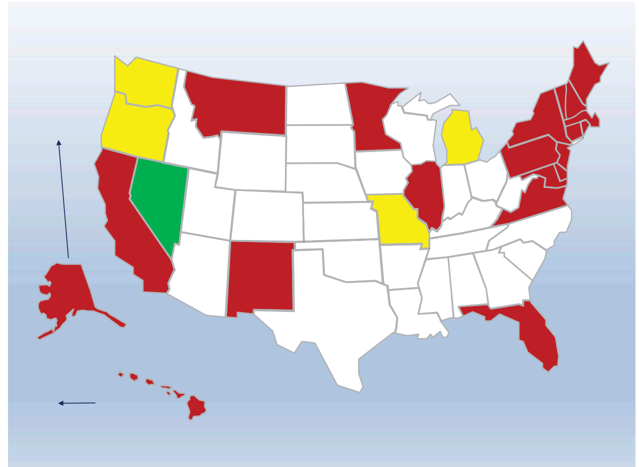

Figure 5. States experience a bubble during the early 2000s based on the two model specifications. A State experiencing a bubble under both models is coloured in red. A State experiencing a bubble only under the assumption without an intercept is coloured in yellow while a State experiencing a bubble only under the assumption with an intercept is coloured in green (For interpretation of the references to colour in this figure legend, the reader is referred to the web version of this article.).

Table 1. States with a bubble during the early 2000s boom.

State	Bubble episode(s) under the model with an intercept	Bubble episode(s) under the model without an intercept
Alaska	2003M06–2006M09	2003M10–2006M12
California	2001M02–2005M11	2001M03–2006M08
Connecticut	2001M10–2005M10	2002M01–2006M08
District of Columbia	2001M02–2005M11	2000M12–2005M12
Delaware	2002M12–2006M10	2003M08–2007M09
Florida	2002M04–2005M11	2003M03–2006M05
Hawaii	2002M07–2006M06	2001M10–2006M07
Illinois	2002M04–2005M11	2002M10–2006M11
Massachusetts	2001M01–2005M08	2000M05–2005M12
Maryland	2002M05–2003M12 2004M02–2006M06	2003M03–2006M11
Maine	2001M04–2005M12	2001M08–2006M11
Michigan		2001M06–2004M11
Minnesota	2001M11–2003M11	2001M04–2006M09
Missouri		2003M05–2003M09 2004M05–2005M10
Montana	2004M11–2006M12	2004M06–2007M11
Nevada	2003M10–2005M02	
New Hampshire	2001M03–2005M10	2001M11–2006M07
New Jersey	2001M11–2005M12	2002M10–2006M09
New York	2001M06–2005M11	2001M06–2006M10
New Mexico	2005M05–2007M01	2004M10–2007M07
Oregon		2004M02–2007M11
Pennsylvania	2004M11–2005M11	2003M05–2006M11
Rhode Island	2001M04–2005M11	2002M03–2006M01
Virginia	2003M03–2006M06	2003M01–2006M11
Vermont	2002M02–2006M05	2002M04–2006M12
Washington		2003M05–2007M08

States that experienced a bubble during this period include California, District of Columbia, Florida, Hawaii, Massachusetts, Nevada, Pennsylvania and Rhode Island, which are well aligned with our results. Crucially, our results allow us to draw an important conclusion – the bubble of the 2000s is more widespread than the 1980s. This phenomenon is clear via close inspection of Figure 3. Our results also seem to suggest that the housing bubble in the early 2000s is not a national bubble, as many States experience no bubble activity during this period. This conclusion is consistent with the views of Greenspan in 2005. The States that experience a bubble are mainly located in Northeast or West with the exception of Florida, New Mexico, Montana, Michigan, Missouri, Minnesota and Illinois. Wheelock (2006) also concluded that housing booms and busts in the 1980s and early 2000s are different in terms of their magnitude and coverage where the early 2000s boom appeared more widely spread out than that of the 1980s.

California (CA)

A number of studies focus on the Californian housing market (e.g. Case and Shiller (1994), Gabriel, Mattey, and Wascher (1999),[6] Riddel (1999) and Riddel (2011).[7]). Our date-stamping strategy results for the Californian price–income ratio are presented in Figure 6(a), where the price–income ratio reaches its lowest point in 1984/1985 and starts to climb to a peak in 1989. After 1989/1990, the ratio declines sharply then reaches another peak in 2005/2006. We find several episodes from the price–income ratio for California: 1983M07–1985M06, 1988M05–1989M10 and 2001M02–2005M11. Case (1986) argues that there is a house price boom in California in 1976–1980; however, our results indicate that such a 'housing boom' was not a bubble.

Based on our analysis, there are two episodes in the 1980s (1983M07–1985M06, 1988M05–1989M10). The first period between 1983M07 and 1985M06 is an example of a collapse episode. The papers by Case (1994), Case and Shiller (1994), Riddel (1999), Zhou and Sornette (2003) and Gupta and Miller (2012) offer some support for our findings. Case and Shiller (1988) were the first to uncover the role of high expectations in driving up the California boom in late 1980s using a questionnaire survey. First, Case and Shiller (1988), surveyed a sample of 2000 households who bought homes in 1988 in Los Angeles, San Francisco, Boston and Milwaukee. The Los Angeles and San Francisco housing markets were chosen to represent two 'boom markets' and Boston was selected as a 'post-boom market'. Milwaukee was treated as a control market to reflect a 'normal' housing market. A key finding from Case and Shiller (1988) suggests that three housing markets (Los Angeles, San Francisco and Boston) have gone through a house price bubble period. A house price boom in California over the period 1987–1989 is also identified in Case (1994), which coincides with our second episode 1988M05–1989M10.[8] Second, Case and Shiller (1994) compared two house price boom/burst cycles in Boston and Los Angeles and they focused on the period post-1983 in Boston and the period post-1985 in Los Angeles. The prices in Los Angeles increased more than 100% between 1985 and 1989. They concluded that these two house price booms cannot be fully explained by economic fundamentals. The run-up in prices seem to have be driven by 'speculation' as most home buyers paid higher prices for properties and aimed at future capital gains. Moreover, Riddel (1999) concluded that there was a speculative housing bubble from late 1987 to mid-1990 in the Santa Barbara County of California. Zhou and Sornette (2003) also found evidence of a housing bubble that originated around 1984 and burst in 1989 in the Californian market based on the characteristic of a bubble defined as 'a super-exponential growth' phase. They draw the same conclusion for Los Angeles and San Francisco as these two cities led the Californian market. This finding is consistent with our analysis. Lastly, Southern California experienced a run-up and subsequent fall in house prices during the late 1980s and early 1990s (Gupta and Miller 2012).[9] This is due to the fact that California experienced considerable economic growth during 1983–1989 and suffered a major decline in economic activity during the 1990–1991 recession.

[6]Gabriel, Mattey, and Wascher (1999) explored the housing price patterns in California's two largest Metropolitan areas (Los Angeles and San Francisco) prior to 2000.

[7]Riddel (2011) estimated an error correction model that spanned 1978Q2 to 2008Q1 using quarterly housing price data for Las Vegas and Los Angeles with both national and regional economic variables. Riddel (2011) provided support for the contagion hypothesis that income and price in Los Angeles contributed to the run-up house values in Las Vegas from 2002 to 2006.

[8]Case (1994) reviewed the house prices in the US since the 1950s at national and regional level and discussed the causes of house price behaviours across regions.

[9]Gupta and Miller (2012) explored the cointegration relationships between house prices in eight Southern California metropolitan statistical areas (MSAs).

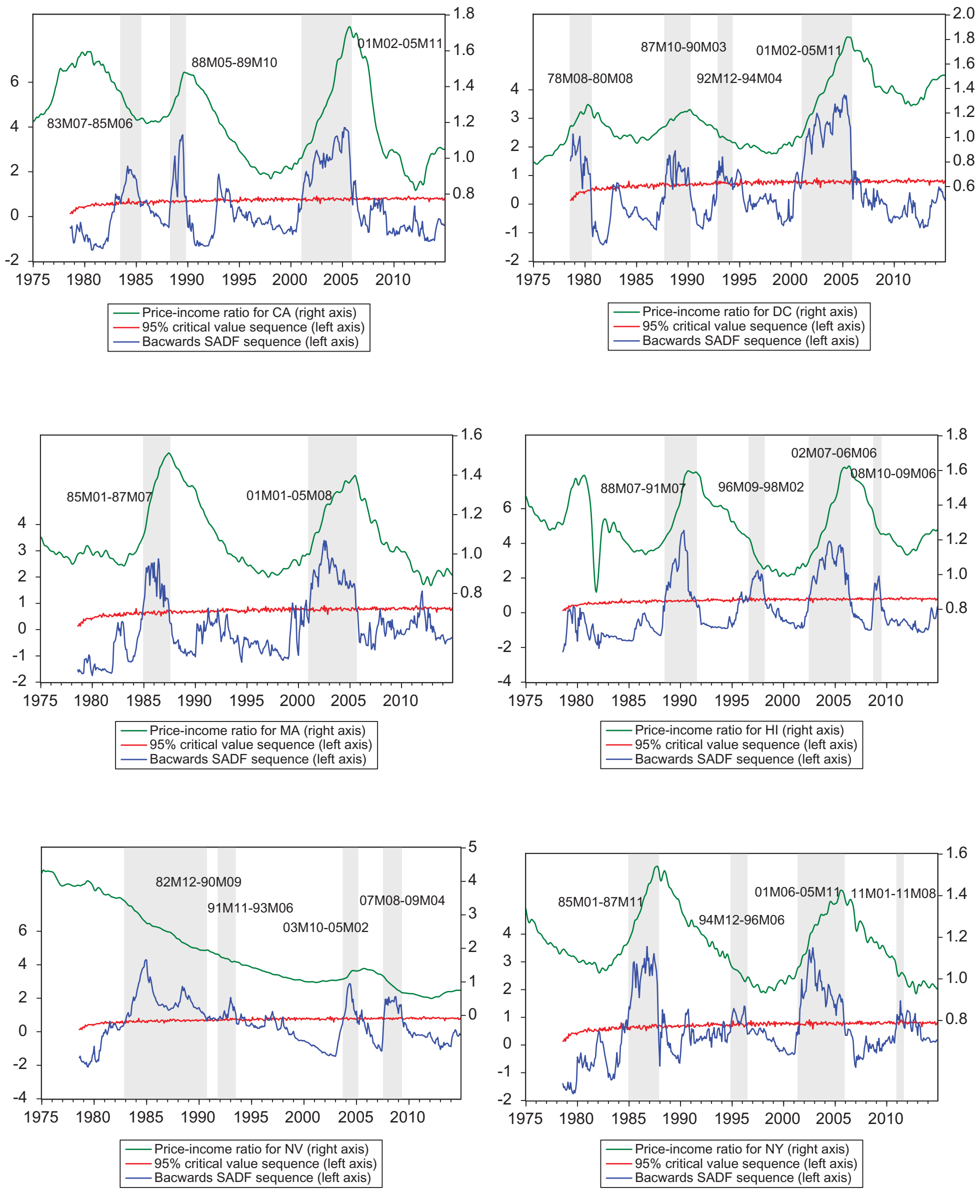

Figure 6. Date-stamping strategy of price–income ratio for several key States based on the regression model formulation with an intercept. (a) California (CA); (b) District of Columbia (DC); (c) Massachusetts (MA); (d) Hawaii (HI) (e) Nevada (NV); and (f) New York (NY).

House prices in California accelerated again in the early 2000s then declined from their peak in 2005/2006. Our results confirm this period as a bubble. The finding of a bubble during this period coincides with several existing studies in the literature (e.g. Case and Shiller 2003; Gupta and Miller 2012). For example, Case and Shiller (2003) replicated a questionnaire survey in 2003 for the same markets (e.g. Los Angeles, San Francisco, Boston and Milwaukee). Unlike the first survey in 1988, these three markets were in booms while Milwaukee was a control city. The survey was conducted among 2000 persons who had bought houses between March and August 2002. They concluded the characteristics of bubbles in these four markets were strong in 2003, but not as strong as in 1988. Results obtained from Case and Shiller (2003) also seem to suggest the existence of bubbles. Moreover, house prices rose dramatically in Southern California MSAs in the early 2000s, peaking in 2005 or 2006 depending on the MSA (Gupta and Miller 2012). It is interesting that our identified episode (2001M02–2005M11) is in line with the period of run-up and decline in house values in California.

We obtain quite different results, however, using the regression model specification without an intercept. According to Figure 7(a), we detect an early episode in 1979, which is not detected earlier. This finding is consistent with Case (1986), as discussed earlier, who concluded a housing boom at that time in California. No explosive behaviour is found in the 1980s as the collapse-only episode between 1983M07 and 1985M06 is not identified. This result does not support the view that California experienced a bubble in the 1980s, which is a key difference between the two models. However, the housing boom in the 2000s is shown to be a bubble.

District of Columbia (DC)

Based upon the assumption 'with an intercept', the data-stamping strategy for the District of Columbia is shown in Figure 6(b). The empirical results suggest that the District of Columbia experiences three house price bubbles in 1978M08–1980M08, 1987M10–1990M03 and 2001M02–2005M11. These occurrences are quite unusual as, based upon our results, few other areas experienced a housing bubble between the late 1970s and early 1980s. However, the District of Columbia experiences a housing bubble and the price–income ratio reaches a peak in 1979/

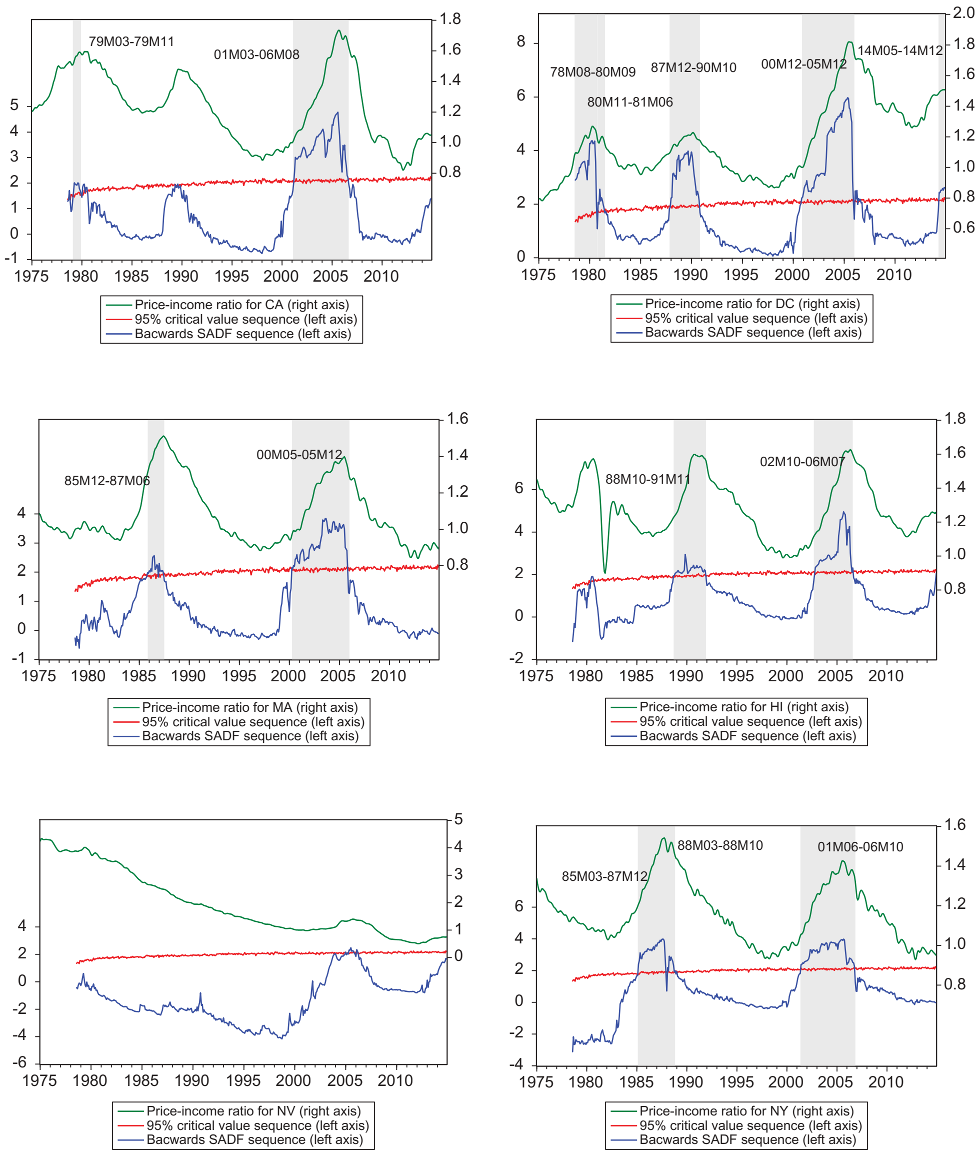

Figure 7. Date-stamping strategy of price–income ratio for several key States based on the regression model formulation without an intercept. (a) California (CA); (b) District of Columbia (DC); (c) Massachusetts (MA); (d) Hawaii (HI) (e) Nevada (NV); and (f) New York (NY).

1980. A collapse episode is also identified during the period 1992M12–1994M04. The recent run-up in housing prices during the early-mid 2000s is shown to be a bubble.

Moreover, as can be seen in Figure 7(b), we obtain similar date-stamping outcomes when the intercept is removed from the regression model, where we are able to identify the following bubbles: 1978M08–1980M09, 1980M11–1981M06, 1987M12–1990M10, 2000M12–2005M12 and 2014M05–2014M12. The most recent episode (2014M05–2014M12) suggests that the District of Columbia is experiencing a new housing bubble, which is of particular interest. Our results, based on tests of the regression model without an intercept, indicate the District of Columbia is the only area experiencing a housing bubble post-2010. This is clearly shown in Figure 3.

Massachusetts (MA)

Based on the model formulation with an intercept, we find significant evidence of bubbles in the price–income ratio for the period 1985M01-–1987M07 and 2001M02–2005M08 presented in Figure 6(c). The GSADF statistics is 3.3926, which is significant at the 1% level. A series of papers by Case (1986, 1994) and Case and Shiller (2003) enable us to understand house price dynamics in Boston during the 1980s. House prices rose rapidly in Boston in 1984 with house prices in the Boston Metropolitan area increasing by 39% in 1985 and more than 140% in 1988 (Case and Shiller 2003). House prices in Boston remained unchanged at or near their peak for almost 3 years from 1987Q2 to 1990Q1. If the price increases in 1994/1995 were maintained, house prices would double approximately every 3 years in Boston (Case 1986). Most importantly, Case (1986) concluded that Boston had experienced a housing price bubble as market fundamentals, (e.g. population growth, employment growth, increasing income, mortgage rates, construction costs) did not fully explain the rapid increase in house values. Case (1994) also described several house price booms in the US including the boom in Boston 1983–1987. Therefore, our findings for the first explosive period (1985M01–1987M07) coincide with the historical experiences of house price boom/bubble in Case (1986, 1994) and Case and Shiller (2003). The run-up in house values of Massachusetts in the early 2000s is a bubble, which is consistent with the conclusion drawn by Case and Shiller (2003).

We obtain similar date-stamping outcomes when the intercept term is removed from the regression model specification. As shown in Figure 7(c), two episodes (1985M12–1987M06 and 2000M05–2005M12) are identified. Results based on the two model formulations seem to suggest that Massachusetts experienced a bubble in the 1980s and early 2000s, respectively.

Hawaii (HI)

As illustrated by Figure 6(d), Hawaii experienced a bubble period between July 1988 and July 1991. As suggested in a series of papers by Miller, Sklarz, and Ordway (1988), Wheelock (2006), Krainer and Wilcox (2011), Krainer and Wilcox (2013), there was rapid growth in Hawaii's real estate market from the late 1980s where the appreciation of house prices was driven by Japanese investors. The house values in Hawaii rose dramatically during the years of the Japanese economy boom in the 1980s and house prices dropped significantly in the early 1990s when the Japanese asset price bubble collapsed (Krainer and Wilcox 2013). Our results support the historical house price dynamics in Hawaii during the late 1980s and early 1990s.

When the intercept term is removed from the regression model, the PSY approach no longer detects the collapse episodes (i.e. 1996M09–1998M02 and 2008M10–2009M06). The exclusion of the intercept term has affected the asymptotic theory and date-stamping outcomes of the PSY approach. Overall, based on the two models, our results seem to indicate that Hawaii experienced a bubble during the late 1980s and early 1990s. This finding coincides with the house price boom/bust in Hawaii. Moreover, the housing boom in the early 2000s in Hawaii is also a bubble.

Nevada (NV)

The null of no explosive behaviour for Nevada's price–income ratio is rejected at the 5% level ($4.2925 > 2.7960$). As illustrated by Figure 6(e), we find several episodes in the price–income ratio series (e.g. 1982M12–1990M09, 1991M11–1993M06, 2003M10–2005M02, 2007M08–2009M04). Zhou and Sornette (2008) analysed the real estate market of Las Vegas between June 1983 and March 2005 and argue for the existence of a housing bubble during 2003 to mid-2004.[10] This finding partially supports our analysis for the episode during 2003M10–2005M02. Riddel (2011) presented the view that contagious price and income growth from the Los Angeles market contributed to the bubble formed in Las Vegas house prices during 2002–2006.

However, a close inspection of the price–income ratio series in Figure 6(e) shows that the general trend of the price–income ratio is declining. Several identified episodes are collapse episodes. When the intercept is removed from the regression model, we find no explosive behaviour as the GSADF statistics is much lower than the 10% critical values (2.4837 < 3.4989). Thus, there is no evidence to indicate the presence of bubbles in Nevada for the whole period as suggested in Figure 7(e).

New York (NY)

The null of no explosive behaviour is rejected at the 1% level for New York (3.5519 > 2.5343). As shown in Figure 6(f), we identify several episodes: 1984M09–1987M12, 1994M12–1996M06, 2001M07–2005M12 and 2011M01–2011M07. Case (1986) also discussed a house price boom in New York during the 1980s, where the median sale prices of existing single-family homes rose by 30% in New York in 1985. The rapid increase in house prices in Boston and New York in the mid-1980s doesn't suggest a national house market boom as many cities faced a decline in nominal values Case 1986. A house price boom in New York in 1983–1987 is identified in Case (1994). A finding of a bubble period 1984M09–1987M12 from our analysis seems to be in line with house price boom in New York during the mid-1980s. Wheelock (2006) also support a housing boom in New York between 1985Q1 and 1987Q3 and a decline in house price followed.[11] Moreover, based on our results, the rapid appreciation of house price–income ratio between 2001M07 and 2005M12 is a bubble.

We no longer detect the collapse episodes of 1994M12–1996M06 and 2011M01–2011M07 in Figure 7(f) when the intercept is removed from the regression model specification. We obtain similar date-stamping outcomes under the regression model without an intercept except for the omission of the two collapse episodes.

Farm and 'rust belt' states: Iowa (IA), Michigan (MI), Wisconsin (WI) and West Virginia (WV)

The date-stamping strategy for the price–income ratio of the farm and the so-called 'Rust Belt' States is presented in Figure 8. The general trend for the price–income ratio in these figures is downward sloping. The null hypothesis of no explosive behaviour in the price–income ratio is rejected at the 1%, 5% and 5% level for Iowa, Michigan and Wisconsin, respectively. The GSADF statistics for the house price–income ratio in West Virginia is 1.8273, which is below the 10% level significance.

The house price busts occurred during 1980–1982 in these four States (Wheelock 2006). Of particular interest is whether these house price busts are bubbles. As can be seen in Figure 8, the price–income ratio reaches a peak or maintains a high level in 1980 for Iowa (IA), Michigan (MI), West Virginia (WV) and Wisconsin (WI). The income of these States relied heavily on older manufacturing industries (e.g. automobiles and steel) where there was a large decline during the early 1980s recession (Wheelock 2006).[12] Although these States do not experience a high growth in house prices, the PSY approach identifies the collapse and recovery episodes in the early 1980s based (e.g. 1981M01–1981M12 for Iowa, 1982M06–1982M11 for Michigan, 1981M10–1982M12 for West Virginia and 1980M10–1982M08 for Wisconsin).

As shown in Figure 9, quite different results are obtained from the GSADF test under the assumption of no intercept. The null hypothesis of no explosive behaviour cannot be rejected at the 10% level and we find no significant evidence of bubbles in these four States. We also find no explosive periods during

[10] Zhou and Sornette (2008) defined a bubble as a price acceleration faster exponential. This definition is the same as Zhou and Sornette (2006).
[11] Wheelock (2006) summarized that US States experienced 20 house price booms between 1980 and 1999.
[12] During 1980–1982, Iowa, Wisconsin, West Virginia and Michigan ranked 42nd, 44th, 45th, and 50th, respectively, among all States in real personal income growth, and 45th, 40th, 48th and 50th in employment growth (Wheelock 2006).

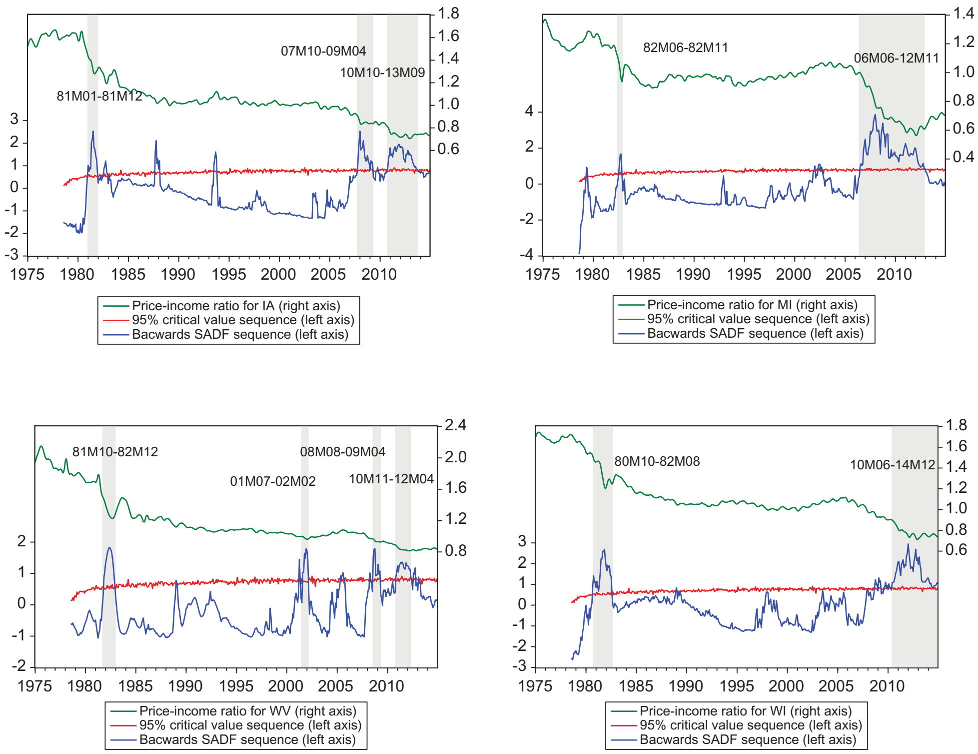

Figure 8. Date-stamping strategy of price–income ratio for farm and the 'Rust Belt' States based on the regression model formulation with an intercept. (a) Iowa (IA); (b) Michigan (MI); (c) West Virginia (WV); (d) Wisconsin (WI).

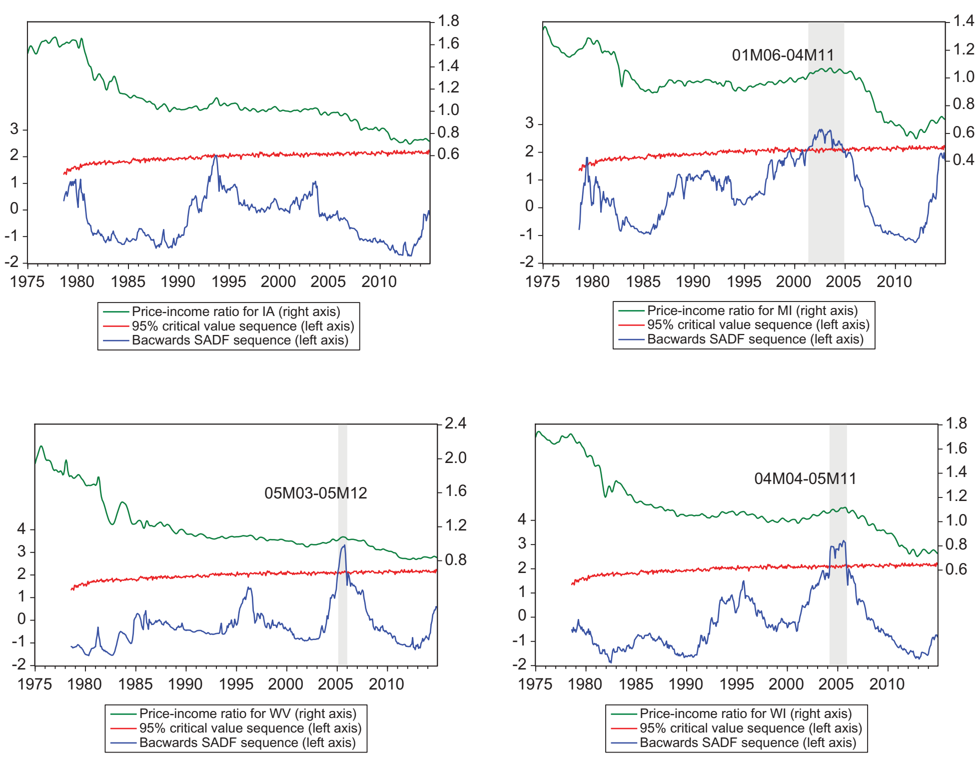

Figure 9. Date-stamping strategy of price–income ratio for farm and the 'Rust Belt' States based on the regression model formulation without an intercept. (a) Iowa (IA); (b) Michigan (MI); (c) West Virginia (WV); and (d) Wisconsin (WI).

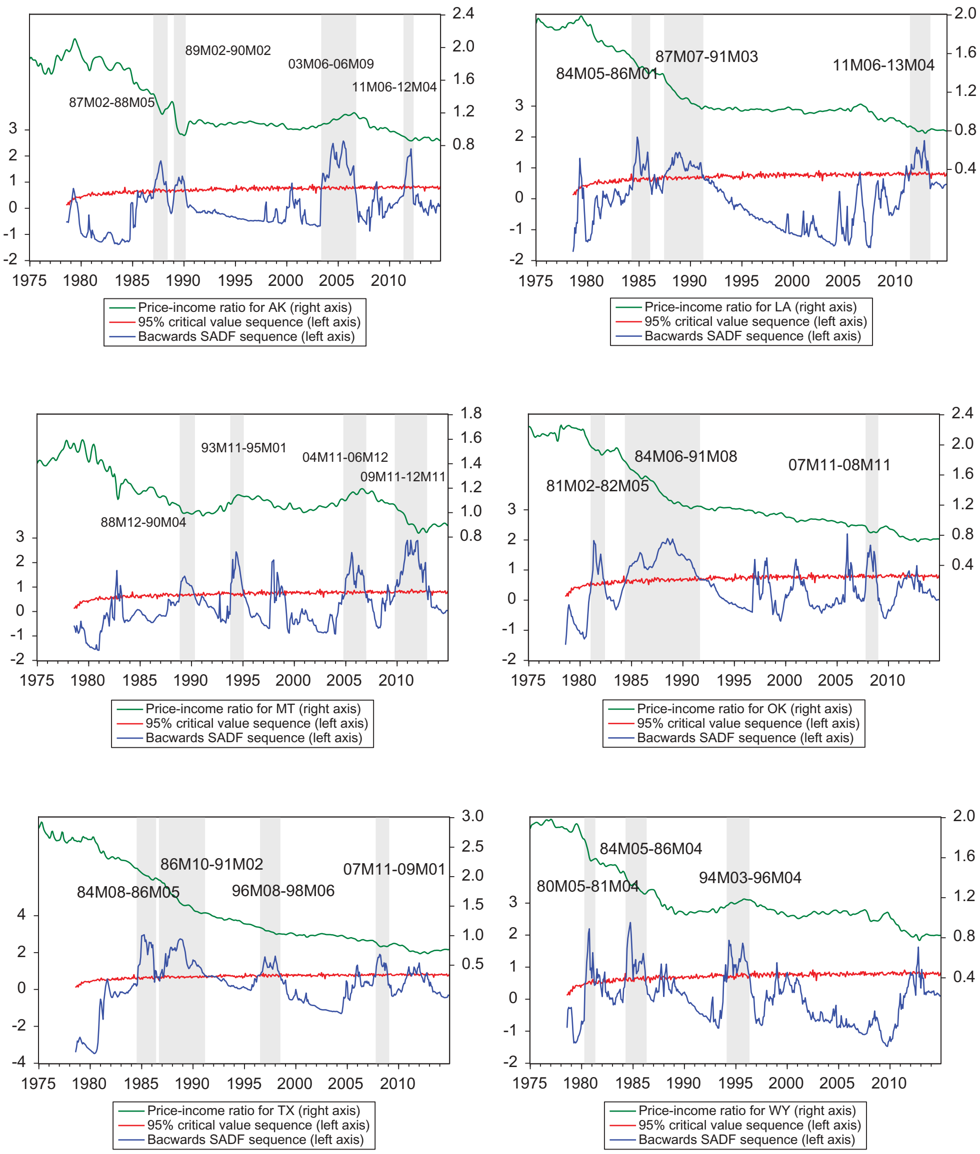

Figure 10. Date-stamping strategy of price–income ratio for energy-producing States based on the regression model formulation with an intercept. (a) Alaska (AK); (b) Louisiana (LA); (c) Montana (MT); (d) Oklahoma (OK); (e) Texas (TX); and (f) Wyoming (WY).

1980–1982 in these four States, which indicates that the house price busts are not bubbles.

Energy-producing states: Alaska (AK), Louisiana (LA), Montana (MT), Oklahoma (OK), Texas (TX) and Wyoming (WY)

The oil prices rise sharply during the 1970s energy crisis and it reached a peak in 1980 at more than $35 per barrel before plunging to less than $10 per barrel in 1986. Wheelock (2006) argues that the house price busts in these energy-producing States was associated with a sharp decline in energy prices. We therefore aim to investigate whether these house price busts during the early-mid 1980s are bubbles. The trend for price–income ratio of these six energy-producing States is downward sloping (see Figure 10). The null hypothesis of no bubbles is rejected for all six States; however, we find significant evidence of 'collapse' episodes and 'collapse and recovery' episodes in Figure 10.

Based on the regression model without an intercept, the null hypothesis of no explosive periods can be rejected for Alaska, Montana and Wyoming at the 5%, 5% and 10% significance level, whereas the null hypothesis cannot be rejected for Louisiana and

Texas. Perhaps more importantly, we find insufficient evidence to support that house price busts associated with the sharp decline in energy prices during the early-mid 1980s are bubbles in Figure 11.

VI. Conclusion

In this article, we investigate the presence of bubbles in the US housing market at the State level based on price–income ratio data, January 1975–December 2014. The recently developed right-tailed unit root test of Phillips, Shi, and Yu (2015,) is adopted in our study. Our results are summarized as follows. First, we find the presence of a bubble in several States in the 1980s (i.e. California, Hawaii, Massachusetts and New York), which is consistent with some existing studies that investigate housing bubbles or housing booms and busts (i.e. Case (1986), Case and Shiller (1988, 1994, 2003), Riddel (1999) and Zhou and Sornette (2003)). Our article completes analysis in this area by formally testing for bubbles (rather than simply analysing graphs or house price indicators) in all US States (and DC) during the 1980s, providing empirical evidence of housing bubbles in several States. Second, we identify the existence of a housing bubble that originates in the early 2000s and

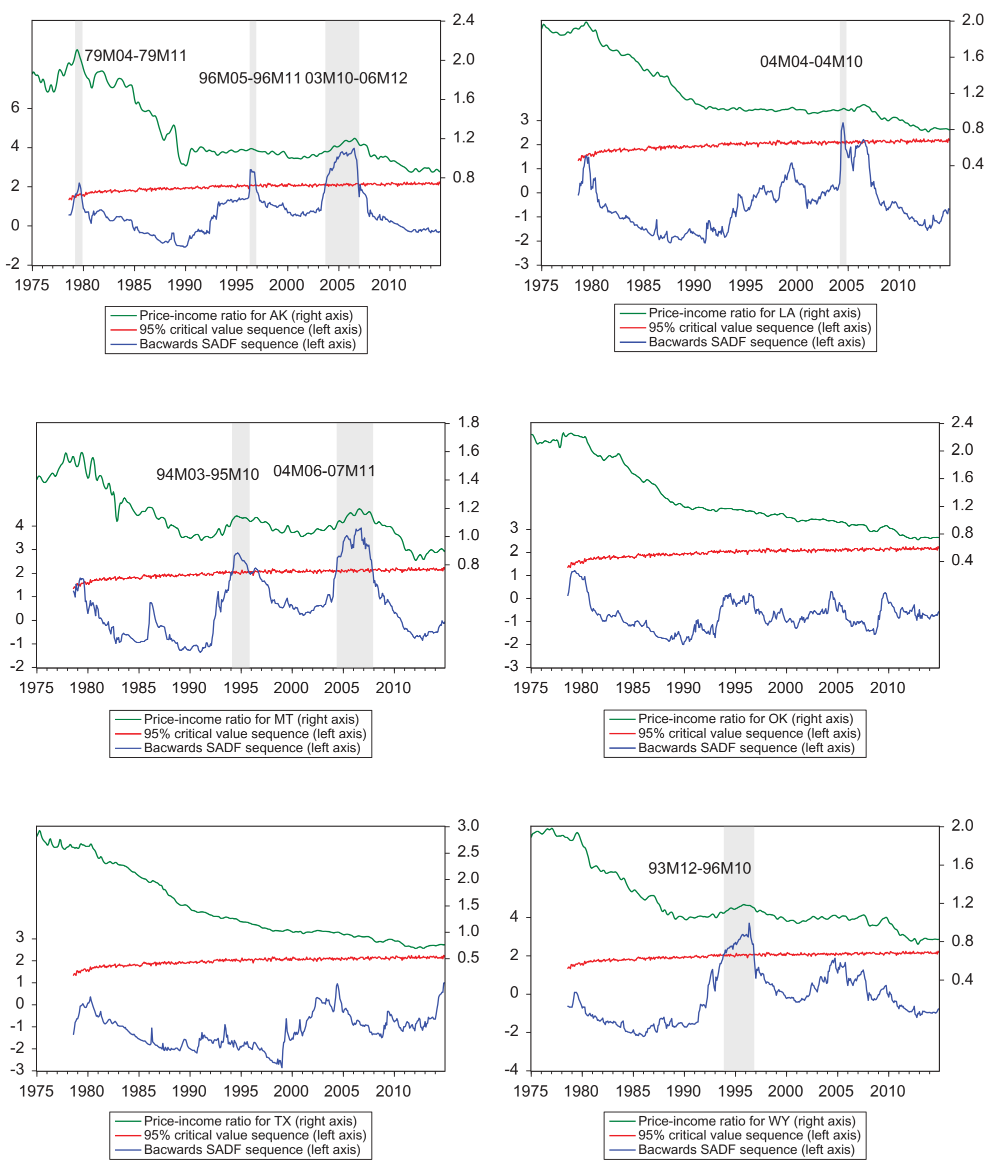

Figure 11. Date-stamping strategy of price–income ratio for energy-producing States based on the regression model formulation without an intercept. (a) Alaska (AK); (b) Louisiana (LA); (c) Montana (MT); (d) Oklahoma (OK); (e) Texas (TX); and (f) Wyoming (WY).

collapses in the mid-2000s in more than 20 States and the District of Columbia. Our results also suggest that the housing bubble in the early 2000s is more widespread than the 1980s. Moreover, the rapid rise in house values during the early 2000s does not seem to be a nationwide bubble which is in agreement with the talk given by Alan Greenspan in 2005. Lastly, the exclusion of an intercept in the regression model has shown to affect the asymptotic theory and date-stamping outcomes for the PSY approach. When the intercept is removed from the regression model specification, we no longer detect the collapse or recovery episode only the potential bubble. We do not try to suggest a particular model specification, which always provides the best date-stamping outcomes. One of the take-home messages from our study is there is a need to consider different regression model specifications for assessing evidence of bubbles in the right tailed unit root tests. This represents an important example of how care is required when using the PSY approach.

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Appendix

Table A1. Testing for explosiveness in the US house price–income ratio at the State level based on regression model formulation with an intercept (1).

Territorial authority	GSADF statistic	Episode(s)
Alaska (AK)	2.5753**	87M02–88M05, 89M02–90M02, 03M06–06M09, 11M06–12M04
Alabama (AL)	2.8677***	80M09–82M06, 88M08–89M06, 89M10–91M03, 91M10–92M07, 97M10–98M07, 09M12–13M04
Arkansas (AR)	3.4767***	81M05–82M05, 88M11–90M04, 97M07–98M08, 07M08–09M05, 10M02–13M05
Arizona (AZ)	4.0760***	84M05–90M12, 96M11–98M11, 04M05–05M11, 08M10–09M05
California (CA)	3.9756***	83M07–85M06, 88M05–89M10, 01M02–05M11
Colorado (CO)	2.5833**	83M08–92M01, 97M10–98M08, 06M10–09M03, 10M10–12M05
Connecticut (CT)	2.8532***	86M01–87M08, 01M10–05M10, 10M10–13M04
District of Columbia (DC)	3.8213***	78M08–80M08, 87M10–90M03, 92M12–94M04, 01M02–05M11
Delaware (DE)	3.8356***	87M04–88M03, 90M10–02M01, 02M12–06M10, 08M10–13M02
Florida (FL)	3.7602***	81M02–81M11, 84M09–86M05, 02M04–05M11
Georgia (GA)	4.3743***	83M02–85M06, 88M07–95M05, 07M08–12M12
Hawaii (HI)	4.7460***	88M07–91M07, 96M09–98M02, 02M07–06M06, 08M10–09M06
Iowa (IA)	2.5337**	81M01–81M12, 07M10–09M04, 10M10–13M09
Idaho (ID)	2.5736**	80M09–81M09, 83M12–85M06, 87M07–90M03, 97M05–03M03, 05M05–05M11, 06M02–06M08, 09M06–12M04
Illinois (IL)	2.3791**	81M09–83M02, 97M07–98M06, 07M09–09M04, 10M06–13M03
Indiana (IN)	2.2692**	81M11–83M02, 83M06–85M07, 07M08–09M04, 10M10–13M04
Kansas (KS)	2.6143**	80M10–87M02, 88M08–89M08, 89M10–91M02, 91M09–92M06, 06M10–07M04, 07M09–08M11, 10M10–12M04, 12M08–13M04
Kentucky (KY)	3.0950***	80M10–82M08, 90M01–92M12, 97M08–99M02, 07M08–09M04, 10M03–13M07
Louisiana (LA)	1.9972*	84M05–86M01, 87M07–91M03, 11M06–13M04
Massachusetts (MA)	3.3926***	85M01–87M07, 01M01–05M08
Maryland (MD)	3.3083***	82M02–83M04, 83M10–86M03, 92M11–01M02, 02M05–03M12, 04M02–06M06, 08M08–09M04
Maine (ME)	3.6362***	87M01–88M03, 92M12–93M08, 01M04–05M12
Michigan (MI)	3.8590***	06M06–12M11
Minnesota (MN)	2.2989**	83M10–86M04, 88M12–90M12, 01M11–03M11, 04M11–05M11, 07M08–09M04, 10M06–12M04
Missouri (MO)	3.6546***	88M10–89M07, 90M04–93M03, 07M06–09M04, 10M06–13M04
Mississippi (MS)	3.7950***	81M06–82M03, 87M08–90M10, 10M04–13M03

** indicates significance at the 5% level.
*** indicates significance at the 1% level.
* indicates significance at the 10% level.

Table A2. Testing for explosiveness in the US house price–income ratio at the State level based on regression model formulation with an intercept (2).

Territorial authority	GSADF statistic	Episode(s)
Montana (MT)	2.9110***	88M12–90M04, 93M11–95M01, 04M11–06M12, 09M11–12M11
North Carolina (NC)	3.4271***	80M10–82M01, 88M01–90M09, 91M07–93M05, 09M03–13M05
North Dakota (ND)	3.1273***	11M06–12M11
Nebraska (NE)	4.3557***	84M06–85M06, 87M11–89M08, 06M12–09M07, 10M06–13M12
New Hampshire (NH)	2.7671**	80M09–83M04, 86M03–87M04, 90M12–92M07, 01M03–05M10, 07M09–09M04, 10M010–12M11
New Jersey (NJ)	2.8940***	86M01–87M08, 95M11–96M05, 01M11–05M12
New Mexico (NM)	3.0967***	87M09–92M08, 97M10–99M01, 99M09–02M04, 05M05–07M01, 11M06–12M04
Nevada (NV)	4.2925***	82M12–90M09, 91M11–93M06, 03M10–05M02, 07M08–09M04
New York (NY)	3.5519***	85M01–87M11, 94M12–96M06, 01M06–05M11, 11M01–11M08
Ohio (OH)	2.8279***	81M02–82M10, 83M10–85M12, 07M08–14M12
Oklahoma (OK)	2.2009*	81M02–82M05, 84M06–91M08, 07M11–08M11
Oregon (OR)	3.1929***	81M02–82M02, 04M12–07M06
Pennsylvania (PA)	2.5222**	81M06–83M02, 92M05–93M07, 94M11–98M06, 04M11–05M11, 10M10–12M04
Rhode Island (RI)	2.8820***	81M10–82M10, 86M06–87M07, 92M02–93M09, 01M04–05M11
South Carolina (SC)	3.5162***	83M11–85M03, 88M02–89M06, 92M12–93M08, 08M09–14M03
South Dakota (SD)	2.7747**	86M04–87M02, 87M07–92M09, 10M10–12M05
Tennessee (TN)	3.5187***	80M07–81M04, 88M01–95M04, 97M11–98M11, 10M05–13M02
Texas (TX)	2.9688***	84M08–86M05, 86M10–91M02, 96M08–98M06, 07M11–09M01
Utah (UT)	2.2061*	81M05–82M10, 83M08–91M01, 94M01–94M10, 97M10–02M02, 10M09–12M03
Virginia (VA)	3.6578***	80M11–81M08, 82M01–82M11, 90M10–91M05, 91M08–99M06, 03M03–06M06
Vermont (VT)	4.5995***	79M06–80M03, 81M02–82M09, 87M11–88M11, 91M08–00M03, 02M02–06M05, 11M03–12M05
Washington (WA)	2.7979***	95M10–97M05, 05M04–07M03, 10M09–12M06
Wisconsin (WI)	2.9467***	80M10–82M08, 10M06–14M12
West Virginia (WV)	1.8273	81M10–82M12, 01M07–02M02, 08M08–09M04, 10M11–12M04
Wyoming (WY)	2.3944**	80M05–81M04, 84M05–86M04, 94M03–96M04

*** indicates significance at the 1% level.
** indicates significance at the 5% level.
* indicates significance at the 10% level.

Table A3. Testing for explosiveness in the US house price–income ratio at the State level based on regression model formulation without an intercept (1).

Territorial authority	GSADF statistic	Episode(s)
Alaska (AK)	3.9690**	79M04–79M11, 96M05–96M11, 03M10–06M12
Alabama (AL)	0.8321	
Arkansas (AR)	0.6238	
Arizona (AZ)	2.7832	04M03–04M08
California (CA)	4.7830***	79M03–79M11, 01M03–06M08
Colorado (CO)	2.3487	79M03–79M08
Connecticut (CT)	4.2990**	02M01–06M08
District of Columbia (DC)	5.9705***	78M08–80M09, 80M11–81M06, 87M12–90M10, 00M12–05M12, 14M05–14M12
Delaware (DE)	5.1558***	87M08–90M06, 03M08–07M09
Florida (FL)	4.2377**	03M03–06M05
Georgia (GA)	1.3040	
Hawaii (HI)	4.9623***	88M01–91M11, 02M10–06M07
Iowa (IA)	2.0512	
Idaho (ID)	2.3192	
Illinois (IL)	4.2508***	02M10–06M11
Indiana (IN)	0.2466	
Kansas (KS)	1.6406	
Kentucky (KY)	0.8426	
Louisiana (LA)	2.9086	04M04–04M10
Massachusetts (MA)	3.8427**	85M12–87M06, 00M05–05M12
Maryland (MD)	4.0003**	89M04–90M04, 03M03–06M11
Maine (ME)	5.1095***	86M05–88M10, 01M08–06M11
Michigan (MI)	2.8259	01M06–04M11
Minnesota (MN)	4.1407**	01M04–06M09
Missouri (MO)	2.8779	03M05–03M09, 04M05–05M10
Mississippi (MS)	0.8537	

** indicates significance at the 5% level.
*** indicates significance at the 1% level.

Table A4. Testing for explosiveness in the US house price–income ratio at the State level based on regression model formulation without an intercept (2).

Territorial authority	GSADF statistic	Episode(s)
Montana (MT)	3.9162**	94M03–95M10, 04M06–07M11
North Carolina (NC)	0.1472	
North Dakota (ND)	1.7347	
Nebraska (NE)	1.7782	
New Hampshire (NH)	5.9468***	86M04–87M07, 01M11–06M07
New Jersey (NJ)	4.2077**	86M07–87M10, 02M10–06M09
New Mexico (NM)	3.8277*	04M10–07M07
Nevada (NV)	2.4837	
New York (NY)	3.9831**	85M03–87M12, 88M03–88M10, 01M06–06M10
Ohio (OH)	1.2067	
Oklahoma (OK)	1.2085	
Oregon (OR)	3.5237*	92M05–92M11, 93M04–97M07, 04M02–07M11
Pennsylvania (PA)	3.8800**	88M04–88M12, 03M05–06M11
Rhode Island (RI)	4.6938***	02M03–06M01
South Carolina (SC)	0.9467	
South Dakota (SD)	1.6170	
Tennessee (TN)	1.6736	
Texas (TX)	1.0062	
Utah (UT)	2.1127	
Virginia (VA)	4.6772***	03M01–06M11
Vermont (VT)	3.9377**	87M12–88M12, 02M04–06M12
Washington (WA)	3.6748*	79M01–79M11, 03M05–07M08
Wisconsin (WI)	3.1672	04M04–05M11
West Virginia (WV)	3.3261	05M03–05M12
Wyoming (WY)	3.7067*	93M12–96M10

*** indicates significance at the 1% level.
** indicates significance at the 5% level.
* indicates significance at the 10% level.

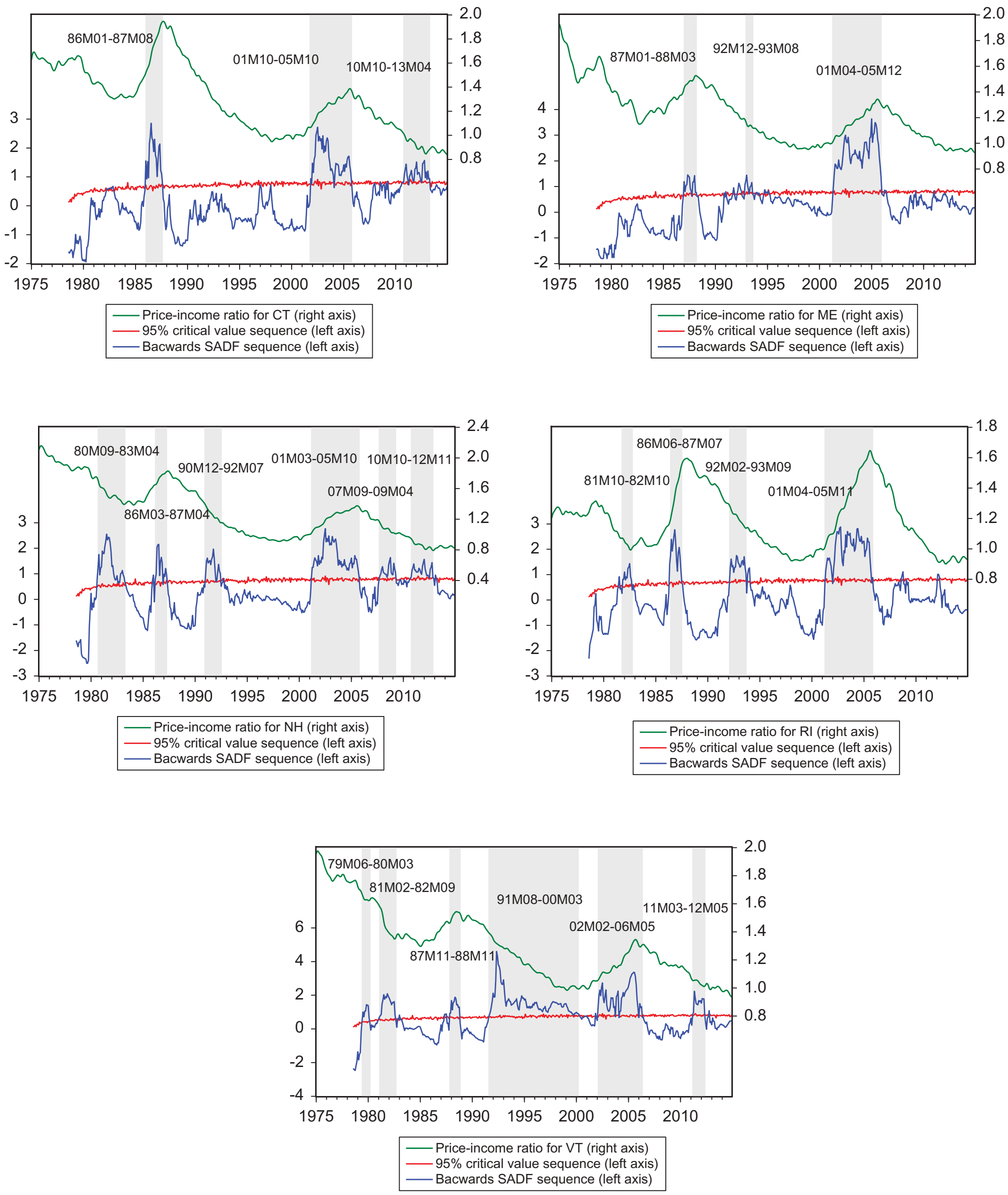

Figure A1. Date-stamping strategy of price–income ratio for Northeast-New England States based on the regression model formulation with an intercept. (a) Connecticut (CT); (b) Maine (ME); (c) New Hampshire (NH); (d) Rhode Island (RI); and (e) Vermont (VT).

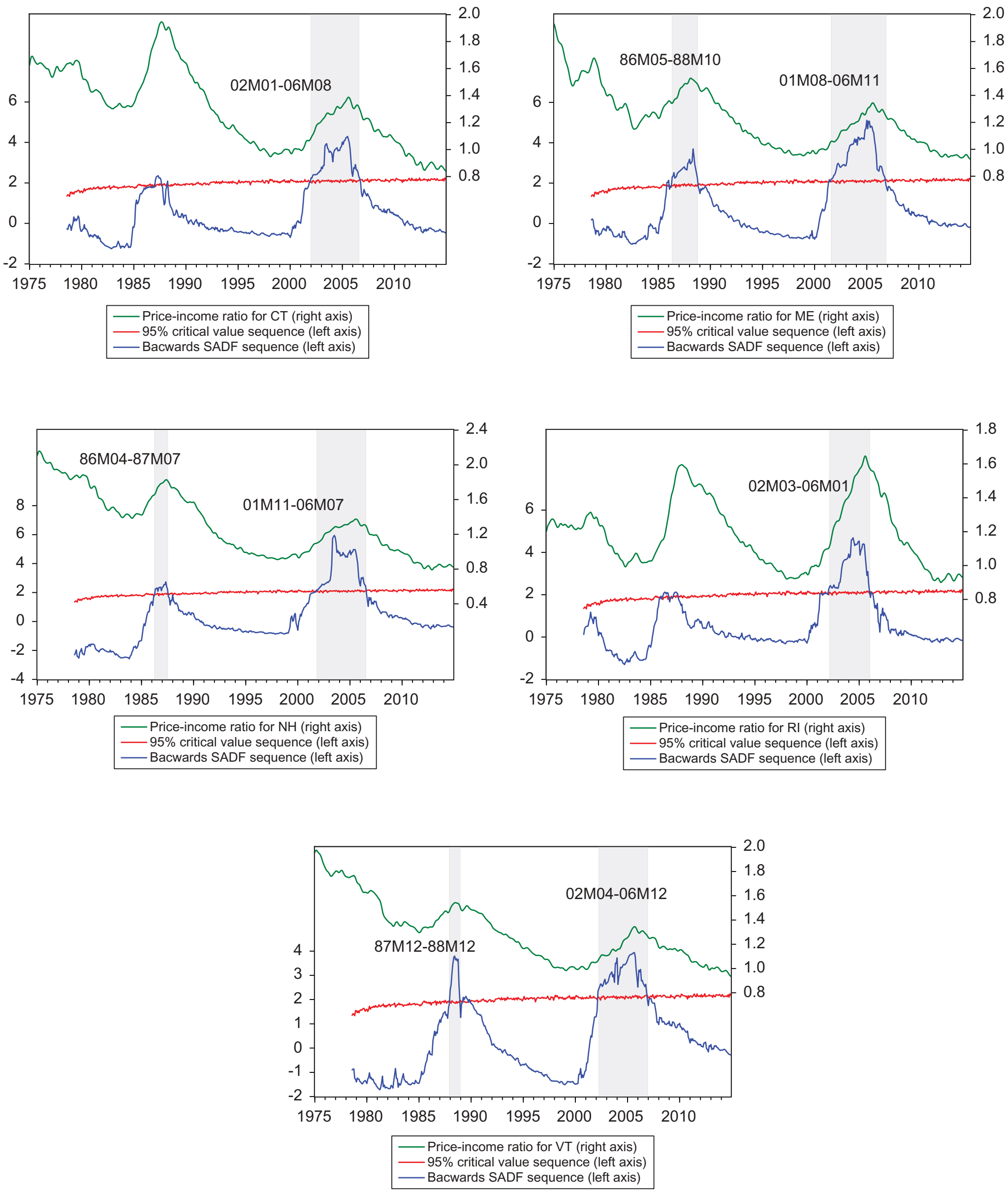

Figure A2. Date-stamping strategy of price–income ratio for Northeast-New England States based on the regression model formulation without an intercept. (a) Connecticut (CT); (b) Maine (ME); (c) New Hampshire (NH); (d) Rhode Island (RI); and (e) Vermont (VT).

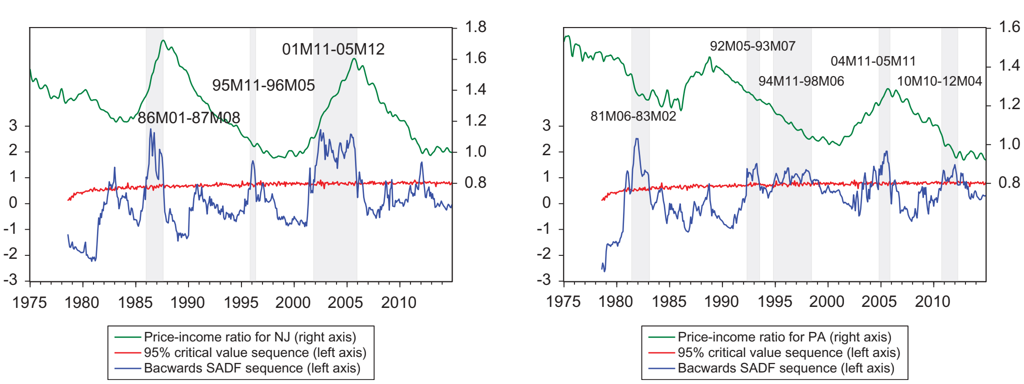

Figure A3. Date-stamping strategy of price–income ratio for Northeast-Mid-Atlantic States based on the regression model formulation with an intercept. (a) New Jersey (NJ) and (b) Pennsylvania (PA).

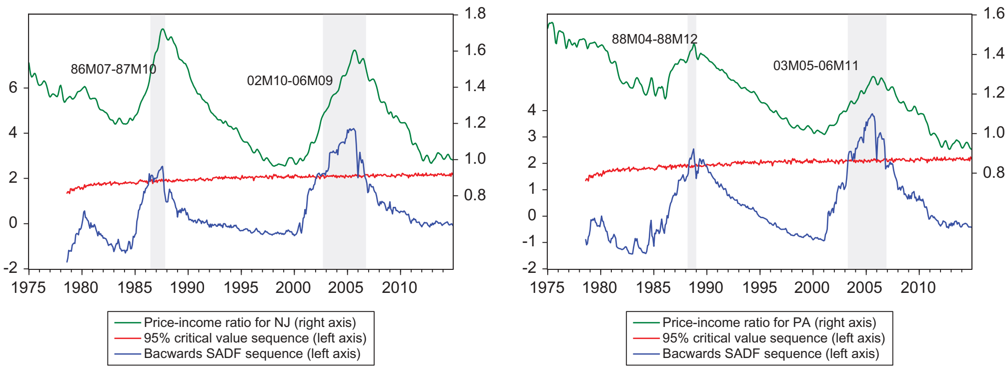

Figure A4. Date-stamping strategy of price–income ratio for Northeast-Mid-Atlantic States based on the regression model formulation without an intercept. (a) New Jersey (NJ) and (b) Pennsylvania (PA).

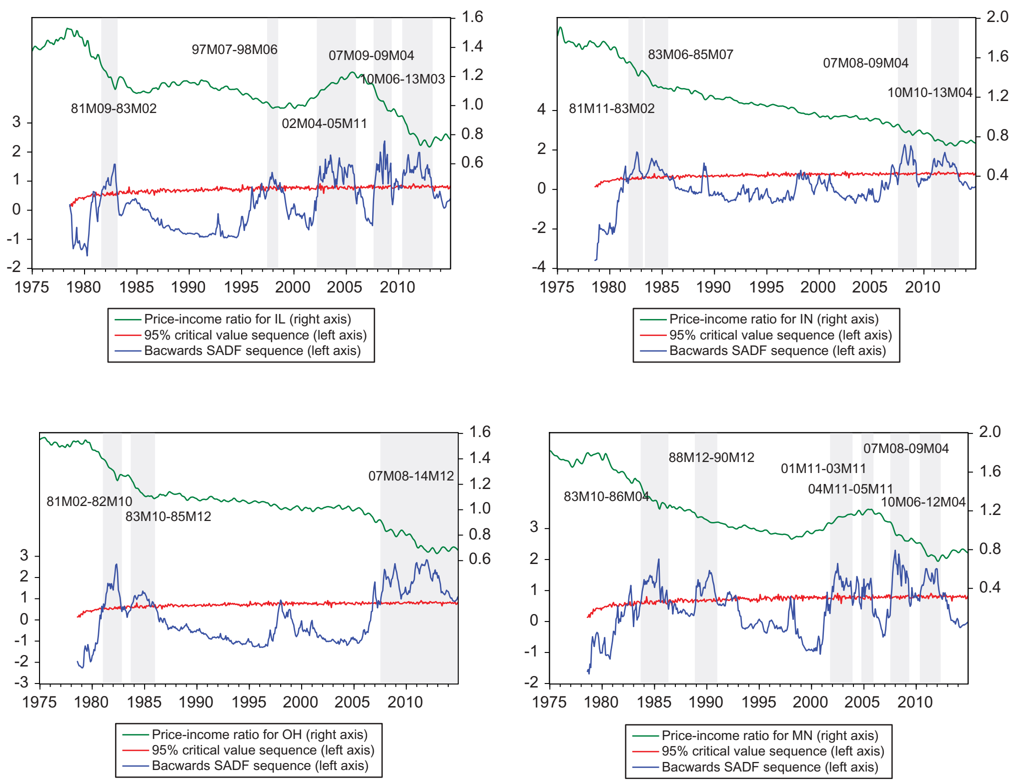

Figure A5. Date-stamping strategy of price–income ratio for Midwest-East/West North Central States based on the regression model formulation with an intercept. (a) Illinois (IL); (b)Indiana (IN); (c) Ohio (OH); and (d) Minnesota (MN).

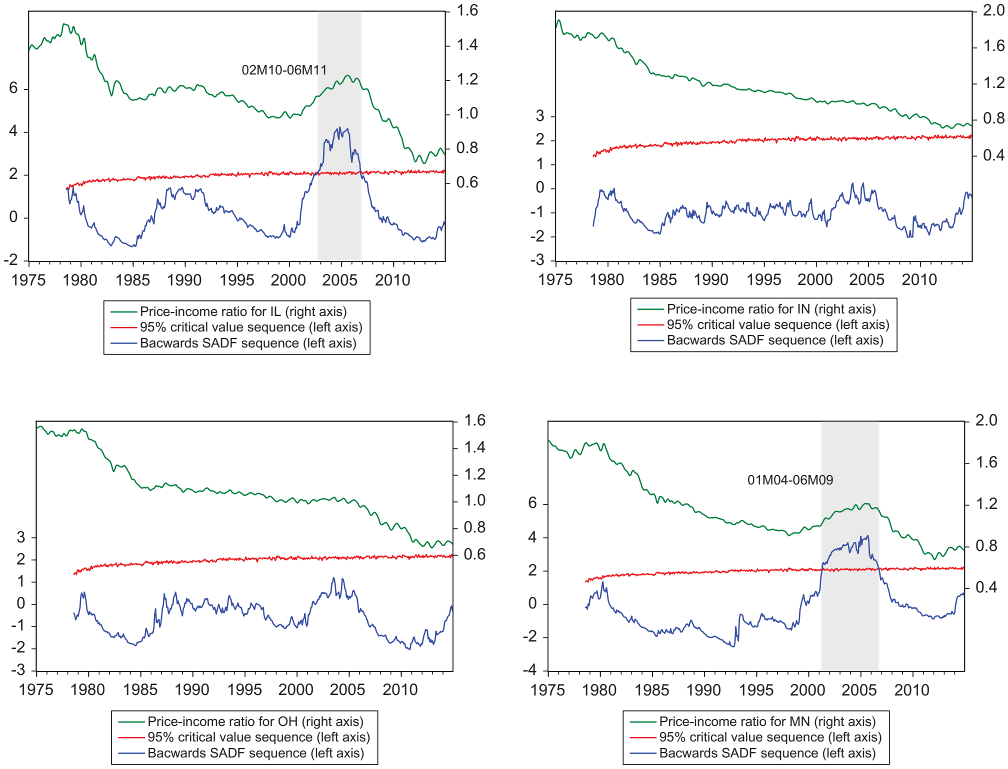

Figure A6. Date-stamping strategy of price–income ratio for Midwest-East/West North Central States based on the regression model formulation without an intercept. (a) Illinois (IL); (b)Indiana (IN); (c) Ohio (OH); and (d) Minnesota (MN).

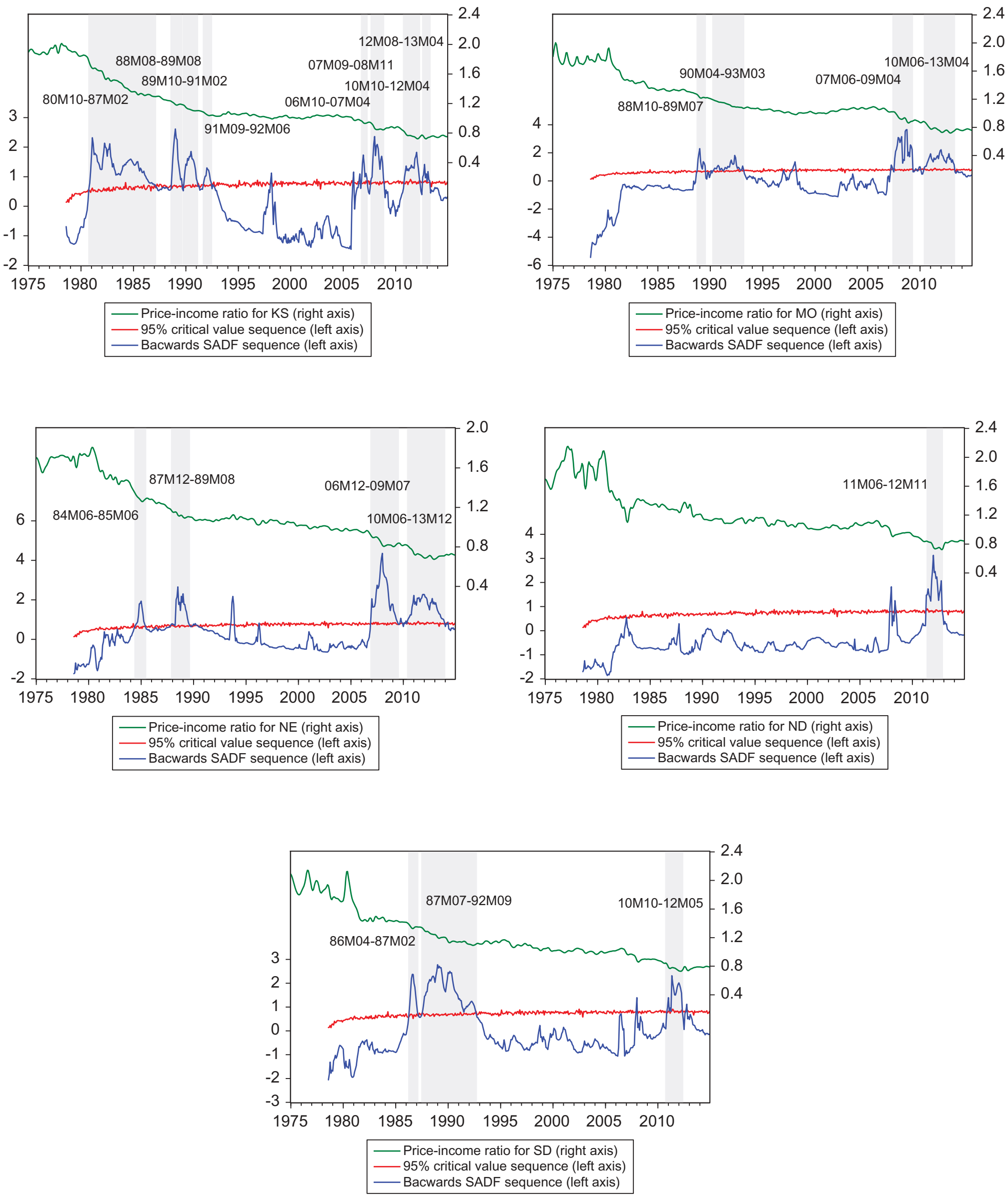

Figure A7. Date-stamping strategy of price–income ratio for Midwest-West North Central States based on the regression model formulation with an intercept. (a) Kansas (KS); (b) Missouri (MO); (c) Nebraska (NE); (d) North Dakota (ND); and (e) South Dakota (SD).

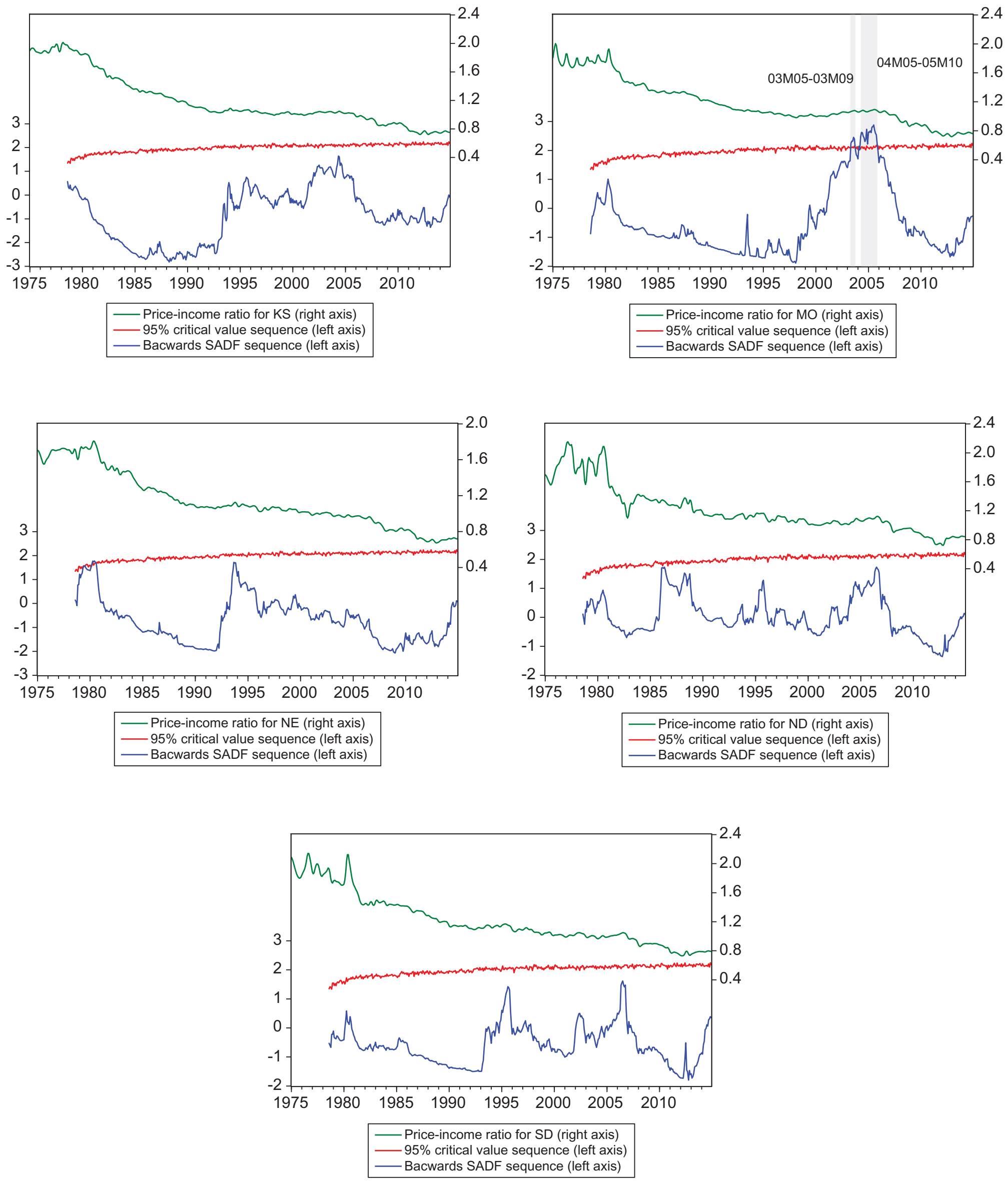

Figure A8. Date-stamping strategy of price–income ratio for Midwest-West North Central States based on the regression model formulation without an intercept. (a) Kansas (KS); (b) Missouri (MO); (c) Nebraska (NE); (d) North Dakota (ND); and (e) South Dakota (SD).

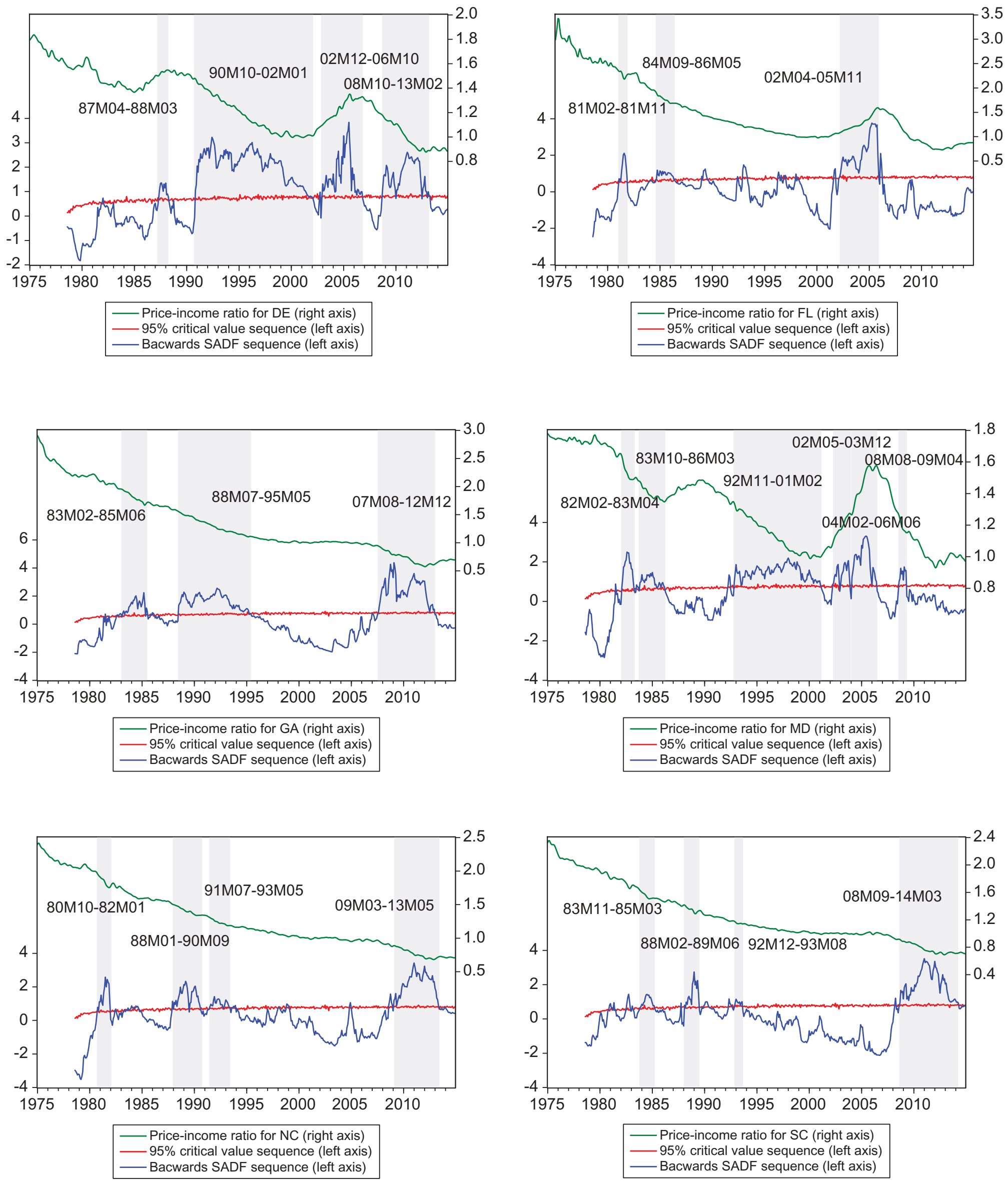

Figure A9. Date-stamping strategy of price–income ratio for South Atlantic States based on the regression model formulation with an intercept. (a) Delaware (DE); (b) Florida (FL); (c) Georgia (GA); (d) Maryland (MD); (e) North Carolina (NC); and (f) South Carolina (SC).

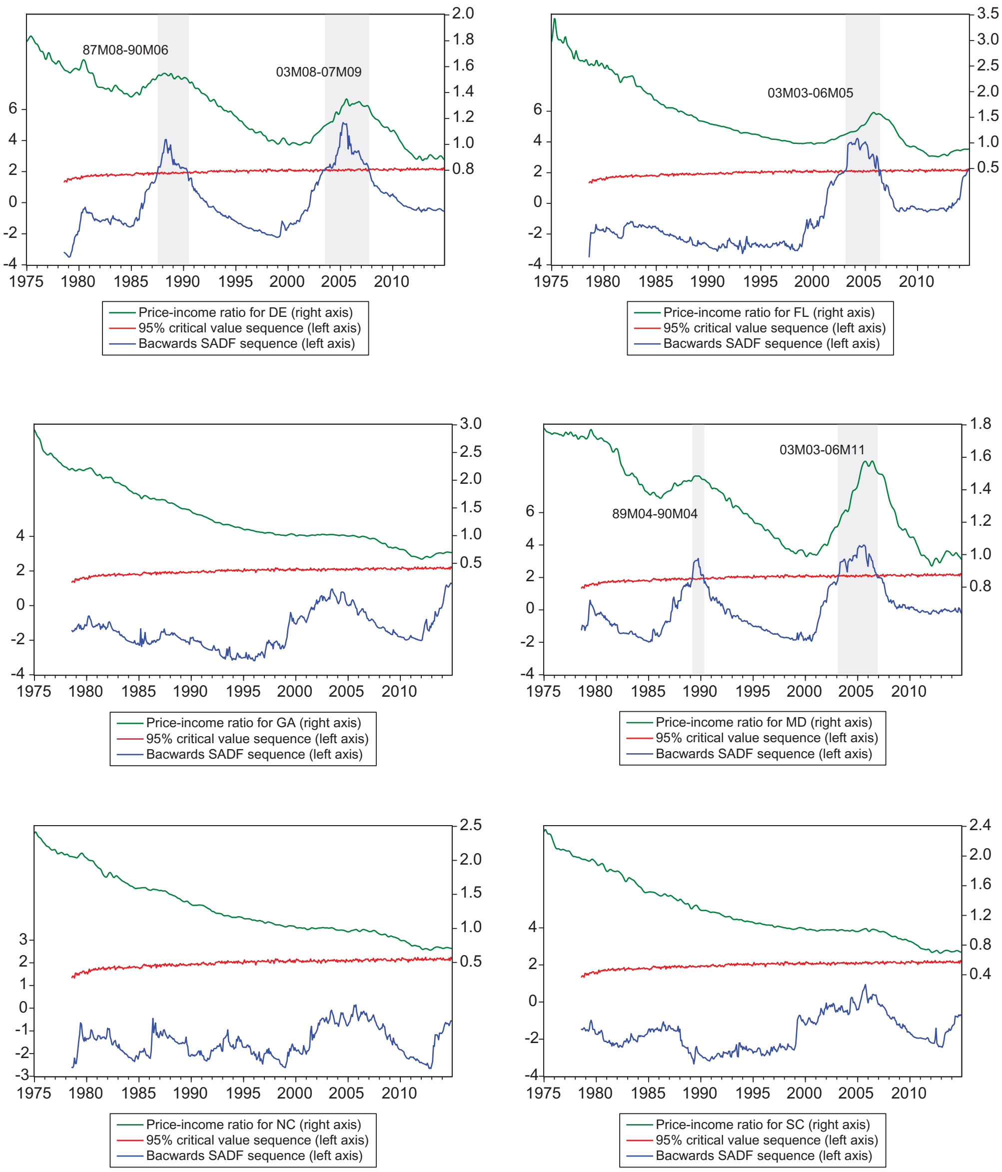

Figure A10. Date-stamping strategy of price–income ratio for South Atlantic States based on the regression model formulation without an intercept. (a) Delaware (DE); (b) Florida (FL); (c) Georgia (GA); (d) Maryland (MD); (e) North Carolina (NC); and (f) South Carolina (SC).

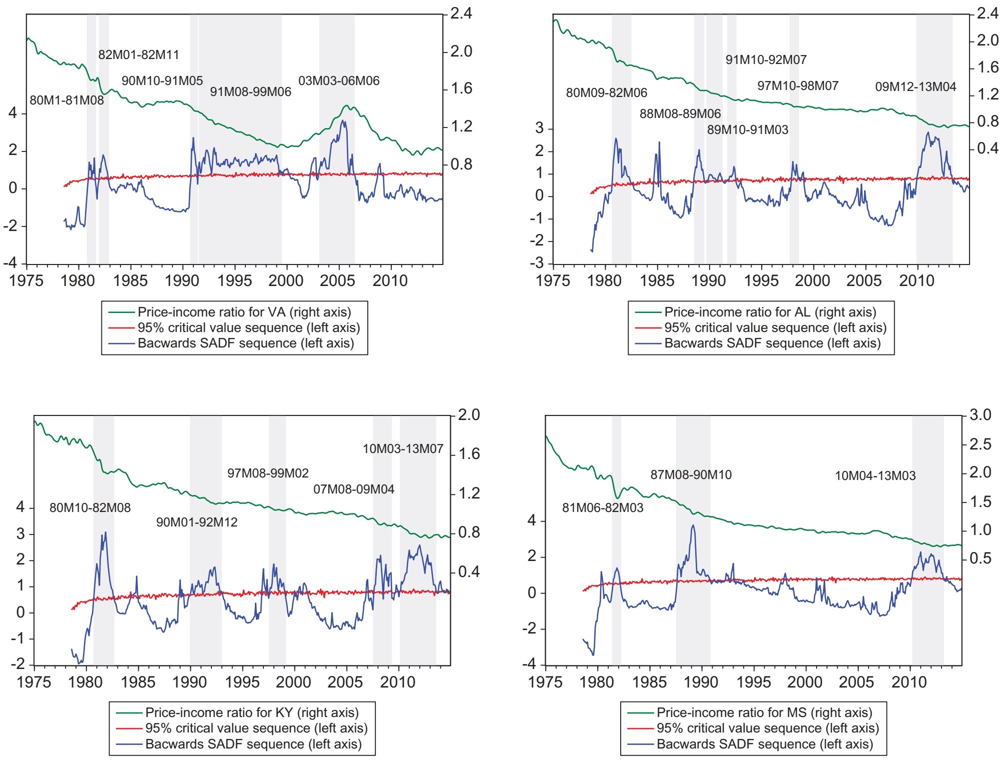

Figure A11. Date-stamping strategy of price–income ratio for South Atlantic/East South Central States based on the regression model formulation with an intercept. (a) Virginia (VA); (b) Alabama (AL); (c) Kentucky (KY); and (d) Mississippi (MS).

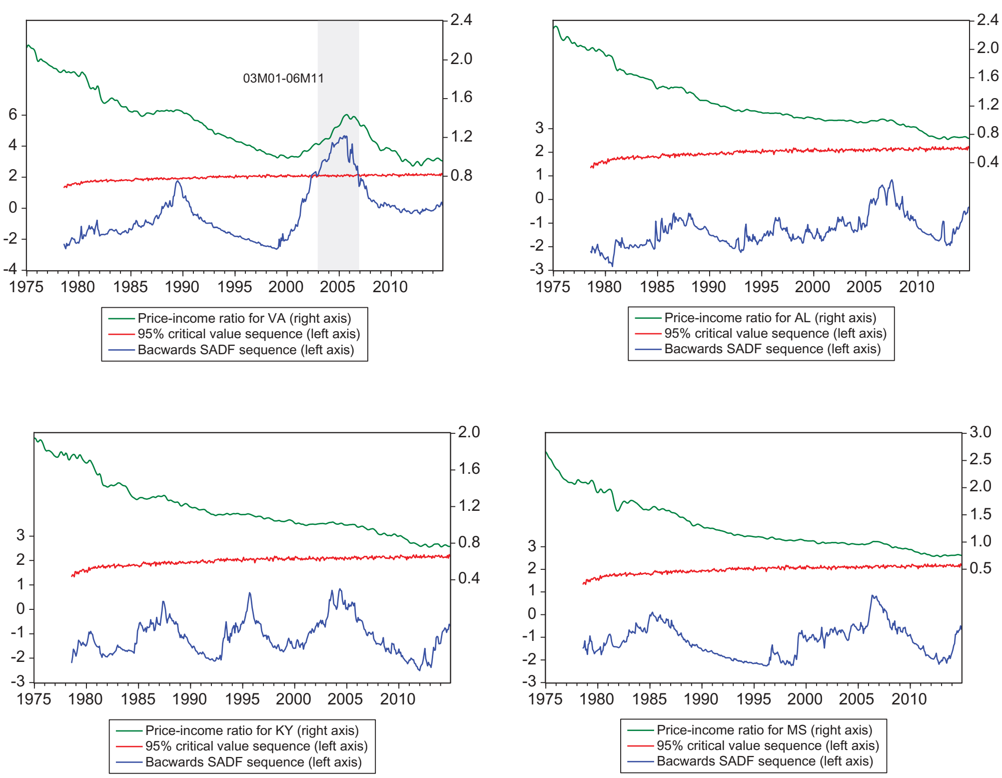

Figure A12. Date-stamping strategy of price–income ratio for South Atlantic/East South Central States based on the regression model formulation without an intercept. (a) Virginia (VA); (b) Alabama (AL); (c) Kentucky (KY); and (d) Mississippi (MS).

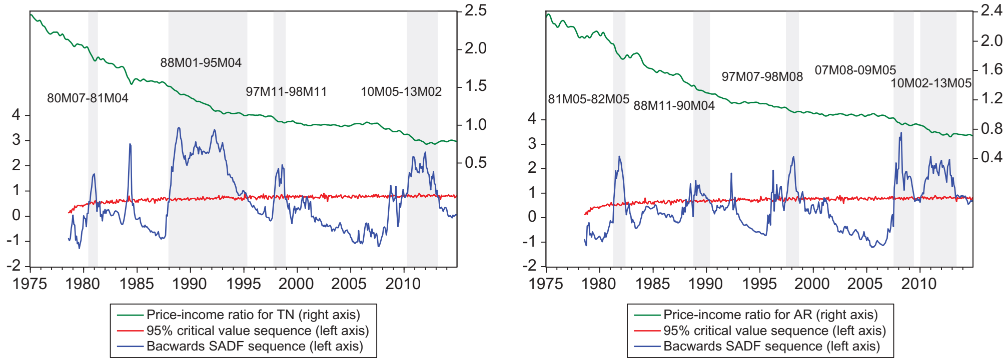

Figure A13. Date-stamping strategy of price–income ratio for East/West South Central States based on the regression model formulation with an intercept. (a) Tennessee (TN) and (b) Arkansas (AR).

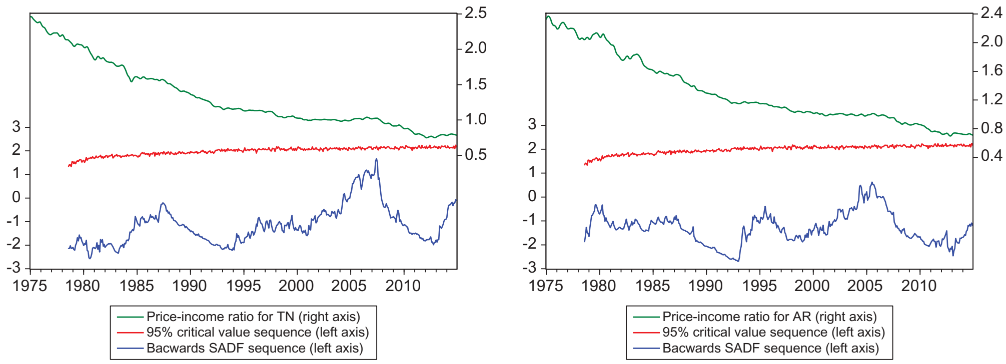

Figure A14. Date-stamping strategy of price–income ratio for East/West South Central States based on the regression model formulation without an intercept. (a) Tennessee (TN) and (b) Arkansas (AR).

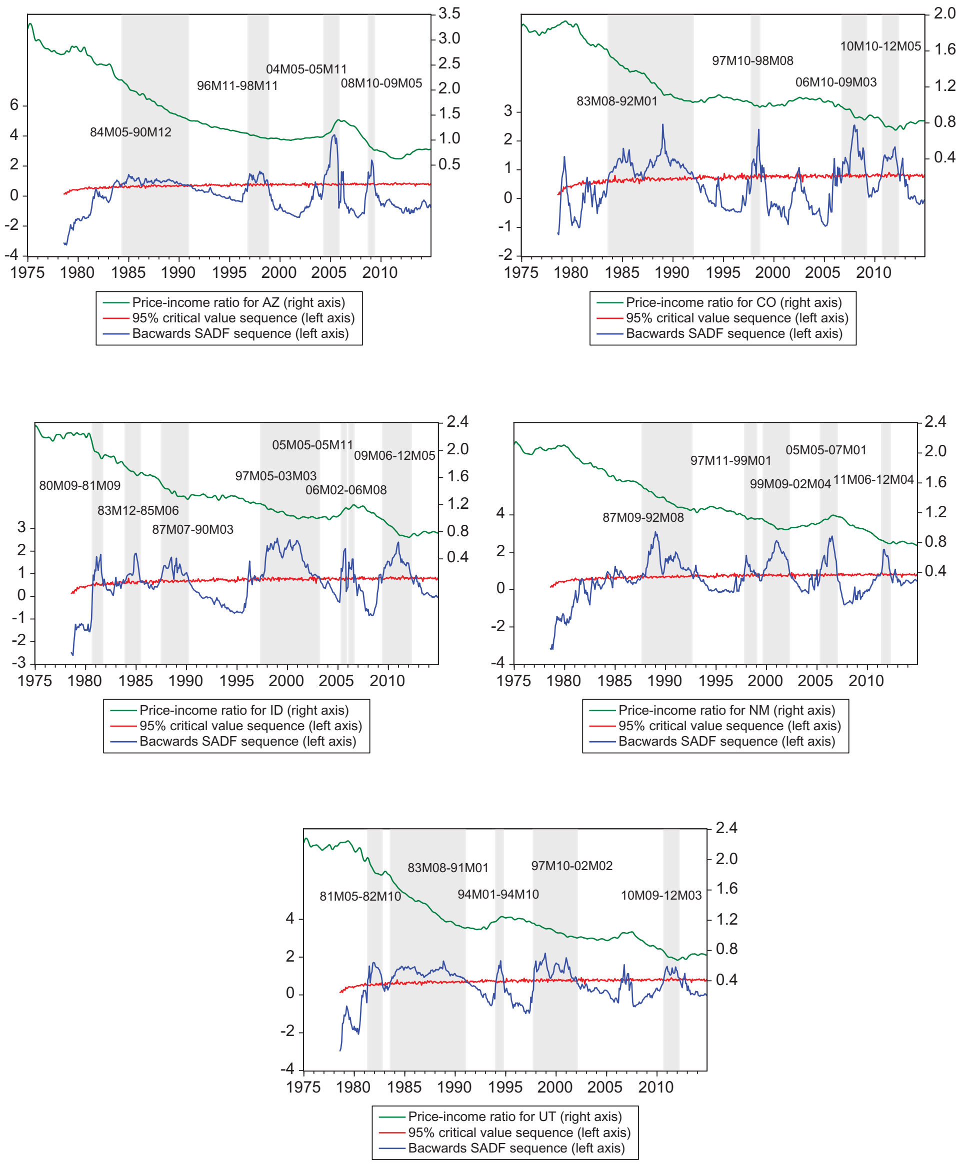

Figure A15. Date-stamping strategy of price–income ratio for West-Mountain States based on the regression model formulation with an intercept. (a) Arizona (AZ); (b) Colorado (CO); (c) Idaho (ID); (d) New Mexico (NM); and (e) Utah (UT).

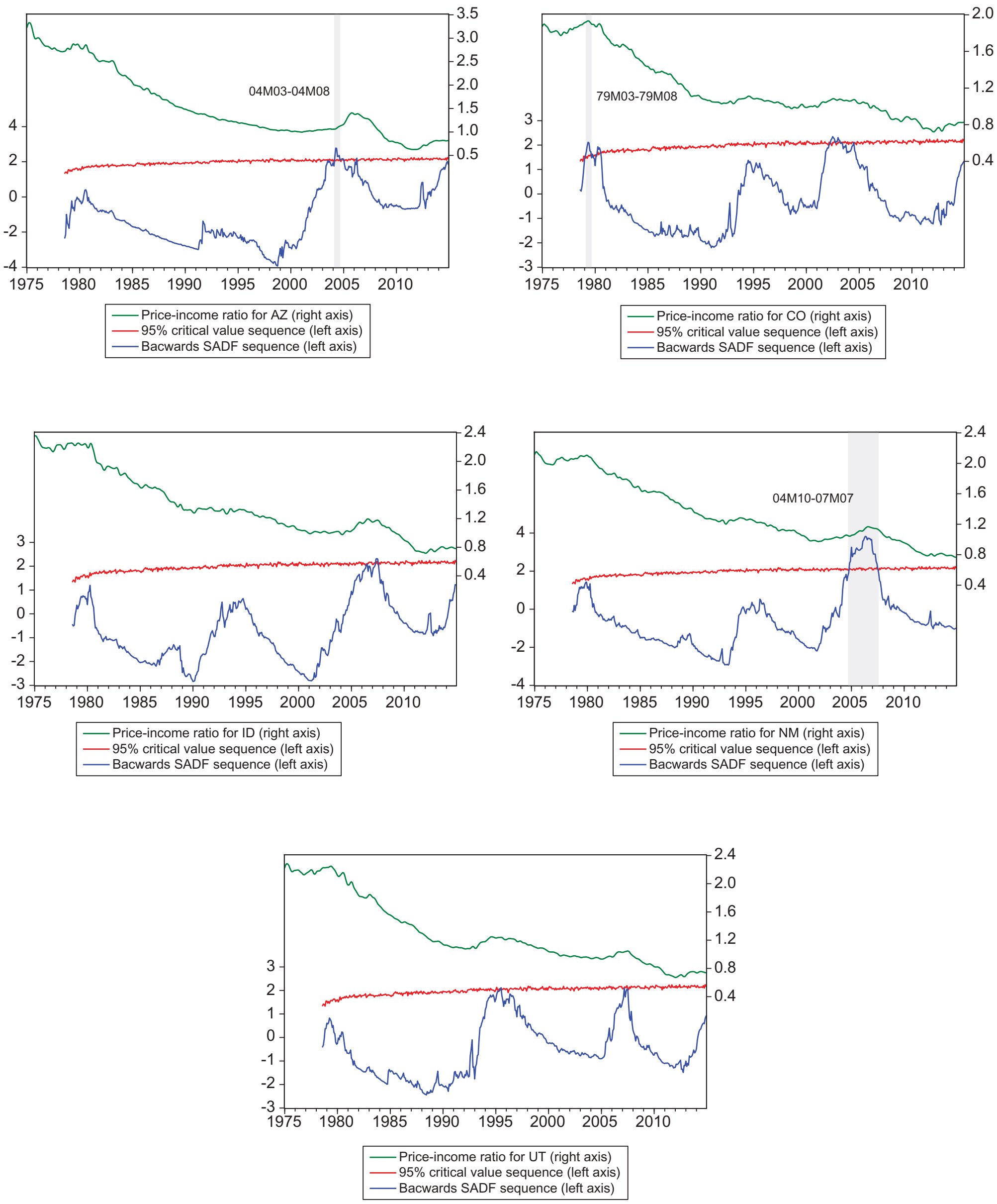

Figure A16. Date-stamping strategy of price–income ratio for West-Mountain States based on the regression model formulation without an intercept. (a) Arizona (AZ); (b) Colorado (CO); (c) Idaho (ID); (d) New Mexico (NM); and (e) Utah (UT).

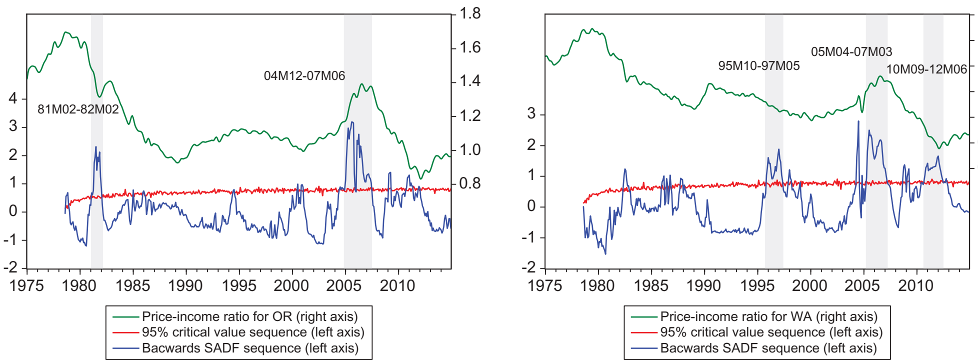

Figure A17. Date-stamping strategy of price–income ratio for West-Pacific States based on the regression model formulation with an intercept. (a) Oregon (OR) and (b) Washington (WA).

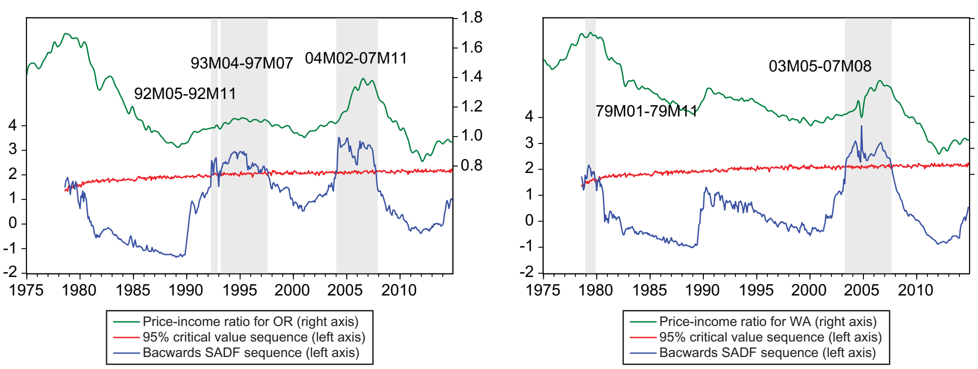

Figure A18. Date-stamping strategy of price–income ratio for West-Pacific States based on the regression model formulation without an intercept. (a) Oregon (OR) and (b) Washington (WA).

Chapter 6

Do 18th Century ‘Bubbles’ Survive the Scrutiny of 21st Century Time Series Econometrics?

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Do 18th century 'bubbles' survive the scrutiny of 21st century time series econometrics?

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HIGHLIGHTS

- We used the Phillips et al. (2015) approach to test for historical bubbles.
- The possibility of non-stationary volatility in the time series was considered using the Harvey et al. (2016) approach.
- We confirm, using modern methods the existence of bubbles for the South Sea Company.
- A number of other 18th century financial company bubbles were identified for the first time.

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ABSTRACT

Applying the methods of Phillips et al. (2015, PSY), while considering the possibility of non-stationary volatility (Harvey et al., 2016), evidence of exuberance in share prices is confirmed for the South Sea Company, and established for a number of other 18th century financial organisations, for the first time. The timings of these bubble episodes show signs of possible contagion.

1. Introduction

Financial history reports the presence of bubbles in a range of commodity markets, for example, Tulipmania of 1634–1637, the Stock Market Crash of 1929, Japan's asset price bubble in the 1980s, the 1990s NASDAQ bubble. Renewed interest in the existence of bubbles has been rekindled with the consequences of the GFC, and recent empirical developments, including the date-stamping tests methods proposed by Phillips, Shi, and Yu (2015, PSY), has introduced a degree of rigour (and flexibility[1]) into their identification.

In this paper we subject two famous price series; the South Sea Company and the Mississippi Company and six other under-researched 18th century financial series, to the rigours of the 21st century tests of Phillips, Shi, and Yu.

2. Background

The motivation of the Mississippi and the South Sea schemes was to refinance the national debts accumulated during the War of the Spanish Succession, see Hamilton (1947) and Dickson (1967). The Mississippi 'bubble' resulted from John Law's 'system' to acquire the French national debt accumulated by the wars of Louis XIV, using equity. Similarly, the South Sea Bubble involved a company (the South Sea Company) that acquired some outstanding British government debt in 1720. Several studies have investigated the Mississippi, the South Sea or related companies for bubbles see Neal (1990), Carlos et al. (2002) and Temin and Voth (2004).

3. Data

The log daily share prices for the relevant companies are obtained from Frehen et al. (2013) and shown as Fig. 1; the Mississippi Company share price in livres between 2 July 1719 and 14 November 1720 (T=385) is shown as Fig. 1a, where T is sample size; the

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[1] Multiple bubble episodes can be identified punctuated by periods of calm.

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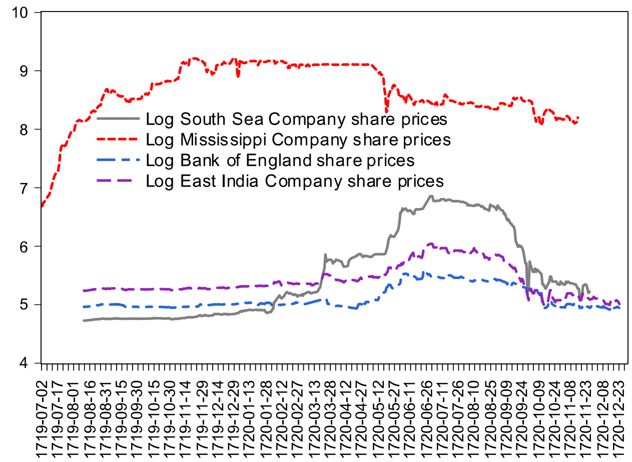
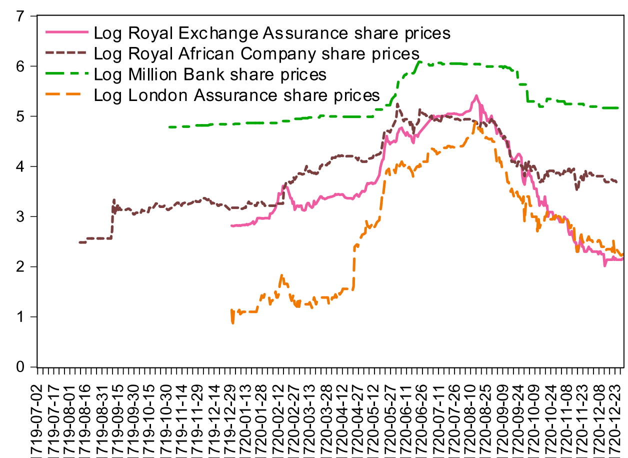
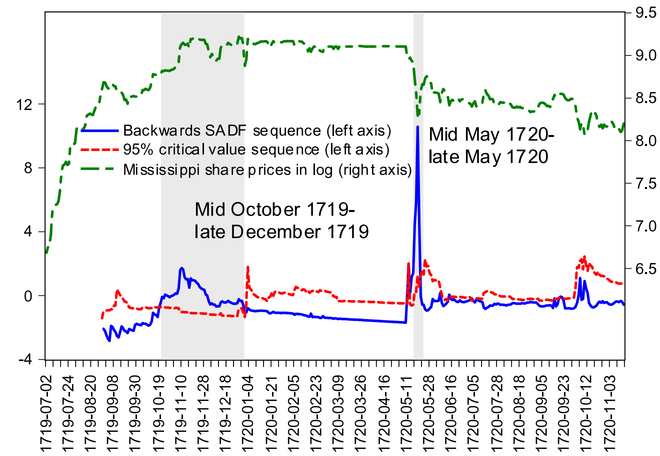


(a)

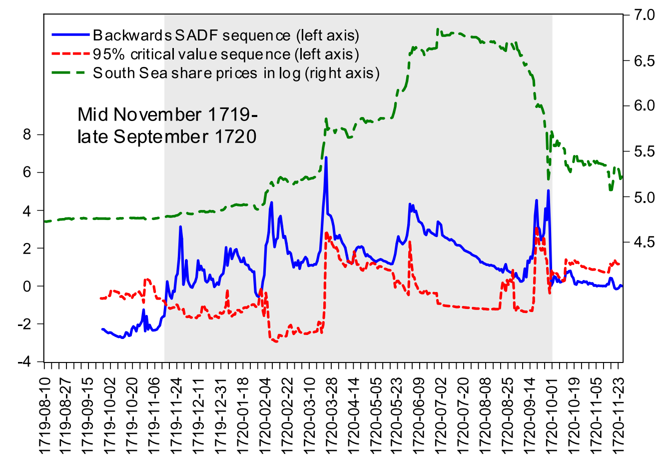


(b)

Fig. 1. The time series plot of the log daily stock prices (Julian dates).



(a) Mississippi Company.



(b) South Sea Company.

Fig. 2. Date-stamping strategies for the Mississippi and South Sea share prices between 1719 and 1720. (For interpretation of the references to colour in this figure legend, the reader is referred to the web version of this article.)

South Sea Company share price in pounds between 10 August 1719 and 23 November 1720 (T=393) is also shown as Fig. 1a with the Bank of England (T=393) and East India Company (T=417). Fig. 1b also shows the time series plot of the share price per pound for London Assurance (T=307), Million Bank (T=348), Royal African Company (T=418) and Royal Exchange Assurance (T=294).

4. Method

We apply the right-tailed unit root test of Phillips, Shi, and Yu (2015) to examine evidence of explosive behaviour in historical stock prices. The martingale null with an asymptotic drift is specified as:

$$H_0 : y_t = dT^{-\eta} + y_{t-1} + \varepsilon_t, \qquad \varepsilon_t \sim \mathrm{NID}(0, \sigma^2), \tag{1}$$

where d is a constant and the localizing coefficient η is great than 1/2. The alternative hypothesis is a mildly explosive process:

$$H_1 : y_t = \delta_T y_{t-1} + \varepsilon_t, \tag{2}$$

where $\delta_T = 1 + cT^{-\theta}$ with $c > 0$ and $\theta \in (0, 1)$.

The following regression model is estimated:

$$\Delta y_t = \alpha + \beta y_{t-1} + \sum_{i=1}^{K} \gamma_i \Delta y_{t-i} + \varepsilon_t, \tag{3}$$

where α is an intercept.

The generalized sup ADF (GSADF) test relies on repeated estimation of the ADF test regression of Eq. (3) on subsamples of the data in a recursive fashion. The window size r_w expands from r_0 to 1, where r_0 is the minimum window size. The ending point r_2 varies from r_0 to 1 and the starting point r_1 varies from 0 to $r_2 - r_0$. The GSADF statistic is the largest ADF statistic over the range of r_1 and r_2:

$$\mathrm{GSADF}(r_0) = \sup_{\substack{r_2 \in [r_0, 1] \\ r_1 \in [0, r_2 - r_0]}} \mathrm{ADF}_{r_1}^{r_2}.$$

The backward SADF (BSADF) statistic is defined as the sup value of the ADF statistic sequence:

$$\mathrm{BSADF}_{r_2}(r_0) = \sup_{r_1 \in [0, r_2 - r_0]} \mathrm{ADF}_{r_1}^{r_2},$$

where the BSADF statistic and its corresponding critical value are used for dating the origination and termination dates of a bubble. The minimum window size r_0 is equal to $0.01 + 1.8/\sqrt{T}$. A fixed lag order of 0 is also selected. Critical values are obtained by following the method of Harvey et al. (2016), which uses a wild bootstrap

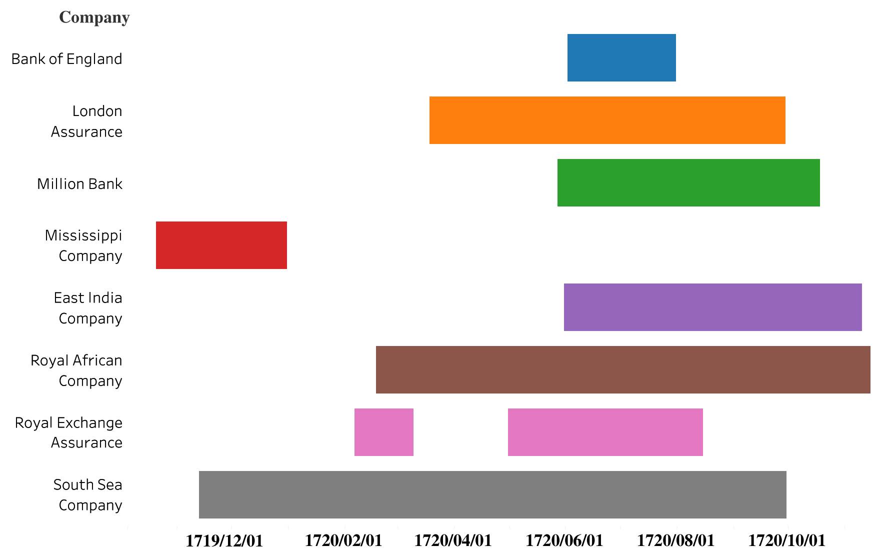


Fig. 3. Date-stamping strategies for all eight companies using wild bootstrapping.

with 2000 replications, to take into account the presence of any possible heteroscedasticity.[2]

The PSY procedure is often applied to a price-fundamental ratio to assess explosive behaviour where the rejection of the null hypothesis of a unit root implies explosive behaviour for y_t. If the time series y_t involves an economic fundamental, we conclude that a finding of explosive behaviour denotes the presence of *a bubble*. Otherwise, we may only conclude that a finding of explosive behaviour in a price series is evidence of *exuberance*, and such an episode is described as *an exuberant episode*, see Hu and Oxley (2016).

5. Results

We present our bubble detection results for two most famous Mississippi and South Sea share prices in Fig. 2. Results for the date-stamping outcomes for the Mississippi share price are presented as Fig. 2a. The null of no explosive behaviour is rejected at the 1% level as the GSADF test statistic is much greater than the critical value (10.5665>4.8017), suggesting very strong evidence of explosive behaviour and the date-stamping outcomes in Fig. 2a seem to provide 'some evidence', where the test statistic (blue solid line) exceeds the critical value sequences (red dashed line) in May 1720. However, we should not interpret such results as evidence of exuberance in the share prices due to the fact that rejection is caused by a 'collapse and recovery' episode (not exuberance) in May 1720, which is clearly shown in Fig. 2a. During this period, the share price (green dashed line) declined sharply from mid May and rebounded at the end of May 1720.[3]

The date-stamping outcomes for the South Sea share price are presented as Fig. 2b. The null hypothesis of no explosive behaviour is rejected at the 1% level (6.8010>3.1459), indicating strong evidence of exuberance. As shown in Fig. 2b, we observe an exuberant episode between mid November 1719 and late September 1720. Such an exuberant episode is closely related to the rapid growth and burst of the well-known South Sea Bubble. Thus we provide some signs of exuberance to support the famous South Sea episode in 1720 by applying the PSY procedure, which is the novel in the literature.

The South Sea Bubble is related to the spectacular rise and fall in the South Sea stock price, however, as discussed in Frehen et al. (2013), the South Sea Company does not experience the largest price increase and several other major companies also experience significant increases and falls during 1720. For example, the East India share price increased over 100% and the Bank of England share price surged by 60% before they fall back (Hoppit, 2002). Carlos et al. (2006) also point out that the Royal African Company is more speculative than other joint stock companies during the South Sea Bubble. We, therefore, test for explosiveness in stock prices for the other six major corporations in the British market. Fig. 3 displays the identified exuberant episodes for all eight companies considered in our study. It is known that the PSY procedure can sometimes lead to the identification of collapse episodes, rather than bubbles, see Hu and Oxley (2016). However, Phillips and Shi (2017, forthcoming) now recommend that users of their procedure be careful to identity such episodes ex-post, hence we only present and plot the exuberant episodes (not those identified as collapses) in Fig. 3.

Several interesting results can be concluded from Fig. 3. Firstly, the South Sea Company experiences the first exuberant episode in the British market, and such an episode is closely followed by those of the Royal African Company, London Assurance and other companies. We also notice that the South Sea episode is not the first one to burst and it lasts the longest period. Secondly, several British share prices also exhibit strong signs of exuberance (e.g., East India Company, London Assurance, Million Bank, the Royal African Company and the Royal Exchange Assurance) as the null of no explosive behaviour can be strongly rejected at the 1% level. We can clearly see that these companies experience exuberant episodes during the South Sea episode in Fig. 3. Thirdly, there is little evidence of exuberance for the Bank of England as the null of no explosive behaviour cannot be rejected at the 10% level although a very short exuberant episode seems to exist between June 1720 and July 1720. This result is perhaps not surprising as the share price of Bank of England is widely regarded as the least speculative stock among the major joint-stock companies (Carlos and Neal, 2006). Finally, our results seem to draw a very interesting conclusion for this period in that the British share market was generally much more speculative, as the South Sea Company was not the only one experiencing exuberance in its share price.

2 The wild-bootstrap critical values are generated using an add-in package for the EViews, see Caspi (2017, forthcoming).

3 We also apply the PSY procedure to the Mississippi share price under the regression model without an intercept using wild bootstrapping. We identify an exuberant episode between September 1719 and May 1720, which coincides with the traditional view of the Mississippi Bubble, see Hu and Oxley (2017). This result will be further investigated in the future.

6. Conclusion

We have subjected two famous and six less famous 18th century share price series to the rigours of 21st century tests in the form of Phillips, Shi, and Yu's (2015) procedure. Exuberance is confirmed in the South Sea Company and established for several other British companies. We also identify the timing and collapse of each of these company's periods of exuberance. The timing of these relationships is provided as some possible evidence of spillovers or contagion in exuberance in the financial market more generally during this period. The famous 'South Sea Bubble' survives the rigours of the 21st century tests and the 'tests' identify a new member of the bubbly club.

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We are grateful to an anonymous referee for her/his constructive comments, which significantly improve the quality of this paper. We particularly wish to thank Professor Peter Phillips for his detailed and insightful comments on this and other papers where we have utilized his PSY test. Thanks also go to Professor Rob Taylor and Professor Dave Harvey for discussions on non-stationary volatility.

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Chapter 7

Bubble Contagion: Evidence from Japan's Asset Price Bubble of the 1980s-90s

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R30

ABSTRACT

This paper investigates the most documented asset price bubbles of the 1980-90s in Japan, and subjects them to the rigours of recent econometric tests. We focus on testing for bubbles in Japan's stock and real estate markets from 1970Q1 to 1999Q4 using the right-tailed unit root test of Phillips et al. (2015a, PSY). We also utilize the econometric methods of Greenaway-McGrevy and Phillips (2016) to explore the possibility of contagion between these two markets. The paper offers significant econometric-based evidence of bubbles in both markets during this period in Japan and more importantly, for the first time in the literature, formal tests of bubble contagion from Japan's stock market to its real estate market. Our findings may help to understand why Japan's real estate bubble collapsed after the stock price bubble, as the bubble-like behaviour from the stock market migrates to the real estate market.

1. Introduction

Japan's asset price bubble of the late 1980s to early 1990s has attracted considerable attention in the literature, as it is widely regarded as one of the most famous bubble episodes of the past 400 years, see Brunnermeier and Schnabel (2016) for a recent survey. As discussed in Okina et al. (2001), Japan's bubble economy was characterised by a rapid rise in asset prices; overheating of economic activity and the expansion of the money supply and supply of credit. From the late 1980s, Japan experienced the most intense episode of growth in asset prices (e.g., art, antiques, golf course memberships, land, real estate, stocks) with prices reaching historical highs. However, during the second half of the 1980s, the growth in asset prices was widely believed to be reflected in economic fundamentals, rather than a bubble, suggesting future, positive, sustained economic growth in Japan. This we now know did not materialize.[1]

What makes Japan's recent asset price bubble interesting is that it has several very unusual distinguishing features, in particular, unlike most historical financial bubbles, which experience only a single type of bubble asset, for example, the South Sea and Mississippi Bubbles or the British Railway Mania see Hu and Oxley (2018b, 2017). Japan experiences at least two sets of bubbles (i) stock prices and (ii) real estate prices. This makes Japan an ideal candidate to examine evidence of bubble spillovers. The second important feature of Japan's bubble economy is that, despite the passage of three decades, neither the stock nor real estate markets have regained their heights achieved in the late 1980s or early 1990s. The third classic feature is that Japan's bursting bubbles foreshadowed a decade of economic stagnation with a growing government debt. The aftermath of economic stagnation in the 1990s, following the bursting of the asset price bubble is commonly referred to as "the Lost Decade".

The aim of this paper is to subject Japan's asset market and real estate price data to firstly, the rigours of modern time series econometric methods using the right-tailed unit root test of Phillips et al. (2015a, PSY) and secondly, tests of contagion following the approach of Greenaway-McGrevy and Phillips (2016).[2] We focus on the real estate market rather than land market for two obvious reasons. First, the rise in land prices certainly contributes to higher house prices. This is due to the fact that the land cost is more than 60% of housing costs in Japan's large cities, over 80% in the surrounding city of Tokyo and approximately 98.5% in the central district of Tokyo, see Noguchi (1994). Second, it is more difficult to obtain the fundamental value for land, which is the discounted sum of all future rent revenues. Instead, we can certainly obtain a house price-rent ratio for Japan as a measure of the

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[1] The causes of devastating Japan's bubble economy is beyond scope of this paper.

[2] As has now become common in finance, we use the word *contagion* to refer to a situation (market, region, country, etc,) where a shock in one sector/area spreads out and effects other by for example, price movements. In our particular case the contagion we will refer to will likely have a particular time series pattern (which we will seek to test), where the contagion effect first increases in intensity and then diminishes as the shock disperses. Financial contagions may manifest themselves as negative externalities diffused from one (crashing) market to another.

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real estate fundamental. This paper, therefore, empirically examines the presence of bubbles in two inflated asset markets and crucially is able to identify the origination and termination dates in each to ascertain whether these econometrically identified bubble episodes coincide with historical events. Furthermore, an additional and perhaps more important focus of the paper is to examine potential contagion between the two markets in Japan by utilising the recent time-varying regression methodology of Greenaway-McGrevy and Phillips (2016). Existing studies seem to suggest that Japan's stock price bubble migrates to real estate or land market during the 1980s, see Voth (2003) and Basile and Joyce (2001). However, no formal statistical tests have been applied to explore possible spillovers in Japan. This assertion will be formally tested below.

The paper, therefore, makes two significant contributions to the existing literature. The first is that we re-visit the widely assumed Japanese asset price bubble episode and subject it to the rigours of the PSY approach which allows us to 'date-stamp' the identification of start- and -end dates of a bubble, from which we can then focus on identifying whether it is based upon its irrationally exuberant stock and real estate markets or on economic fundamentals. To anticipate, we will present significant evidence of bubbles in Japan's stock price-earnings and house price-rent ratios and identify the origination and collapse dates of both stock and real estate bubbles, which coincide with the traditional stories of Japan's asset price bubble. The second important contribution is that by utilizing the recently developed time-varying regression methodology of Greenaway-McGrevy and Phillips (2016) we can establish for the first time that there is robust and rigourously determined evidence of bubble spillovers between the stock and real estate markets during the Japanese bubble economy period.[3] Our results clearly show signs of migration from a stock price bubble to a real estate bubble.[4] This should help understand why the real estate bubble bursts after the stock price bubble, as the stock market activity not only migrates to, but fuels, the booming real estate market as well.

The paper is organized as follows. Section 2 provides some background on this unique period in Japan's history and then reviews the existing relevant literature. Section 3 provides a brief description of rational bubble models and econometric methods of Phillips et al. (2015a) and Greenaway-McGrevy and Phillips (2016) and Section 4 describes the data. Section 5 presents the empirical findings and Section 6 concludes.

2. Literature review

2.1. Overview

It now seems clear that there has been an unprecedented rise in Japan's land (real estate) and stock prices. According to Malkiel (2003, p.77), the value of real estate and stocks increased over 75 and 100 times from 1955 to 1990. Two 'myths' emerged and were associated with the inflated Japanese real estate and stock markets. The first was that land prices could never fall, and the second was that stock prices could only rise. Both myths were fueled by large amounts of cash savings (Malkiel, 2003, p.78). A common belief related to Japan's price-earnings ratio, was that stock prices were too high or over-valued, see Ueda (1990) and French and Poterba (1991). In particular, Japan's price-earnings ratio during the 1980s was several times higher than in the US. French and Poterba (1991) also suggest that accounting differences alone, cannot explain the doubling of the Japanese price-earnings ratio in 1986, nor its decline in 1990. This provides some support for Ueda (1990) and French and Poterba (1991) believe that 'irrationally exuberant stock prices' produced a bubble. Moreover, the value of the stock market achieved remarkable growth. Stone and Ziemba (1993) stated that the stock market had a value of approximately US$ 4 trillion at its peak in December 1989, which represented approximately 44% of the world's stock market value. Turning to the sharp rise in land prices especially during the late 1980s, see Noguchi (1994), Stone and Ziemba (1993) further conclude that the total value of all Japanese property was valued at approximately 20 trillion US Dollars at the end of 1991, which equated to over 20% of the world's total wealth and almost double the total value of the world's stock markets. This led Malkiel (2003, p.77) to suggest that if this were the case, the Japanese could have brought all the property in America by only selling the Tokyo area. It is generally believed, therefore, that both markets are characterised by bubble-like behaviours, see for example, Hardouvelis et al. (1988), Lee (1995), French and Poterba (1991), and Chirinko and Schaller (2001).[5]

The bursting of Japan's asset price bubble in the early 1990s destroyed these two myths. According to Okina et al. (2001), the Nikkei 225 reached a historical high of 38,130 in December 1989 from a start of 12,598 in September 1985, and then dropped sharply to 14,309 in August 1992, which represented a more than 60% decline from the peak. When we consider land prices in September 1990, they were around four times higher than the level in September 1985, only to drop in 1999 by approximately 80% from the peak in 1990. Alan Greenspan described the sudden and dramatic decline in the Japanese stock market as "a correction of the bubble in asset prices".[6]

2.2. Empirical studies on Japan's asset price bubble

A number of studies have attempted to detect a bubble in asset prices using various approaches. Several studies present evidence of bubbles in Japan's exuberant stock market, for example, Hardouvelis et al. (1988), Ueda (1990), Chung and Lee (1998), Chirinko and Schaller (2001), Binswanger (2004) and Asako and Liu (2013). Chan et al. (1998) is one of the few to argue there is no overwhelming evidence of rational speculative bubbles in Japan's stock

[3] In economics, the notion of a *spillover* has been applied to a range of circumstances. The general characteristic of a spillover is that economic events in one context occur because of something else in a seemingly unrelated context. In this sense it is another type of externality, however, spillovers can generate positive as well as negative externalities.

[4] In terms of finance, *migration* relates to e.g., the movement of assets, including the effects of some of their characteristics (i.e. riskiness) from one market and/or region to another. Migration may involve a series of movements through a range of markets, where the effects could be large or small, negative or positive. Migration typically occurs in response to normal economic/finance drivers e.g., expected profitability, risk diversification and is regarded as quite normal characteristics of dynamic financial markets.

[5] Apart from speculation in the stock and real estate or land price markets, another very famous example of Japan's bubble economy is the price of golf course membership. The land market can be proxied by the golf course membership index, which is created by Nikkei on a weekly basis from 1981 for various parts of Japan. The golf index becomes essentially an indicator for the overheated land or real estate market. Stone and Ziemba (1993) conclude that the total value of all golf courses in Japan was approximately US$500 billion by 1989, a value that equates to the sum of the Australian and Swiss stock exchanges. Not only does the membership allow the membership holder and guests to play, but it becomes more importantly as an investment tool for membership holders. Memberships are actively traded like securities by brokers who earn a 2% commission per transaction (Chancellor, 2000, p. 315). According to the Economist("Golf holes," in 21 Jan 1999), the price of membership in one of Japan's most exclusive golf clubs was more than ¥400 million (US $ 2.8 million) during the peak of the early 1990. The quarterly golf course membership index in Tokyo peaked in early 1990 at 600 from a starting value of 100 in 1981. It then declined to around 250 at the end of 1992, which represented more than 60% fall, see Stone and Ziemba (1993). Shiryaev et al. (2015) also argue that the golf course memberships market contains a larger bubble than the Nikkei stock market. Overall, Japan's golf course membership price boom, seems to be, another symbol of the bubble era and worthy of special mention here.

[6] Washington Post, April 18, 1992.

market. Previous studies have also provided some empirical evidence to support the existence of bubbles in the land market, for instance, Noguchi (1994) and Sato (1995).[7] Lee (1995) and Ito and Iwaisako (1996) also investigate the behaviour of stock and land prices during Japan's bubble economy, indicating the presence of bubbles. It therefore seems clear that there have been previous attempts to detect bubbles in Japan, however, none has sought to test the possibility of potential *bubble contagion between stock market and real estate market* in Japan using a rigorous and appropriate statistical test that is designed to identify migration.

2.3. Empirical studies on bubble contagion

Despite numerous studies seeking to identify asset price bubbles using a range of econometric methods, research on bubble contagion across markets is much less frequently investigated. Basile and Joyce (2001) do investigate potential linkages between asset price bubbles (stock and land bubbles), monetary and bank lending variables in Japan during the bubble period. Both the Granger-causality tests and variance decompositions suggest that these two asset markets are related and changes in stock market speculation lead to changes in land market speculation for the period of 1986–1991. Riddel (2011) suggests that housing bubbles are contagious as house prices in Los Angeles contribute to a rapid growth and subsequent decline of house prices in Las Vegas between 2001 and 2008. Shih et al. (2014) identify real estate bubbles in China for most provinces and document evidence of spillovers in surrounding regions of Beijing and Shanghai. Nneji et al. (2015) confirm evidence of speculative bubbles in five of the nine US regions and also show spillovers of speculative bubbles across both contiguous and non-contiguous regions. Teng et al. (2017) also find housing bubbles spillovers from Taipei City (city centre) to New Taipei City (suburbs) based on an empirical analysis of Taipei metropolitan area from 1973 to 2014 using Engle-Granger cointegration and Granger causality tests.

Recently new econometric methods including the migration test of Phillips and Yu (2011) and time-varying regression approach of Greenaway-McGrevy and Phillips (2016) have been developed to study the effects of bubbles across different markets. In particular, Phillips and Yu (2011) detect the presence of bubbles in the US real estate and a selection of commodities. Their finding shows that the US real estate bubble in the early 2000s migrates to the oil and bond markets after the subprime crisis, separately. Gomez-Gonzalez et al. (2017) also apply this method to explore evidence of bubble migration in the housing, currency and stock markets for seven countries. They report evidence of spillovers in these financial markets for some selected countries. The time-varying regression approach of Greenaway-McGrevy and Phillips (2016) shows evidence of bubble spillovers from the Auckland housing markets to the other regional housing markets in New Zealand. The main advantage of new method of Greenaway-McGrevy and Phillips (2016) over Phillips and Yu (2011) is that the former one allows time-varying effects of bubbles transmission. To our knowledge, only two studies have employed the methods of Greenaway-McGrevy and Phillips (2016) to examine spillovers of market contagion. The first is Gomez-Gonzalez et al. (2016), who apply the PSY approach to date-stamp housing bubbles using quarterly data of 20 OECD countries from 1970 to 2015 and investigate further for potential international bubble transmissions from the US to European countries. The second and more closely aligned study is Deng et al. (2017), who find signs of bubble migration from China's stock market to its housing market between 2005 and 2010.

[7] Asako (1991) concludes that the land price bubble is not always rational.

3. Econometric methods

Here we use a two-stage testing procedure. The first stage is to apply the bubble detection test of Phillips et al. (2015a) to identify the origination and collapse dates of bubbles in Japan's stock and real estate markets. The PSY procedure is widely used as an early warning diagnostic of bubble-like behavior. The second stage utilises the new time-varying regression methodology of Greenaway-McGrevy and Phillips (2016) to consider market contagion between these two markets.

Date-stamping the origination and collapse dates of a financial bubble has recently received considerable attention in the literature. Since the development of econometric tools for assessing mildly explosive behavior in a price-fundamental ratio (e.g., Phillips et al. (2011, PWY) and Phillips et al. (2015a, PSY)), these methods have become popular for identifying the presence of financial bubbles as they provide real-time monitoring of financial market exuberance. In addition to PWY and PSY, CUSUM monitoring techniques are an alternative approach for detecting bubbles, see Homm and Breitung (2012). Harvey et al. (2016) propose and develop a wild bootstrap-based implementations of the PWY test, which can reduce spurious indications of explosive behaviour. Phillips and Shi (2017, forthcoming) also provide the latest work of theoretical developments on bubble detection procedures. The PSY procedure is extended from the PWY recursive testing approach. Both the PWY and PSY tests are shown to have significant power in identifying bubbles for a range of financial applications, for example, see Gutierrez (2012) and Etienne et al. (2014) for commodities, Bohl et al. (2013) for stock markets, and Yiu et al. (2013) and Shi et al. (2016) for real estate. Two studies on historical bubble episodes using the PSY approach are Hu and Oxley (2018b) who subject the South Sea and Mississippi Bubbles during 1719–1720 to test for bubbles and contagion whereas Hu and Oxley (2017) find signs of explosive behavior in British railway shares in the 1840s, which are related to the famous British Railway Mania.

We, therefore, first apply the right-tailed unit root test of PSY to examine evidence of explosive behaviour in stock and real estate markets. The popular PSY procedure is based on a test statistic that is designed for examining mildly explosive behaviour in a price-fundamental ratio time series. The PSY is also extended from an early bubble detection test of Phillips et al. (2011). The martingale null with an asymptotic drift is specified as:

$$H_0\colon y_t = dT^{-\eta} + y_{t-1} + \varepsilon_t, \qquad \varepsilon_t \sim \mathrm{NID}(0, \sigma^2), \tag{1}$$

where d is a constant, the localizing coefficient η is greater than $1/2$ and T is the sample size. The alternative hypothesis is a mildly explosive process:

$$H_1\colon y_t = \delta_T y_{t-1} + \varepsilon_t, \tag{2}$$

where $\delta_T = 1 + cT^{-\theta}$ with $c > 0$ and $\theta \in (0, 1)$.

The following regression model is estimated:

$$\Delta y_t = \hat{\alpha} + \hat{\beta} y_{t-1} + \sum_{i=1}^{k} \hat{\gamma}_i \Delta y_{t-i} + \hat{\varepsilon}_t, \tag{3}$$

where $\hat{\alpha}$ is an intercept and k is optimum lag length.

The generalized sup ADF (GSADF) test relies on repeated estimation of the ADF test regression of Eq. (3) on subsamples of the data in a recursive fashion. The window size r_w ($r_w = r_2 - r_1$) expands from r_0 to 1, where r_0 is the minimum window size. The end point r_2 varies from r_0 to 1 and the starting point r_1 varies from 0 to $r_2 - r_0$. The GSADF statistic is the largest ADF statistic over the range of r_1 and r_2:

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

The GSADF test is used for assessing explosive behavior for the entire sample period, however, it does not provide the origination and termination dates of identified bubble episodes. We will use the backward SADF (BSADF) test to detect the origination and termination dates of bubbles. The BSADF statistic is defined as the sup value of the ADF statistic sequence:

$$BSADF_{r_2}(r_0) = \sup_{r_1 \in [0, r_2 - r_0]} ADF_{r_1}^{r_2},$$

where the BSADF statistic and its corresponding critical value are used for dating the origination and termination dates of a bubble. The BSADF test provides more information and improves detective capacity for bubbles within the sample. This approach, therefore, has greater power in the detection of multiple bubbles. Phillips et al. (2015b) also derive the asymptotic distribution of this test statistic under the null. The minimum window size r_0 needs to be large enough to allow initial estimation but not too large to miss an early bubble episode. As recommended in PSY, the minimum window size r_0 is set equal to $0.01 + 1.8/\sqrt{T}$ and a fixed lag order of $k = 0$ is chosen for Eq. (3). Critical values are simulated using 2000 replications.

We then apply the time-varying regression approach of Greenaway-McGrevy and Phillips (2016) to examine evidence of bubble contagion. Greenaway-McGrevy and Phillips (2016) estimate a series of regressions as follows. They first estimate autoregressions of the form in Eq. (3) for each region recursively, generating the slope coefficient estimates $\widehat{\beta}_{i,s}$, where i indexes the geographic region and s indexes the ending date of the subsample ($s = S, \cdots, T$). They then fit the following regression:

$$\widehat{\beta}_{j,s} = \delta_{1i} + \delta_{2j}\left(\frac{S}{T - S + 1}\right)\widehat{\beta}_{core,s-d} + error_s,\ s = S, \cdots, T, \tag{4}$$

for some initialization date S for $j \neq core$, where *core* denotes a core market where the asset bubble is hypothesized to originate. The non-negative delay parameter d that captures the lag in market contagion from the core market to other regions. In empirical analysis, d is an integer.

The primary coefficient δ_{2j} of the functional regression (Eq. (4)) is time-varying. Greenaway-McGrevy and Phillips (2016)'s procedure checks the time series order of the recursively estimated autoregressive coefficients for the core market and market j. Eq. (4) allows the contagion effect to evolve smoothly over time. The time-varying coefficient δ_{2j} captures the contagion effect during the pre- and post-bubble episodes. The phenomenon of bubble migration can be observed for an inverted U shape of δ_{2j} as the contagion effect of the core market to the market j can increase over certain time when two bubbles merge and then decline when bubbles burst. For the details of estimating δ_{2j}, please refer to Greenaway-McGrevy and Phillips (2016).

4. Data

We obtain the house price-rent ratio for Japan between 1970Q1 and 1999Q4 (2010 = 100) from the OECD Main Economic Indicators. A price-rent ratio is a measure of the profitability of owning a house. The rent is commonly used as a proxy for the housing market (e.g., Yiu et al. (2013), Greenaway-McGrevy and Phillips (2016) and Shi et al. (2016)). The 10-year cyclically adjusted price-earnings (CAPE) ratio (also known as *Shiller PE Ratio*) for Japan between 1970M1 and 1999M12 is obtained from the Global Financial Data and is used for assessing explosive behaviour in the stock market.[8] Due to different data frequencies, we create a quarterly price-earnings ratio by averaging and re-scaling to 2010 = 100. Both price-rent ratio and price-earnings ratio are then transformed to natural logarithms. A time series plot for both ratios is provided as Fig. 1. It can be seen from Fig. 1 that both ratios rise dramatically after the mid-1970s, where the price-earnings ratio and house price-rent ratio climb to peaks in 1989 and 1991 and then decline suddenly and sharply.[9]

5. Results

This section reports empirical findings of the bubbly stock and real estate markets in Japan using the PSY method and a time-varying regression approach for analysing market migration. Date-stamping outcomes are provided for both Japan's stock and real estate markets in Fig. 2. It is important to emphasize that the origination (termination) of a bubble episode is defined as the first chronological observation whose the BSADF test statistic (blue solid line) exceeds (goes below) its corresponding critical value (red dash line) if the actual data series (green dash line) grows significantly and then decline suddenly during this episode.[10]

We firstly focus on date-stamping outcomes of the exuberant stock market in Fig. 2 a. The null of no explosive behaviour in Japan's stock price-earnings ratio is strongly rejected at the 1% level as the GSADF statistic (3.7536) exceed the respective 1% right-tail critical value (2.6738), which suggests very strong evidence of bubbles. As shown in Fig. 2 a, we successfully identify a stock market bubble between 1983Q2 and 1990Q3. The timing of such a bubble episode is clearly related to a period of rapid expansion and sudden collapse of Japan's asset price bubble. During this identified bubble episode, the price-earnings ratio grew remarkably in the 1980s and climbed to the peak in 1989, and then declined sharply after the peak.

When we look at Japan's real estate market, it also exhibits bubble-like behaviour as shown by Fig. 2 b. In this case, the null of no explosive behaviour in Japan's house price-rent ratio is also rejected at the 1% level (GSADF statistic: 8.2139 > critical value: 2.6738), indicating the presence of bubbles. As presented as Fig. 2 b, the PSY procedure detects two bubbles in the 1980s (e.g., 1981Q2-1984Q2 and 1987Q2-1992Q2). Of particular interest in the real estate market is that the second identified bubble, which originated in 1987Q2 and collapsed in 1992Q2, is widely believed to be linked with Japan's asset price bubble. During this period in Japan's history, the sudden decline in the real estate market is quite substantial as the price-rent ratio is continuously declining after its peak in 1991 until the end sample of 1999.

A question may arise here as to why we focus on the second bubble rather than the first. First, it is generally accepted that Japan's asset price bubble started from the latter half of the 1980s, see Ito and Iwaisako (1996), Chirinko and Schaller (2001) and Okina et al. (2001). Second, the first bubble doesn't reach the same level of speculation and magnitude as the second bubble if we compare the relevant price-rent ratio for two periods in Fig. 2 b. We, therefore, concentrate on the second identified bubble (1987Q2-1992Q2) rather than the first (1981Q2-1984Q2). It is also worth mentioning that some false positive identification episodes (collapse episodes) are also detected at the end of the sample data. This is due to the fact that the actual data series (house price-rent ratio) is continuously declining from 1994 to 1999 while the BSADF test statistic (blue solid line) crosses the critical value sequences (red dashed line) as shown in Fig. 2 b. Phillips and Shi (2017, forthcoming) recommend users identifying false positive identification episodes ex-post using the PSY procedure, we therefore only report the presence of a bubble (not collapse episodes) in this paper. Therefore it is very important to interpret PSY's results with care to avoid identifying collapse episodes as bubbles.

Overall, we provide evidence of bubbles for two of the most inflated asset prices during Japan's bubble economy based on a recent

[8] Deng et al. (2017) also treat a price-earnings ratio as the price-fundamental ratio when applying the PSY to detect bubbles in China's stock market.

[9] It should be emphasized that both ratios reach its maximum at a height that it never climbs to even almost 30 years later now.

[10] If the data is continuously declining, we would detect some collapse episodes. It is important to distinguish between a bubble and a collapse episode, see Hu and Oxley (2018a).

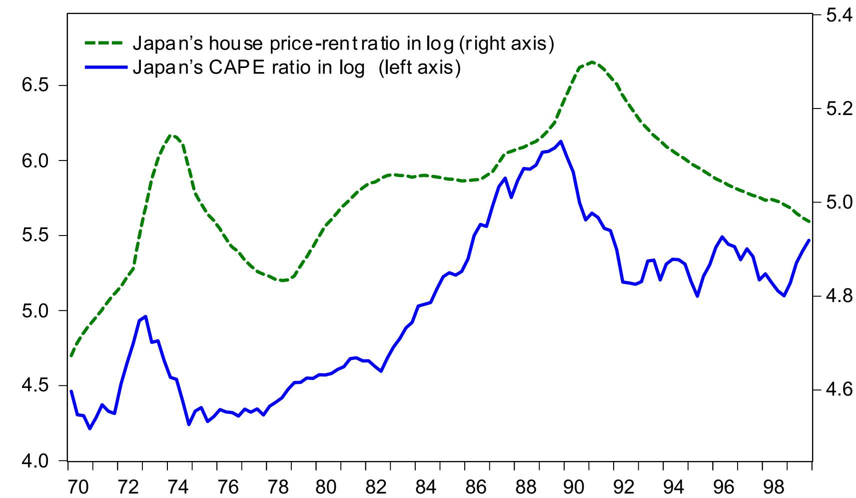

Fig. 1. The time series plot of Japan's house price-rent ratio and cyclically adjusted price-earnings (CAPE) ratio in log scale between 1970Q1 and 1999Q4.

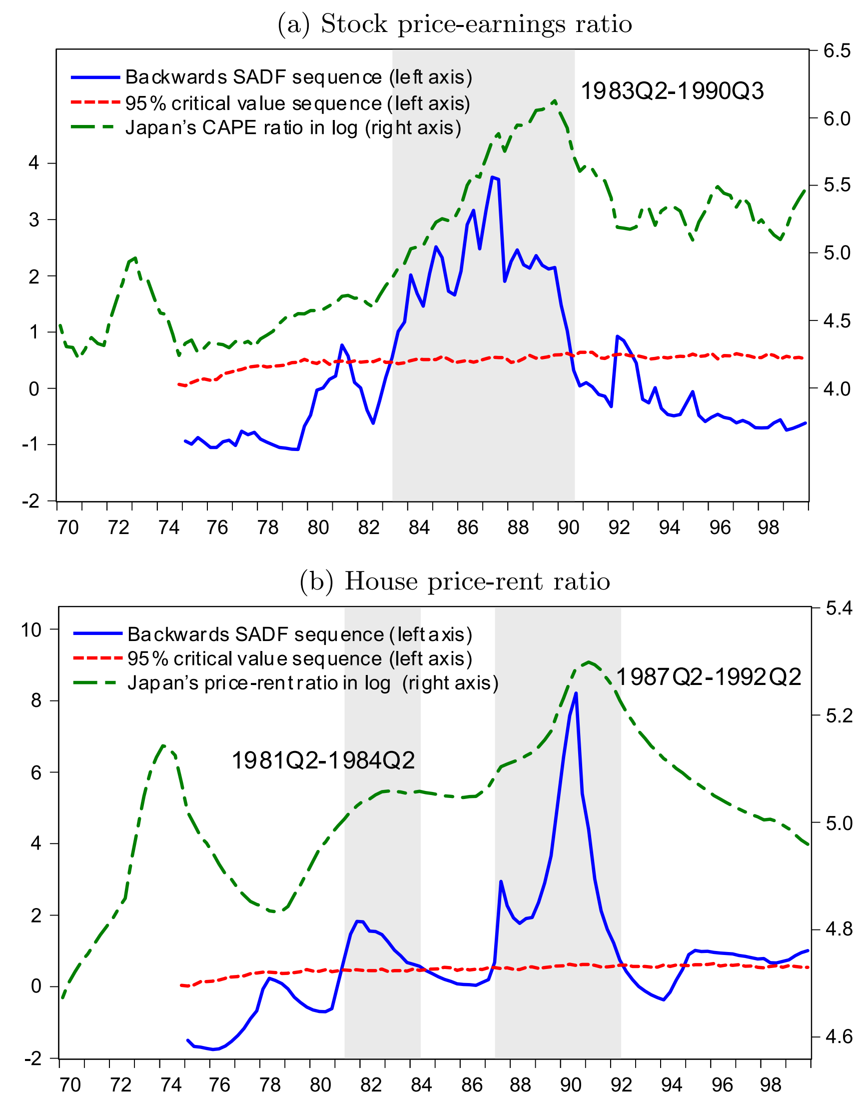

Fig. 2. Date-stamping strategy of cyclically adjusted price-earnings ratio and house price-rent ratio in Japan between 1970Q1 and 1999Q4.(For interpretation of the references to colour in this figure legend, the reader is referred to the web version of this article.)

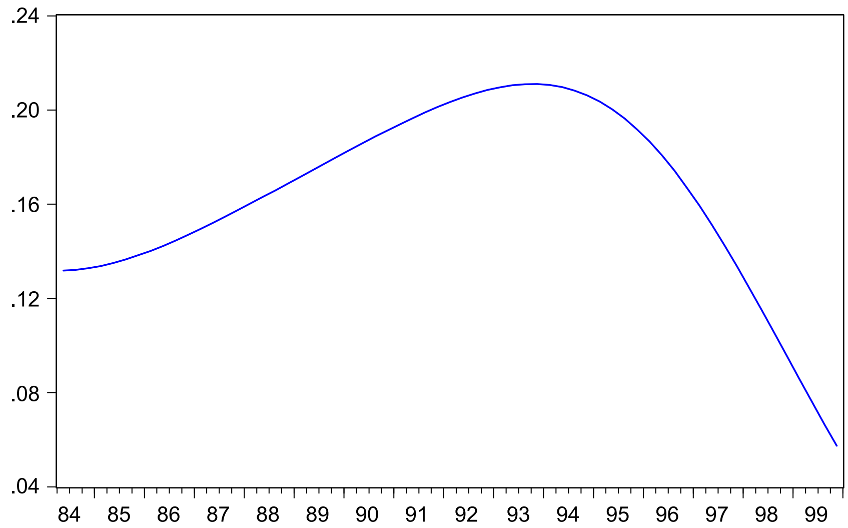

Fig. 3. Time-varying contagion coefficient δ_{2j} from the Japanese stock market to the real estate market.

econometric test of PSY.[11] For both markets, we identify the origination and collapse dates of so-called 'double bubbles', which coincide with the traditional view of the well-known Japan's asset price bubble. We can clearly see that a bubble emerged in the stock market in 1983Q2 before spreading to its real estate market. During the 1980s, our results seem to suggest that the bubble-like behaviour in stock market migrated to the real estate market, creating a bubble in 1987Q2 that collapses in 1992Q2. However, does rigourous testing confirm these 'eyeball based tests'- we next consider the application of the time-varying regression approach of Greenaway-McGrevy and Phillips (2016).

The stock market is regarded as the core market hence we believe that the bubble migrates from the stock market to the real estate market for four reasons.[12] First, Japan's stock market seems to hit a historical market high earlier than the real estate market as the stock price-earnings ratio reached its peak in 1989 while the price-rent ratio climbed to its maximum in 1991. Second, the stock market exhibits bubble-like behaviour earlier than the housing market during the late 1980s, which is clearly shown by comparing Fig. 2 a and Fig. 2 b. Third, the stock price bubble bursts earlier than real estate bubble. In particular, we can see from Fig. 1 that the stock price-earnings ratio started to decline after 1989Q4 while the house price-rent ratio started to fall after 1991Q1. Fourth, transaction costs are much lower in the stock market as people can simply and quickly sell off their shares and then invest in the housing market. Hence it is reasonable to assume Japan's stock market as a core market, and a similar view is also noticed in the literature. Voth (2003) points out that Japan's asset bubble episode migrates from the stock market into the real estate market in the 1980s.[13] Basile and Joyce (2001) also show that changes in stock price bubble lead to changes in the land market bubble during the bubble period.

Using the time-varying regression approach of Greenaway-McGrevy and Phillips (2016), we let the fixed window size S be 50 and delay parameter d be 4. A fixed window subsample method is utilized for selecting the subsample sequence of estimated autoregressive coefficients as suggested by Greenaway-McGrevy and Phillips (2016). An important feature of this econometric method is that the primary coefficient δ_{2j} is time-varying in the proposed regression Eq. (4), indicating contagion effects from the core market to the other market vary smoothly over time.

If a bubble in the stock market migrates to the real estate market, we would expect to see a particular pattern of contagion effects i.e., an inverted U shape of contagion coefficients δ_{2j} due to the intensity of transmission increasing when two bubbles merge, and then decreases when bubbles burst. This is the reason why the so-called inverted U-shape of contagion coefficients δ_{2j} demonstrates the presence of bubble migration as discussed in the method section. In our example, the sensitivity of Japan's real estate market to its stock market is clearly shown in Fig. 3 and it clearly exhibits an inverted U shape as expected.[14] The sensitivity rises gradually and reaches a peak in 1993 (slightly after the peak in Japan's real estate), then declines suddenly, as the response to the core market gets smaller and smaller. Fig. 3 clearly demonstrates signs of bubble migration from the stock market (core market) to the real estate market due to an inverted U shape of the contagion effect, which is a major finding that contributes to the literature. A finding of a bubble migration phenomenon in this paper supports the traditional notions of Japan's asset price bubbles. Our results help to understand why Japan's real estate bubble collapse after Japan's stock market bubble as the stock price bubble migrates to the real estate market.

6. Conclusions

Japan's asset price bubble during the 1980-90s was one of the most significant financial bubbles in history. During the speculative growth of bubble period, Japan's asset prices rose without precedent (e.g., stocks, land, real estate, golf course memberships). With the additional evidence presented here, there remains no doubts that these asset prices were a bubble. However, in the later half of the 1980s, the rise of asset prices was thought to be associated with economic fundamentals rather than irrational exuberance in stock and real estate markets.

This current paper makes two significant contributions to the

[11] In addition, we also apply the PSY procedure to Japan's CAPE and house price-rent ratios to take into account the presence of possible heteroscedasticity using wild bootstrapping of Harvey et al. (2016). We identify a bubble in the stock market (1980Q4-1995Q2) and two bubbles in the real estate market (1981Q1-1986Q4 and 1987Q3-1993Q1). The conclusion of bubbles in both markets still holds. Results are not shown here to conserve the space, but are available upon request.

[12] Deng et al. (2017) also treat the stock market as the core market.

[13] On page 84 of Voth (2003), he notices that 'most other major asset bubble episodes saw "spillovers" from the stock market into the real estate market-such as in Britain and Japan in the 1980s'.

[14] We also try different choices of the fixed window size S and delay parameter d to adjust the sampling period. The results remain unchanged as a similar inverted U shape of the contagion effect is also obtained.

literature. First, it presents empirical evidence to support the existence of an asset price bubble in Japan in the late 20th century using the right-tailed unit root test of PSY as we document overwhelming evidence in both stock and real estate markets between the 1980s and the early 1990s. Second and more importantly, we then apply the Greenaway-McGrevy and Phillips (2016)'s methodology to detect possible contagion between the two markets. Japan's stock market is treated as the core for several reasons to allow the contagion effect to transmit to real estate market. It is expected to see that the contagion effect of the core market on the other market exhibits an inverted U shape, in which the contagion effect increases when two bubbles merge and declines when two bubbles burst. Our results clearly show signs of spillovers from the Japan's stock market to its real estate market for the first time in the literature due to an inverted U shape of contagion effect as shown in Fig. 3. The results we present contribute to the small, yet growing empirical literature that combines PSY type tests with Greenaway-McGrevy and Phillips (2016)'s time-varying regression approach and provides the final set of empirical results to rigourously confirm contagious, spillovers across asset and real estate markets in Japan. In terms of potentially fruitful areas for future research which may offer some new, unique empirical findings, one could consider investigating the presence of housing bubbles in Japan's regional and metropolitan real estate market. In particular evidence of contagion from a core area or market to other markets and/or areas may be localised or more widespread. This may offer valuable insights in our understanding of how Japan's real estate bubble developed and spread, spatially, during the 1980s and 90s. This new knowledge could be invaluable for stemming the development and subsequent migration of bubble effects in Japanese markets in the future.

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Supplementary material

Supplementary material associated with this article can be found, in the online version, at 10.1016/j.jjie.2018.09.002

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Chapter 8

Do US House Prices Reflect Economic Fundamentals? Evidence from Quantile Cointegration Tests.

Abstract

In this chapter, we re-visit the long-run relationship between house prices and economic fundamentals using the US data for the period 1978M1-2016M12. Unlike the existing studies that employ conventional cointegration tests, this chapter applies two recent quantile ARDL models of Cho et al. (2015) and Greenwood-Nimmo et al. (2011) to explore the equilibrium relationship at the national and state levels. The empirical results from this paper also suggest that the cointegrating relationship may be absent for conventional cointegration tests, but it may be established by quantile based models, and vice versa.

Keyword

Quantile cointegration; ARDL; US house prices

JEL classifications: *C59; R19; R39*

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8.1 Introduction

Possible cointegration between house prices and fundamentals is important for understanding house price dynamics. The presence of cointegration between house prices and fundamentals would imply convergence to a stable long-run equilibrium relationship. In other words, cointegration would indicate whether a temporary increase or drop in house prices would eventually come back to an equilibrium. Among many housing markets, US housing market has drawn a lot of attention in the literature. Glaeser (2013) concludes that the United States seems to have a long history of real estate speculation as the nation experienced several periods of speculation. This chapter, therefore, aims to explore the existence of a stable long-run relationship between house prices and economic fundamentals using US aggregate and disaggregate data, by utilising quantile cointegration approaches.

An important question that has been researched extensively in the literature is whether US house prices reflect their economic fundamentals. There have been many studies that attempted to answer this question. However, the existing studies have mainly focused on cointegration analysis do not reach a firm conclusion. This chapter, therefore, tries to address this issue by applying several autoregressive distributed lag (ARDL) based conditional mean- and quantile cointegration tests. Two conditional mean cointegration models are considered in this study including the linear ARDL cointegration model of Pesaran and Shin (1998) and Pesaran et al. (2001) and the non-linear ARDL version of Shin et al. (2014). These two conditional mean-based models are abbreviated as ARDL-M and NARDL-M, respectively. The ARDL model of Pesaran and Shin (1998) and Pesaran et al. (2001) have become popular in analysing a long-run cointegrating relationship.

Shin et al. (2014)'s model is an asymmetric extension of the linear ARDL approach to modelling long-run cointegrating relationship based on partial sum decompositions suggested by Schorderet (2003). Not only does this paper explore a long-run cointegrating relationship at the conditional mean, but also investigates the relationship for the whole conditional distribution. Two quantile ARDL cointegration models of Cho et al. (2015) and Greenwood-Nimmo et al. (2011) are also employed in this study. These two quantile-based models are abbreviated as NARDL-Q and QARDL, respectively. Apart from the mean, other locations in the distribution affect the cointegration analysis, which is well demonstrated by several empirical studies in the literature especially since the work of Xiao (2009), for example, see Lee and Zeng (2011), Burdekin and Siklos (2012) and Tsong and Lee (2013). Conventional cointegration approaches focus on the conditional mean behaviour, which is perhaps not informative (Cho et al., 2015). As noted in Chevapatrakul et al. (2009), conventional estimation methods such as LS, IV, or GMM evaluate the relationship between these observations at the mean of the conditional distribution. It is assumed that the estimated relationship holds at the mean and even in the tails. Xiao (2009) further points out that the absence of cointegration has frequently been observed in many applications using conventional cointegration despite the fact that economic variables are supposed to be cointegrated. Xiao (2009) explains that this phenomenon may be explained by the existence of time-varying cointegrating coefficients, which characterise the long-run relationship but vary over the time. Conventional cointegration tests focus on the conditional mean relationship between variables while quantile cointegrating models concentrate on a long-run relationship in a range of quantiles. The quantile-based cointegration tests are potentially more informative as they provide additional insights.

Several ARDL based cointegration tests are employed in this chapter to model a long-run equilibrium relationship. There have been increasing interests in modelling economic phenomenon by non-linear models. Three regime-switching models have been popularised in the literature including the threshold ECM of Balke and Fomby (1997), the Markov-switching ECM of Psaradakis et al. (2004) and the smooth transition regression ECM of Kapetanios et al. (2006). It is known that linear models do not always provide optimal inference. The majority of studies on cointegration analysis rely on the assumption that the long-run relationship may be represented as a symmetric linear combination of nonstationary stochastic regressors. However, such an assumption may be too restrictive.

There has been little research on non-linear cointegration with a few exceptions. Two approaches are worth special mention to modelling non-linear cointegration using similar ideas. The first approach is asymmetric cointegration by Schorderet (2003), who introduces a nonlinear version of a cointegration test by decomposing a time series into its positive and negative partial sums. This new generalisation of cointegration is appealing to model asymmetric behaviour. Schorderet (2003) applies this technique to the European exchange rate data yielding evidence of asymmetric cointegration. In contrast, Johansen's (1991) maximum likelihood procedure shows no linear cointegration. This asymmetric cointegration approach has been applied to several empirical applications. For example, Lardic and Mignon (2006) and Lardic and Mignon (2008) investigates the existence of a long-term relationship between oil prices and GDP in 12 European countries, and G7 countries and the US, respectively. The second approach is the hidden cointegration of Granger and Yoon (2002). Granger and Yoon's (2002) hidden cointegration decomposes the positive and negative components of each time series and tests for cointegration in these components. If the components of two data series (negative or positive) are cointegrated, then the data are said to have a hidden cointegration. Hidden cointegration is a type of nonlinear cointegration. Granger and Yoon (2002) demonstrate the usefulness of hidden cointegration using two sets of US data- short-term and long-term interest rates, output and unemployment. The hidden cointegration approach has also been applied in other area, for instance, Richard (2012) examines the energy-growth nexus in sub-Saharan Africa countries using annual data from 1971 to 2008. Honarvar (2009) studies asymmetry in crude oil and retail gasoline price movements for the period

1981M09-2007M12. Tiwari et al. (2015) investigate the hypothesis of asymmetric effects between economic growth and renewable and nonrenewable energy production in sub-Saharan Africa using the annual data 1971-2011. Alexakis et al. (2017) explore a long-run relationship between Islamic and conventional equity indices based on daily data from 2000M03 to 2014M06.

Apart from cointegration tests of Pesaran et al. (2001) and Cho et al. (2015), this chapter will investigate whether US house prices cointegrate with several economic fundamentals using two non-linear cointegration models-the NARDL model of Shin et al. (2014) and the quantile NARDL model of Greenwood-Nimmo et al. (2011). To the best knowledge, this study is also the first to attempt to examine a long-run quantile-varying cointegration relationship between US house prices and economic fundamentals using the development of linear (symmetric) and non-linear (asymmetric) quantile ARDL cointegration models of Cho et al. (2015) and Greenwood-Nimmo et al. (2011). By applying the linear and non-linear quantile autoregressive distributed-lag models of Cho et al. (2015) and Greenwood-Nimmo et al. (2011), quantile cointegration models allow us to study the entire conditional distribution of the dependent variable compared with those of conditional mean-based cointegration models. As the housing market is our core focus, we may interpret the conditional distribution of housing market with lower quantiles (at a relatively low price level) as a bear market, the median quantiles as the normal phase of the market, and the upper quantiles (at a relatively high price level) as a bull market. A similar interpretation can be found in Christou et al. (2018).

This chapter makes several contributions to the existing literature. The first contribution is that formal cointegration testing procedures for two quantile ARDL models of Cho et al. (2015) and Greenwood-Nimmo et al. (2011) are utilised. By applying two quantile cointegration models, this research is the first to explore US house price dynamics. Cho et al. (2015) extends Pesaran and Shin's (1998) autoregressive distributed-lag approach into quantile regression, resulting a new quantile autoregressive distributed-lag (QARDL) model. Similarly, Greenwood-

Nimmo et al. (2011) also develop a non-linear autoregressive distributed-lag quantile (NARDL-Q) regression model, which is extended from the non-linear autoregressive distributed lag (NARDL) model by Shin et al. (2014). However, a formal cointegration testing mechanism that could address the existence of the long run cointegration relationship at each quantile is provided in neither Cho et al. (2015) nor Greenwood-Nimmo et al. (2011). This paper makes use of the t-test and Wald test to jointly test for quantile cointegration at each quantile. The first testing procedure of Banerjee et al. (1998) is to carry out a t-test. The second testing procedure is a Wald test, which is used to test for quantile cointegration based on asymptotic theory. Existing studies have not adopted formal cointegration testing procedures using the QARDL model, see Zhu et al. (2016) and Lahiani et al. (2017). This research formally tests for the long run cointegration relationship at each of quantiles by two formal testing procedures of the two linear and non-linear quantile ARDL based models for the first time in the literature.

The second contribution of this chapter is to make use two forms of non-linear (asymmetric) cointegration tests at the conditional mean and conditional quantile applied to the US housing market. Prior empirical studies in the literature mainly employ the linear model based cointegration tests or panel version tests to test for a long-run cointegrating relationship. Little research has focused on non-linear (asymmetric) versions of cointegration models excluding Zhou (2010) and Bahmani-Oskooee and Ghodsi (2016). Adoption of the NARDL-Q model of Greenwood-Nimmo et al. (2011) makes it empirically possible to examine non-linear (asymmetric) quantile cointegration at the conditional quantile for the first time. Moreover, none of the existing studies has adopted quantile cointegrating models to investigate the existence of a long-run equilibrium relationship between house prices and fundamentals. Therefore this chapter certainly contributes to this literature by applying two new quantile cointegrating models to explore a long-run cointegrating relationship between house prices and fundamentals across the whole conditional distribution.

The third contribution of this paper is that it presents several interesting findings based on empirical evidence. The US national house prices are not cointegrated with economic fundamentals using both conditional mean- and quantile ARDL based models. The empirical results from this paper also suggest that the cointegration relationship may be absent at the conditional mean-based model, but it may be established at the quantile based model. Under some instances, cointegration may be established based on the conditional mean model, but there is no overwhelming evidence of cointegration based on its quantile model. Additionally, it is also possible to identify a cointegration relationship at the conditional mean and the relationship may be held for the whole conditional distribution by quantile models.

The following paper is organized as follows. Section 8.2 reviews the existing studies in the literature. Section 8.3 describes the data. Section 8.4 gives a brief description of several autoregressive distributed lag (ARDL) based models. Section 8.5 presents the empirical findings and Section 8.6 concludes.

8.2 Literature Review

8.2.1 A review of quantile cointegration models

This section reviews several quantile cointegrating models in the literature including Xiao (2009), Kuriyama (2016), Cho et al. (2015) and Greenwood-Nimmo et al. (2011). Xiao's (2009) work has made a significant contribution to the quantile regression with non-stationary variables. Xiao (2009) advances a quantile cointegration approach based on the estimators proposed by Phillips and Hansen (1990) and Saikkonen (1991). A Phillips-Hansen type quantile estimator is used for removing serial correlation and long-run endogeneity.

Xiao (2009) considers quantile regression of the following cointegration model:

$$y_t = \alpha + \beta' x_t + u_t = \theta' z_t + u_t, \tag{8.1}$$

where x_t is a k-dimensional vector of integrated regressors, $z_t = (1, x_t')'$, and u_t is mean zero stationary.

To extend the model of Equation (8.1), Xiao (2009) applies the idea proposed by Saikkonen (1991) that decomposes u_t into lead-lag terms $\sum_{j=-K}^{K} \Delta x_{t-j}$ and a pure innovation component ε_t to deal with endogeneity in traditional cointegration models and also model the time varying cointegrating coefficient β_t as a function of the pure innovation component. The following model is extended from Equation (8.1):

$$y_t = \alpha + \beta_t' x_t + \sum_{j=-K}^{K} \Delta x_{t-j}' \Pi_j + \varepsilon_t \tag{8.2}$$

If we denote the τ-th quantile of ε_t, as $Q_\varepsilon(\tau)$, let $\mathcal{F}_t = \sigma\{x_t, \Delta x_{t-j}, \forall_j\}$, then, conditional on $\mathcal{F}_t$, the τ-th quantile of y_t is given by:

$$Q_{yt}(\tau|\mathcal{F}_t) = \alpha + \beta(\tau)' x_t + \sum_{j=-K}^{K} \Delta x_{t-j}' \Pi_j + F_\varepsilon^{-1}(\tau), \tag{8.3}$$

where $F_\varepsilon(\cdot)$ is the c.d.f of ε_t. Let Z_t be the vector if regressors consisting $z_t = (1, x_t')$ and $(\Delta x_{t-j}', j = -K, ..., K)$, $\Theta = (\alpha, \beta_t', \Pi_{-K}', ..., \Pi_K')'$, $\Theta(\tau) = (\alpha(\tau), \beta(\tau)', \Pi_{-K}', ..., \Pi_K')'$, where $\alpha(\tau) = \alpha + F_\varepsilon^{-1}(\tau)$, then, we can re-write the above regression as:

$$y_t = \Theta' Z_t + \varepsilon_t = \Theta(\tau)' Z_t + \varepsilon_{t\tau}, \tag{8.4}$$

$$Q_{yt}(\tau|\mathcal{F}_t) = \Theta(\tau)' Z_t. \tag{8.5}$$

If $\varepsilon_{t\tau} = \varepsilon_t - F_\varepsilon^{-1}(\tau)$, then $Q_{\varepsilon_{t\tau}}(\tau) = 0$. In the above model, the value of the cointegration coefficients are affected by the innovation received at each period. The cointegrating vector may be quantile dependent. This model can be referred to as the "quantile cointegration model". Xiao's (2009) model can be treated as a stochastic cointegration model which includes conventional cointegration as a special case where $\beta(\tau)$ is a vector of constants. This special case can be presented as:

$$y_t = \alpha + \beta' x_t + \sum_{j=-K}^{K} \Delta x_{t-j}' \Pi_j + \varepsilon_t, \tag{8.6}$$

and

$$Q_{yt}(\tau|x_t) = \alpha + \beta' x_t + \sum_{j=-K}^{K} \Delta x'_{t-j}\Pi_j + F_{\varepsilon}^{-1}(\tau). \tag{8.7}$$

In Xiao's (2009) application to US S&P stock price and dividend data from January 1974 and September 1998, the null hypothesis of constant cointegrating coefficients is rejected at the 1% level, displaying strong evidence of time-varying cointegrating coefficient behaviour.

Xiao's (2009) approach has gained popularity in recent years. As shown in Table 8.1 and Table 8.2, below this approach has been applied to many interesting applications including the relationship between spot and futures oil prices; stock market integration; the validity of the Fisher hypothesis for six OECD countries; the relationship between globalization and economic performance; the relationship between government expenditures and revenues in US; the classic PPP hypothesis in China, Japan and South Korea and the relationship between spot and futures prices of gold and silver.

More recently, Kuriyama (2016) suggests a cumulated sum (CUSUM) test for the null hypothesis of quantile cointegration. Kuriyama (2016) extends the analysis of Xiao and Phillips (2002) to the case of conditional quantiles. A residual-based test for the null of cointegration is proposed in Xiao and Phillips (2002) using a conventional cumulated sum (CUSUM) test for structural change. The linear model is considered in Xiao and Phillips (2002) as follows:

$$y_t = \alpha' d_t + \beta' x_t + u_t = \theta' z_t + u_t, \tag{8.8}$$

where $\theta = (\alpha', \beta')'$, $z_t = (d'_t, x'_t)'$ and d_t is a vector of deterministic trend component. A CUSUM statistic is constructed based on the fully modified OLS residuals to test cointegration between y_t and x_t.

The highlight of Kuriyama's (2016) approach is that it allows the cointegrating coefficient to change across the whole conditional distribution of the dependent variable and simultaneously test for the null of cointegration at each quantile.

Kuriyama (2016) introduces the quantile version of Equation (8.8) as:

$$y_t = \alpha'(\tau)d_t + \beta'(\tau)x_t + u_t(\tau) = \theta'(\tau)z_t + u_t(\tau), \tag{8.9}$$

where $\theta(\tau) = (\alpha'(\tau), \beta'(\tau))'$. The quantile dependent regression coefficient vector $\theta(\tau)$ can characterize the possibly non-linear long-run relationship between y_t and x_t.

Kuriyama (2016) uses a Phillips-Hansen type fully modified estimator to correct serial correlation and endogeneity. The fully modified estimator $\hat{\beta}^+(\tau)$ of $\beta(\tau)$ is given by:

$$\hat{\beta}^+(\tau) = \hat{\beta}(\tau) - \left[f(\widehat{F^{-1}(\tau)}) \sum_{t=1}^{T} \underline{x}_t^d \underline{x}_t'^d \right]^{-1} \left[\sum_{t=1}^{T} \underline{x}_t^d \hat{\Omega}_{\psi x} \hat{\Omega}_{xx}^{-1} \Delta x_t + T\hat{\Delta}_{x\psi}^+ \right], \tag{8.10}$$

where $\underline{x}_t'^d$ denotes the demeaned or detrended regressors, and $f(\widehat{F^{-1}(\tau)})$ is a non-parametric consistent estimator of the density function $f(F^{-1}(\tau))$.

For the quantile regression model, the residuals from the fully modified quantile regression are calculated as $\hat{u}_t^+(\tau) = y_t^+ - z_t'\hat{\theta}^+(\tau)$. The cumulated sum of $\psi_\tau(\hat{u}_t(\tau))$ meets the following conditions:

$$\begin{aligned} \max_{n=1,\ldots,T} T^{-\frac{1}{2}} \left| \sum_{t=1}^{n} \psi_\tau(\hat{u}_t(\tau)) \right| &= O_p(1) \qquad \text{under } H_0, \\ &= O_p(T) \qquad \text{under } H_1. \end{aligned} \tag{8.11}$$

The CUSUM test statistic for the null hypothesis of quantile cointegration is given by:

$$CS_T(\tau) = \max_{n=1,\ldots,T} \frac{1}{\hat{\omega}_{\psi x}\sqrt{T}} \left| \sum_{t=1}^{n} \psi_\tau(\hat{u}_t^+(\tau)) \right|, \tag{8.12}$$

This new approach has been applied to several sets of US interest rates. In particular, the expectation hypothesis for the term structure is retained only in part of the distribution for certain data sets. In addition, Christou et al. (2018) applies Kuriyama's (2016) approach to examine a quantile-varying long-run relationship between US house prices and non-housing Consumer Price Index from 1953 to 2016, which both variables are cointegrated at lower quantiles only.

Cho et al. (2015) extends Pesaran and Shin's (1998) autoregressive distributed-lag approach into a quantile regression context, which allows us to jointly address short-run dynamics and long-run cointegrating relationships across a range of quantiles of the conditional distribution of the dependent variable. The resulting QARDL framework has been applied to analyse US dividends and earnings for the period 1871Q3-2010Q2. There has been growing interest in using the QARDL model. Zhu et al. (2016) employ the QARDL model to explore quantile cointegration relationship between silver and gold prices and their results suggest that the prices of silver and gold are cointegrated in tail quantiles but not in the middle quantiles. This shows the cointegration holds when the silver price is at a relatively low price level (lower quantiles) or relatively high price level (upper quantiles)-conditional on gold prices. Lahiani et al. (2017) also apply the QARDL model to investigate the short- and long-run linkages between oil prices and a set of energy prices (e.g., gasoline, diesel, heating and natural gas). They find that all considered energy prices are shown to be cointegrated with the oil price across quantiles and the oil price is a predictor of individual energy prices in the short run.

Shin et al. (2014) propose a non-linear autoregressive distributed lag (NARDL) model to model short- and long-run asymmetries jointly based on partial sum decompositions suggested by Schorderet (2003). Based on this new development, Greenwood-Nimmo et al. (2011) further develop a non-linear autoregressive distributed-lag quantile (NARDL-Q) regression model, which is capable of simultaneously identifying three fundamental forms of asymmetry: long-run (reaction), short-run (adjustment) and quantile-specific (location) asymmetry. A set of hypotheses are proposed regarding testing for each type of asymmetry. In their application to US monetary policy, the NARDL-Q model provides more insights compared with the conditional mean-based model.

Table 8.1 – Empirical studies using quantile cointegration models of Xiao (2009).

Authors	Sample data	Method	Finding
Lee and Zeng (2011)	The daily spot and futures oil prices of West Taxas Intermediate (WTI) from 2 January 1986 to 6 July 2009.	Xiao (2009)	Significant evidence of non-linear cointegration.
Burdekin and Siklos (2012)	The daily data for 12 stock market indices in the Asia-Pacific region as well as the S&P 500 from 4 January 1995 to 15 July 2010.	Xiao (2009)	Cointegration is established between Shanghai market, the US market and many regional exchanges.
Tsong and Lee (2013)	The 3-month Treasury-Bill rate and the annual growth rate of CPI from 1957Q1 to 2010Q2.	Xiao (2009)	The Fisher hypothesis is held in Australia, Belgium, Canada, Sweden, the UK and the US as two. variables are cointegrated in a quantile sense.
Chang et al. (2015)	The annual real GDP and indices of globalization of G7 countries from 1970 to 2006.	Xiao (2009)	The long-run relationship is established in most models.

Table 8.2 – Empirical studies using quantile cointegration models of Xiao (2009).

Authors	Sample data	Method	Finding
Chen (2016)	The US nominal government spending and government revenues or per-capital basis between 1960Q2 and 2010Q3.	Xiao (2009)	A quantile-dependent cointegrating relationship exists.
Ma et al. (2017)	The nominal interest rate and price differential for Japan, China and Korea from as early as 1974M1 to 2015M08.	Xiao (2009) Koenker and Xiao (2004)	Mixed evidence to support PPP theory.
Schweikert (2018)	The gold and silver spot prices and the corresponding futures prices. from 1971 to 2017.	Xiao (2009)	The long-run relationship is established.

Table 8.3 – Empirical studies using quantile cointegration models of Cho et al. (2015) and Kuriyama (2016).

Authors	Sample data	Method	Finding
Zhu et al. (2016)	The weekly prices of gold and silver from 3 April 1968 to 27 April 2016.	QARDL model of Cho et al. (2015)	Cointegration is due to the tail quantiles.
Lahiani et al. (2017)	The daily US spot closing prices for WTI crude oil, gasoline, diesel, heating, and Henry Hub natural gas.	QARDL model of Cho et al. (2015)	Energy prices are cointegrated with oil prices.
Christou et al. (2018)	The monthly US house prices and non-housing Consumer Price Index (CPI) from 1953 to 2016.	Kuriyama (2016)	House prices are cointegrated with non-housing Consumer Price Index at lower quantiles between 5% and 20%.

8.2.2 A Review of Cointegration Relationship between US House Prices and Fundamentals

Table 8.4, Table 8.5, Table 8.6 and Table 8.7 present results from the previous literature considering the existence of a cointegration relationship between US house prices and fundamentals. Those studies typically conclude that cointegration exists and have a long-run equilibrium likewise from then they adopt and estimate an error-correction model see, for example, Abraham and Hendershott (1996), Malpezzi (1999), Meen (2002) and Gallin (2008).

However, Gallin (2006), Mikhed and Zemčík (2009a), Mikhed and Zemčík (2009b) and Clark and Coggin (2011) provide conflicting results on a long-run cointegrating relationship. Gallin (2006) finds no evidence of cointegration between house prices and fundamentals at the national level between 1975Q1 and 2002Q2. More importantly, Gallin (2006) concludes no evidence of cointegration in a panel of 95 metro areas from 1978 to 2000 using the panel test of Maddala and Wu (1999) and Pedroni (1999). Both Mikhed and Zemčík (2009a) and Mikhed and Zemčík (2009b) apply the panel test of Pedroni (1999) for cointegration between house prices and various economic fundamentals, concluding no evidence to support a long-run cointegrating relationship. Moreover, Clark and Coggin (2011) also present no evidence of cointegration from 1975Q1 to 2005Q2 at the national and regional levels using the Engle and Granger (1987) and Gregory and Hansen (1996) approaches.

On the other hand, several studies show that house prices and fundamentals are cointegrated, for example, McCarthy and Peach (2004), Gallin (2008), Duca et al. (2011), Arestis and González (2014) and Chen and Cheng (2017). Zhou (2010) and Bahmani-Oskooee and Ghodsi (2016) argue that ignoring non-linear cointegration may lead to no cointegrating relationship. For example, Zhou (2010) shows that the non-linear cointegration approach of Granger and Hallman (1991) and Granger (1991) presents overwhelming evidence of cointegration at the US city level. A similar conclusion can be drawn from Bahmani-Oskooee and Ghodsi

(2016) as the non-linear ARLD model of Shin et al. (2014) shows more evidence of cointegration in the US states. These two studies highlight the importance of non-linear cointegration in modelling the long-run relationship.

8.3 Data

As the US regional housing market could behave differently at the aggregated national market, the focus of this research is to explore the possibility of a long-run cointegrating relationship between house prices and fundamentals at the both national (aggregate) and State (disaggregate) levels. The eight regional states and District of Columbia (DC) are selected for this study including Arizona (AZ), California (CA), Florida (FL), Hawaii (HI), Massachusetts (MA), Nevada (NV), New York (NY) and Pennsylvania (PA).[1] In summary, these selected States have potential to exhibit bull & bear characteristics of Christou et al. (2018) at lower and upper quantiles. If these States do not support quantile cointegration, it is unlikely that other States would.

These states are located in different parts of the US and are selected for their distinguishing features. For example, California is the most populous state and the largest economy in the United States. Notably, the San Francisco Bay Area and the Greater Los Angeles Area in California are famous as centres of the global technology and entertainment (firm) industries, respectively. The District of Columbia is the capital of the US. Florida has the longest coastline in the United States. Florida has the nickname of "Sunshine State" as it is well-known for its balmy weather in the world and preference for retirement. Hawaii is the only U.S. state located in Oceania and is a world-renowned vacation spot. Massachusetts is a true hub of higher education. The Medicine and the life science industry are also

[1] We focus on regional data at the state level rather than city level as economic variables (personal income, construction costs and the corresponding CPI) are more easily obtained at the monthly or quarterly frequency without resorting to interpolation. When an economic variable is available at the monthly frequency, a quarterly variable is created by averaging.

Table 8.4 – Cointegration between US house prices and economic fundamentals.

Authors	Sample data and variables	Method	Finding
Abraham and Hendershott (1996)	Regional: 1977-1992 for 30 cities Variables: house prices, construction costs, disposable income, interest rate.	Error correction model	Cointegration
Malpezzi (1999)	Regional: 1979-1996 for 133 MSAs Variables: house prices, income, population, geographical variables, interest rates and regulation.	Error correction model	Cointegration
Meen (2002)	Regional: 1977-1992 Variables: house prices, income, rent, construction costs, population, mortgage rates and etc.	Error correction model	Cointegration
Gallin (2006)	National: 1975Q1-2002Q2 Regional: 1978Q1-2000Q4 for 95 cities Variables: house prices, per capital income, user cost, stock market, construction wage.	Engle and Granger (1987) Panel test of Maddala and Wu (1999) and Pedroni (1999)	No cointegration

Table 8.5 – Cointegration between US house prices and economic fundamentals.

Authors	Sample data and variables	Method	Finding
McCarthy and Peach (2004)	National: 1981Q1-2003Q3 Variables: house prices, housing stock, and user cost.	Johansen (1988)	Cointegration
Gallin (2008)	Regional: 1970Q1-2001Q4/2005Q4 Variables: house prices, rent, rent-price ratio, mortgage rate, property tax rate, marginal income tax rate.	Error correction model	Cointegration holds 1970Q1-2001Q4 only.
Mikhed and Zemčík (2009a)	National: 1980Q2-2008Q2	Engle and Granger (1987)	No cointegration
	Regional: 1978-2007 for 22 MSAs. Variables: house prices, rent, construction cost, personal income, population, mortgage rate, stock market.	Panel test of Pedroni (1999), Pedroni (2004)	No cointegration
Mikhed and Zemčík (2009b)	Regional: 1978H1-2006H2 for 23 MSAs Variables: house prices, rent.	Panel test of Pedroni (1999, 2004)	No cointegration

Table 8.6 – Cointegration between US house prices and economic fundamentals.

Authors	Sample data and variables	Method	Finding
Zhou (2010)	National: 1978Q1-2007Q4	Engle and Granger (1987), Johansen (1991, 1995),	Mixed.
	Regional: 1978Q1-2007Q4	Granger and Hallman (1991), Granger (1991)	Linear cointegration: Cleveland.
	for 10 cities Variables: house prices, personal income, construction costs, mortgage rates.		Non-linear cointegration: the national, Chicago, Dallas, Philadelphia, Richmond, Seattle, St. Louis. No cointegration: Boston, Los Angeles, New York.
Clark and Coggin (2011)	National: 1975Q1-2005Q2	Engle and Granger (1987), Gregory and Hansen (1996)	No cointegration
	Regional: 1975Q1-2005Q2 for 4 regions Variables: house prices, mortgage rates, unemployment rate, homeowner rate, stock index, price-rent ratio and etc.	Qu (2007)	
Duca et al. (2011)	National: 1981Q1-2007Q2 Variables: price-rent ratio, user costs, loan-value ratio.	Johansen (1991, 1995)	Cointegration

Table 8.7 – Cointegration between US house prices and economic fundamentals.

Authors	Sample data and variables	Method	Finding
Arestis and González (2014)	National: 1970-2011 Variables: real house price, real disposable income, real residential investment, volume of banking credit, mortgage rate, ratio of taxation to property/house price, the rate of unemployment and the evolution of population.	Johansen (1991)	Cointegration
Bahmani-Oskooee and Ghodsi (2016)	Regional: 1975Q1-2014Q3 for 52 states Variables: house prices, personal income, mortgage rates.	Pesaran et al. (2001), Shin et al. (2014)	Linear cointegration is established in 17 states and DC. Non-linear cointegration is established in 24 states.
Chen and Cheng (2017)	National: 1979Q1-2015Q3 Variables: house price-to-income ratio, interest rate, inflation rate, real income growth rate and housing demand.	Johansen (1991)	Cointegration

one of the top sectors as they are surrounded by world-class universities, research institutes and hospitals. Massachusetts is also famous for its high-tech industry. New York state is known as the Empire State. The largest city and the most populous city in New York state is New York City, which has more than 40% of the state's population. New York city is well-known as a global hub of business, commerce and high-tech industries. New York City's most important economic sector is its financial industry, where the New York Stock Exchange and the NASDAQ represent the world's largest and second largest stock exchanges. Silicon Alley in Manhattan is an area of New York City, which is associated with high-tech industries. Nevada is famous for its tourism and gaming industry, especially, Las Vegas, which is an internationally renowned major resort city, known primarily for its gambling, shopping, fine dining, entertainment, and nightlife.

The Freddie Mac House Price Index (FMHPI), January 1978 and December 2016, is used to measures home price inflation at the both national and state level in this study, where the house price index represents the value of single-family housing.[2] These aggregated national and disaggregated state indices are not seasonally adjusted. The FMHPI is calculated using a repeat-transactions methodology. Repeat transactions indices measure price appreciation while holding constant housing type and location, by comparing the price of the same property over two or more transactions. As we work at a quarterly frequency, we create a quarterly house price index from 1978Q1 to 2016Q4 by averaging.

Apart from house prices, the choice of housing fundamentals is also crucial. Following Zhou (2010), three critical economic fundamentals are also selected in this research: personal incomes, construction costs and mortgage rates. Following Mikhed and Zemčík (2009a), Zhou (2010) and Bahmani-Oskooee and Ghodsi (2016), total personal income from the Bureau of Economic Analysis (BEA) is used at the state level, where the total personal income is reported as the product

[2]The national index is defined as a weighted average of the 50 states and Washington, D.C. indices.

of per-capita and population.[3]

In order to take account of construction costs, we use the relevant indexes at the national- and regional-level from United States Census Bureau. For the national data, the monthly indexes for houses under construction are used. The under construction indexes are based on data for a sample of all single-family houses. However, land and other nonconstruction costs are not included. For individual states, the house sold index, which incorporates the value of the land, is used for measuring construction costs and is available annually by four regions (Northeast, Midwest, South and West). A state located within the four areas is used as an indicator. In particular, sold indexes are based on data for a sample of houses that are built for sale with the land upon which the house is situated. A quarterly house sold index is interpolated by region.

The weekly 30-year fixed mortgage rates are available from Freddie Mac since 1971. Freddie Mac surveys approximately 125 lenders with a mix of lender types thrift institutions, credit unions, commercial banks and mortgage lending companies for their most popular mortgage products. The survey is based on first-lien prime conventional conforming home purchase mortgages with a loan-to-value of 80 percent. The 30-year fixed mortgage rates on a quarterly basis between 1978Q1 and 2016Q4 are created by averaging.

The respective CPI less shelter in four regions (Northeast, Midwest, South and West) and the national are obtained from U.S. Bureau of Labor Statistics to deflate the house price indexes, personal incomes, and construction costs into real values. These variables are then normalized to 100 at 1978Q1. A real series of mortgage rates can be derived using the inflation rates calculated from the respective CPI series by four regions as well. All variables are then converted into natural logarithm. [4]

[3]Personal income are available at the state level for our sample chosen period. However, BEA does not allow us to separate per-capita and population variables during our selected sample period.

[4]For mortgage rate variable, it is common to take log of one plus the mortgage rates.

8.4 Method

This section reviews the conditional mean and quantile based cointegration models of Pesaran and Shin (1998) and Pesaran et al. (2001), Shin et al. (2014), Cho et al. (2015) and Greenwood-Nimmo et al. (2011).

8.4.1 ARDL Model

According to Pesaran and Shin (1998) and Pesaran et al. (2001), the error correction model of the ARDL specification can be defined as:

$$\Delta y_t = \rho y_{t-1} + \theta \mathbf{x}_{t-1} + \sum_{j=1}^{p-1} \phi_j \Delta y_{t-j} + \sum_{j=0}^{q-1} \psi_j \Delta \mathbf{x}_{t-j} + u_t. \tag{8.13}$$

The long-run cointegration relationship can be tested using: the t-test of Banerjee et al. (1998) on ECM coefficient for testing the restriction $\rho = 0$ and the F test for testing the restriction $\rho = \theta = 0$. Both the asymptotic distribution of the F-statistic and t-statistic are non-standard under the null hypothesis of no cointegration. The relevant critical values are obtained from Pesaran et al. (2001). The ARDL approach can be applied regardless of whether the underlying regressors are integrated of order one I(1), I(0) or mutually integrated.

8.4.2 QARDL Model

The autoregressive distributed lag process can be defined as:

$$y_t = \alpha_* + \sum_{j=1}^{p} \phi_{j*} y_{t-j} + \sum_{j=0}^{q} \boldsymbol{\theta}'_{j*} \boldsymbol{x}_{t-j} + U_t. \tag{8.14}$$

The quantile autoregressive distributed lag (QARDL) process is given by:

$$y_t = \alpha_*(\tau) + \sum_{j=1}^{p} \phi_{j*}(\tau) y_{t-j} + \sum_{j=0}^{q} \boldsymbol{\theta}'_{j*}(\tau) \boldsymbol{x}_{t-j} + U_t(\tau), \tag{8.15}$$

where $\tau \in (0,1)$ is a quantile index, p and q are lag orders, and $|\sum_{j=1}^{p} \phi_{j*}(\tau)| < 1$ for $\tau \in (0,1)$.

A long run cointegrating relationship between Y_t and $\boldsymbol{X}_t$ is captured by the long-run parameter $\boldsymbol{\beta}_*$,

$$\boldsymbol{\beta}_*(\tau) = \left\{1 - \sum_{i=1}^{p} \phi_{j*}(\tau)\right\}^{-1} \sum_{j=0}^{q} \boldsymbol{\theta}'_{j*}(\tau). \tag{8.16}$$

The quantile ARDL error correction model allows us to investigate the quantile-dependent long-run relationship and the associated dynamic adjustments, simultaneously.

$$\Delta y_t = \alpha_*(\tau) + \rho_*(\tau) y_{t-1} + \boldsymbol{\theta}_*(\tau)\boldsymbol{x}_{t-1} + \sum_{j=1}^{p-1} \tilde{\phi}_{j*}(\tau)\Delta y_{t-j} + \sum_{j=0}^{q-1} \tilde{\boldsymbol{\theta}}_{j*}(\tau)\Delta \boldsymbol{x}_{t-j} + U_t(\tau) \tag{8.17}$$

No long-run relationship testing procedure is discussed in Cho et al. (2015) as they assume the existence of a long-run equilibrium relationship. The null hypothesis of no (linear) quantile cointegration in Equation (8.15) can be tested in two ways: a Wald test for the joint restriction $\rho_*(\tau) = \boldsymbol{\theta}_*(\tau) = 0$ and the t-test of Banerjee et al. (1998) for the restriction $\rho_*(\tau) = 0$. It should be pointed out that Prof Yongcheol Shin suggests to use a Ward test for the joint restriction $\rho_*(\tau) = \boldsymbol{\theta}_*(\tau) = 0$ rather than the F test under the null of no quantile cointegration.

8.4.3 NARDL Model

Shin et al. (2014) advance a simple technique for modelling both short- and long-run asymmetries based on extension of the linear ARDL approach of Pesaran and Shin (1998) and Pesaran et al. (2001). Shin et al. (2014) introduces the following asymmetric cointegrating relationship:

$$y_t = \boldsymbol{\beta}^{+'}\boldsymbol{x}_t^+ + \boldsymbol{\beta}^{-'}\boldsymbol{x}_t^- + u_t, \tag{8.18}$$

where $\boldsymbol{x}_t$ is a $k \times 1$ vector of regressors decomposed as:

$$\boldsymbol{x}_t = \boldsymbol{x}_0 + \boldsymbol{x}_t^+ + \boldsymbol{x}_t^-, \tag{8.19}$$

where $\boldsymbol{x}_t^+$ and $\boldsymbol{x}_t^-$ are partial sum process if positive and negative values changes in $\boldsymbol{x}_t$ defined by

$$\boldsymbol{x}_t^+ = \sum_{j=1}^{t} \Delta\boldsymbol{x}_j^+ = \sum_{j=1}^{t} \max(\Delta\boldsymbol{x}_j, 0), \boldsymbol{x}_t^- = \sum_{j=1}^{t} \Delta\boldsymbol{x}_j^- = \sum_{j=1}^{t} \min(\Delta\boldsymbol{x}_j, 0), \quad (8.20)$$

and $\boldsymbol{\beta}^+ = -\boldsymbol{\theta}^+/\rho$ and $\boldsymbol{\beta}^- = -\boldsymbol{\theta}^-/\rho$ are the associated asymmetric long-run parameters. Using a similar idea, the corresponding asymmetric ARDL error correction model may be defined as:

$$\Delta y_t = \rho y_{t-1} + \boldsymbol{\theta}^+\boldsymbol{x}_{t-1}^+ + \boldsymbol{\theta}^-\boldsymbol{x}_{t-1}^- + \sum_{j=1}^{p-1} \varphi_j \Delta y_{t-j} + \sum_{j=0}^{q} \left(\boldsymbol{\pi}_j^+ \Delta\boldsymbol{x}_{t-j}^+ + \boldsymbol{\pi}_j^- \Delta\boldsymbol{x}_{t-j}^-\right) + \epsilon_t. \quad (8.21)$$

Shin et al. (2014) refer to Equation (8.21) as the "asymmetric or non-linear ARDL (NARDL) model". They also demonstrate that the NARDL approach has overwhelming advantages over regime-switch models. First, Equation (8.21) can be estimated by OLS. Second, the null hypothesis of no long-run relationship between y_t, $\boldsymbol{x}_t^+$ and $\boldsymbol{x}_t^-$ can be tested using the F bounds-testing procedure ($\rho = \boldsymbol{\theta}^+ = \boldsymbol{\theta}^- = 0$) and the procedure remains valid irrespective of whether the regressors are I(0), I(1) or mutually cointegrated. The long-run relationship can be also tested using the t-statistic of Banerjee et al. (1998) for the restriction $\rho = 0$. If $\rho = 0$, Equation (8.21) reduces to regression only involving first difference terms, indicating that there is no long-run relationship between y_t, $\boldsymbol{x}_t^+$ and $\boldsymbol{x}_t^-$. Third, both the long-run reaction symmetry and short-run adjustment symmetry can be tested using standard Wald tests.

8.4.4 NARDL-Q Model

By allowing non-linearity of the form modelled by Equation (8.21) into the conditional quantile function, the NARDL-Q model is obtained as follows:

$$\Delta y_t = \rho(\kappa) y_{t-1} + \theta_{(\kappa)}^+ \mathbf{x}_{t-1}^+ + \theta_{(\kappa)}^- \mathbf{x}_{t-1}^- + \sum_{j=1}^{p-1} \varphi_{(\kappa)j} \Delta y_{t-j} + \sum_{j=0}^{q} \left(\pi_{(\kappa)j}^+ \Delta\mathbf{x}_{t-j}^+ + \pi_{(\kappa)j}^- \Delta\mathbf{x}_{t-j}^-\right) + \epsilon_{(\kappa)t}, \quad (8.22)$$

where κ is a given quantile index in (0,1). The details of the NARDL-Q model can be found in Greenwood-Nimmo et al. (2011). Similar to the NARDL and QARDL model, the null of no cointegration can be tested using the t-test for the restriction $\rho_*(\tau) = 0$ and the Wald-test for the restriction $(\rho_*(\tau) = \boldsymbol{\theta}_\kappa^+ = \boldsymbol{\theta}_\kappa^- = 0)$.

8.4.5 Interpreting Results

For the ARDL and NARDL models, the t-statistic and F-statistic testing procedures can be used for the existence of a long-run relationship. For the QARDL and NARDL-Q models, the t-statistic and Wald-statistic testing procedures are used for testing the existence of a long-run relationship. A 'spurious' cointegrating relationship may arise if the relevant testing procedures yield conflicting results. It should be pointed out that a finding of cointegrating relationship is valid only if the relevant statistics are simultaneously significant.

8.5 Results

This section presents the results of various cointegration models at the both national and State levels. Two conditional mean-based ARDL models and two quantile versions of ARDL models are applied to explore a long-run cointegrating relationship between house prices and fundamentals in the US at the (aggregate) national and (disaggregate) State levels. It is well-known that the F and t testing procedures have a non-standard distribution which depends on the number of regressors, inclusion/exclusion of an intercept or a trend and the sample size. Following the standard practices in the literature, the relevant critical values are obtained for t_{BDM} and F_{PSS} statistic from Pesaran et al. (2001). As the economic fundamentals include three variables, it is reasonable to set $k = 3$ and $k = 6$ in obtaining critical values for linear and non-linear models.[5] Pesaran et al. (2001) tabulate the $k = 3$ critical values of t_{BDM} at the 10%, 5% and 1% as -3.46, -3.78

[5]Due to the decomposition for the non-linear models of Shin et al. (2014) and Greenwood-Nimmo et al. (2011), we have 6 exogenous variables. The critical values are selected for $k = 6$.

and -4.37, respectively. They also tabulate the relevant critical values of F_{PSS} as 3.77, 4.35 and 5.61 at the 10%, 5% and 1% level for $k = 3$. For the non-linear models of NARDL-M and NARDL-Q, we select the $k = 6$ critical values due to the partial sum decompositions from Pesaran et al. (2001). The $k = 6$ critical values of t_{BDM} are -4.04, -4.38, -4.99 at the 10%, 5%, and 1% significance level and the equivalent values of F_{PSS} are 3.23, 3.61 and 4.43, respectively. All the analysis is implemented by Gretl. The general-to-specific lag selection is performed with a maximum lag length of 4 using a sequential 10% rule. As we work with the quarterly data, it is reasonable to set the maximum lag order to be 4.

8.5.1 Unit Root Tests

We first apply unit root tests of the augmented Dickey-Fuller (ADF), Phillips-Perron (PP) and KPSS to check the stationarity of time series. Results are based on the model specification of a constant, and a constant with a trend and are presented in Table 8.14 through Table 8.23. In addition, the quantile Kolmogorov-Smirnov (QKS) test from Koenker and Xiao (2004) is also utilised to assess the unit root behavior across quantiles. The conventional unit root tests concentrate on the conditional central tendency of the time series while the QKS test provides a general perspective of the unit root behavior over all the quantiles. Among these above mentioned tables, the house price, income, construction costs and mortgage rates are abbreviated as HP, IC, CC and MR, respectively. The relevant partial sum process of positive and negative changes in the regressors are denoted by adding a P or a N after its variable name. For example, the partial sum process of positive and negative changes in income are denoted as 'ICP' and 'ICN'.

Unit root results at the national level are shown in Table 8.14. The house price (HP) seems to be I(1) under most unit root tests. However, there is evidence to suggest that this variable is I(0) under the model specification with a constant and a trend for ADF and KPSS tests. The income variable (IC) is I(1) as suggested by all tests. The construction cost (CC) is I(1) except the KPSS test under the assumption with a constant and a trend. The mortgage rate (MR) is I(0) under the assumption with a constant and trend for ADF, PP and KPSS tests. The partial sum process of positive and negative changes in income are denoted in ICP and ICN, respectively. ICP and ICN are I(1) in most cases. A similar conclusion may be drawn for the partial sum process of positive and negative changes in construction costs (CCP and CCN) and in mortgage rates (MRP and MRN). Overall, there is no overwhelming evidence to suggest that all variables are I(1) at the national level except the income variable.

Table 8.15 shows unit root results for Arizona. HP is I(1) except the KPSS test under the assumption with a constant and trend, and the QKS test. IC, ICP and CCP variables are I(1) as suggested by all tests. CC and ICP are likely to be I(1) except the quantile-based QKS test. It is not clear that the exact order of integration for MR as MR is hardly likely to be I(1). CCN, MRP and MRN are likely to be I(1) except based upon the KPSS test. Overall, there is strong evidence that all variables are I(1) based on ADF and PP tests only except the MR variable. When additional KPSS and QKS tests are considered, there is no firm evidence to conclude the presence of I(1) variable except IC, ICP and CCP.

Table 8.16 presents unit root results for California. HP is I(0) as the ADF and KPSS tests under the assumption with a constant and trend. QKS also indicates that HP is I(0). IC and CCN are strongly supported to be I(1) variable by all tests. There is no clear evidence to conclude that MR and ICP are I(1) variables based on three conventional unit root tests. The ADF and PP unit root tests suggest that CC, ICN, CCP, MRP and MRN are I(1) under different cases. However, the KPSS test does not provide the same conclusion regarding the order of these variables. Overall, there is no firm evidence to suggest that all variables are I(1) except IC and CCN.

Table 8.17 also shows mixed order of integration results among variables for the District of Columbia. There is evidence that HP, IC, CC and CCN are I(1) based on most tests. ICP is I(1) except for the QKS test. All tests suggest that MRP and MRN are I(1). It seems that MR and ICN are I(0) for conventional unit root tests especially with an estimated equation with a constant and trend.

According to Table 8.18, there is strong evidence that ICP, CCN and MRP for Florida are I(1) variables based on three conventional unit root tests but not the QKS test. HP is I(1) under most cases except ADF and KPSS tests with a a constant and a trend. The QKS test also supports an I(0) variable for HP. Thus the exact order of HP is not clear. Similar conclusion may be drawn for MR as well, as it seems to be an I(1) variable under a constant only in the equation while it is an I(0) variable under a constant and a trend in the equation. There are ambiguous results for IC as it is an I(0) variable for the ADF test and an I(2) variable for the KPSS test under the assumption of a constant only. CC, ICN, CCP and MRN are all identified as I(1) variables under the ADF and PP tests only. When the KPSS is applied, the exact order of these variables is not sure.

As presented in Table 8.19, there is strong evidence that IC, CC and CCN for Hawaii are I(1) by all three conventional unit root and QKS tests. There is evidence that HP, ICP, ICN, MRP are I(1) as suggested by three conventional unit root tests except the KPSS test with a constant and a trend in the equation. On the other hand, CCP and MRN are I(1) except based on the KPSS test with a constant only in the equation.

As shown in Table 8.20, several variables for Massachusetts are I(1) based on all three conventional unit root and QKS tests including IC, ICP, CCP, MRP and MRN. ICN also shows strong signs of I(1) except the QKS test. CC and CCN are I(1) variables except the KPSS test with a constant and a trend. HP and MR are I(0) only for the ADF and KPSS tests with a constant and a trend. Hence the exact order of these two variables is not certain.

As indicated in Table 8.21, ICP and CCN for Nevada are clearly I(1) variables. CC can be regarded as an I(1) variable by three conventional unit root tests. It is reasonable to believe that HP is an I(0) variable as the ADF, KPSS and QKS tests suggest. IC, ICN, CCP, MRP and MRN are I(1) based on the ADF and PP tests only. All three conventional unit root tests indicate that MR is I(0) under the assumption of a constant and a trend in the equation together with the QKS test.

As illustrated by Table 8.22, there is strong evidence that all variables for New York are I(1) in most cases except the HP and MR. Both HP and MR exhibit evidence of I(0) variables based upon conventional unit root tests.

Table 8.23 provides a clear order of integration for most variables in Pennsylvania. There is strong evidence that IC, ICP, CCP, MRP and MRN are I(1). CC, MR and CCN are also believed to be I(1) variables for most cases. The only exception is the HP variable. It is identified as an I(0) in certain cases. In short, there is mixed evidence to suggest that HP is an I(1) variable.

8.5.2 Johansen Cointegration Tests

If all variables are assumed to be I(1), we can then apply Johansen cointegration tests to explore a long-run relationship between house prices, and economic fundamentals at the conditional mean. There are five different assumptions regarding trend specification in Johansen (1995). In this paper, we choose trend assumption 3 and 4 as discussed in Johansen (1995), which are the most widely chosen trend specification in applying the test. The trend case 3 assumes the level data assumes to have linear trend but the cointegrating equations have only intercept while the case 4 assumes the level data and the cointegrating equations have linear trends. The corresponding cointegration results are presented in Table 8.8 based on both trace and maximum eigenvalue tests.

As illustrated in Table 8.8, under the trend assumption of case 3, the Johansen test identifies at least one cointegrating relationship at the both national and regional levels except for the District of Columbia (DC), indicating a long-run equilibrium relationship among house prices and fundamentals for these States.

Under the assumption of trend specification 4, the Johansen test identifies at least one cointegrating relationship based on the trace test at the both national and regional levels except the District of Columbia (DC), supporting the cointegration relationship between house price and fundamentals. However, the maximum eigenvalue statistics suggest that the cointegration relationship holds only in Florida (FL), Hawaii (HI), Massachusetts (MA), New York (NY) and Pennsylvania (PA).

8.5.2.1 Other Relevant US Studies using Johansen Cointegration Tests

Several previous studies attempted to examine a long run relationship between the US house prices and fundamentals using Johansen cointegration tests. For example, Duca et al. (2011) conclude that the US price-to-rent ratio, user costs and loan-to-value ratio (LTV) are cointegrated at the national level based on the trace and maximal eigenvalue statistics between 1981Q1 and 2007Q2. However, Zhou (2010) shows that the house price is not cointegrated with fundamentals (income, construction costs, mortgage rates) at the national and city levels except Cleveland covering the period 1978Q1-2007Q4. Arestis and González (2014) present evidence of a long-run relationship at the national level between the real house price and a set of fundamentals (real disposable income, real residential investment, volume of banking credit, mortgage rate, ratio of taxation to property/house price, the rate of unemployment and the evolution of population) for the period 1970-2011. Chen and Cheng (2017) argue that the US house price-to-income ratio, interest rate, inflation rate, real income growth rate and housing demand are cointegrated from 1979Q1 to 2015Q3 at the national level.

8.5.3 ARDL Model Results

8.5.3.1 National-level

Results for the US national data are presented as Table 8.9. The t_{BDM} statistic testing $\rho = 0$ based on Equation (8.13) and Equation (8.21) for the ARDL-M

Table 8.8 – Number of cointegrating relationships based on Johansen cointegrating test based on different trend assumptions for the US and a selection of states.

Trend assumption	US	AZ	CA	DC	FL	HI	MA	NV	NY	PA
Case 3: the level data have linear trends										
but the cointegrating equations have only intercepts										
No. of cointegrating relationships										
Trace	1	2	1	0	3	1	1	4	1	1
Maximum-Eigenvalue	1	2	1	0	1	1	1	1	2	1
Case 4: the level data and the cointegrating										
equations have linear trends										
No. of cointegrating relationships										
Trace	1	1	1	0	3	1	1	1	1	1
Maximum-Eigenvalue	0	0	0	0	1	1	1	0	1	1

and NARDL-M models is -1.746 and -1.664, where both test statistics are not significantly different from zero at the 10% level, meaning that no cointegration. Although the F_{PSS} statistic is significant at the 5% level for both models, it seems to provide evidence of cointegration. However, the two testing procedures yield conflicting results, indicating 'spurious' cointegration relationship. A key assumption regarding the ARDL model is that the model is dynamically stable, suggesting ρ needs to be non-zero. Apart from the well-known F_{PSS} statistic, the t-test on $\rho = 0$ can also be regarded as one way of testing cointegration. Due to the inconsistent results from t_{BDM} and F_{PSS} tests, we conclude that no cointegration exist between house prices and fundamentals. Despite using the powerful NARDL-M model to take into account non-linearity, the above results still suggest no overwhelming evidence of cointegration between house prices and fundamentals at the national level.

We then investigate whether the non-cointegration relationship holds at the conditional quantile using two quantile models. From the upper panel of Table 8.9, the t_{BDM} statistic is not significantly different from zero at all quantiles across the whole conditional distribution of the house price for both the QARDL and NARDL-Q models, where the critical values are -3.46 and -4.04, respectively. The non-significant t_{BDM} statistic implies that house prices are neither symmetrically nor asymmetrically cointegrated with economic fundamentals at the national level. Results from two quantile models suggest that, for either low house prices (when house prices are at low quantile level), average house prices (when house price at the middle quantile level) or higher house prices (when house prices are at high quantile level), they are not cointegrated with economic fundamentals, indicating no long-run equilibrium relationship at the conditional quantile.

Meen (2002) argues that house prices and fundamentals are cointegrated in the US using the error correction model. Zhou (2010) and Clark and Coggin (2011) also draw the same conclusion at the national level using cointegration tests of Engle and Granger (1987), Johansen (1991), Gregory and Hansen (1996) and Granger and Hallman (1991). The finding of no stable long-run relationship at the aggregate national-level is consistent with the conclusion of Gallin (2006), Zhou (2010) and Clark and Coggin (2011). We then concentrate on empirical results from ARDL models. More importantly, the non-cointegrating relationship not only holds for the two conditional mean based linear and non-linear ARDL models but also establishes at two recently proposed quantile versions of ARDL models. Our results from Table 8.9 show neither lower, average nor higher house prices cointegrate with economic fundamentals. For a bear, normal, or bull market, house prices do not move towards to economic fundamentals. A key contribution from this paper is that we provide additional evidence to support further the hypothesis of no cointegration between house prices and various fundamentals at the national level using two quantile-based ARDL models. Given the fact that no stable long-run relationship can be established for both symmetric and non-symmetric quantile ARDL models, as argued by Gallin (2006), it would be inappropriate to model house price dynamics using an error-correction specification. Hence these studies on house price dynamics of Abraham and Hendershott (1996), Capozza et al. (2002) and Malpezzi (1999) based on error correction specifications are fundamentally misleading.

8.5.3.2 Arizona (AZ)

Results for Arizona are presented as Table 8.9. Both the ARDL-M and NARDL-M models indicate the presence of either a symmetric or an asymmetric long-run cointegrating relationhsip between house prices and their economic fundamentals based on the t-statistic testing procedure of Banerjee et al. (1998) and F-test of Pesaran et al. (2001). For the ARDL-M model, both the t_{BDM} and F_{PSS} statistic are significant at the 1% level (e.g., t_{BDM} statistic:-5.933; F_{PSS} statistic: 9.051). Similarly, an asymmetric long-run relationship can also be observed between house prices and their economic fundamentals in Arizona as the t_{BDM} and F_{PSS} tests are significant at the 1% level. Overall, two conditional mean-based ARDL cointegration models present strong evidence of a long-run cointegrating relationship for Arizona.

Two quantile ARDL models also exhibit strong evidence of a cointegration relationship based on the t-statistic of Banerjee et al. (1998) and a Wald test for the joint null at the conditional quantile. House prices in Arizona and their economic fundamentals are cointegrated at a conditional quantile if the statistics from the two testing procedures are simultaneously significant. In particular, Table 8.10 indicates a long-run relationship between house prices and economic fundamentals for the whole conditional distribution except the 10% quantile based on the ARDL-Q and NARDL-Q models. At the 10% quantile, there is no long-run relationship between house prices and fundamentals. This means that, when house prices in Arizona are at the 10% lower level (lower house price), it is neither symmetrically nor asymmetrically cointegrated with economic fundamentals.

This probably means that, a rise or fall in lower house prices is not driven by economic fundamentals as there is no long-run equilibrium relationship between the bear housing market and fundamentals. Hence, market forces will probably not significantly affect lower house prices. Changes in fundamentals are unlikely to lead to a stable long-run relationship with the bear housing market (lower house prices). In other words, if house prices are far too low relative to fundamentals, they lose their cointegration relationship.

8.5.3.3 District of Columbia (DC)

The ARDL-M model finds no long-run equilibrium relationship as the t-statistic with a value of -3.122 is not significant at the 10% significance level. On the other hand, the NARDL-M model shows some interesting results as both the t_{BDM} and F_{PSS} statistic are significant at the 10% level, indicating an asymmetric long-run cointegration relationship between house prices and fundamentals.

Interestingly, when we consider the quantile version model of QARDL, we identify a long-run relationship at the lower quantile (e.g., from 10% to 30%) as both the t_{BDM} and F_{PSS} statistic are significant. At the 20% quantile, the t_{BDM} statistic is -3.454, which is very close to the critical value of -3.46. It is perhaps acceptable to conclude that house prices and fundamentals are symmetrically cointegrated in the lower quantile between 10% and 30%. Results from the QARDL model suggest that there is evidence of symmetric cointegration for a bear housing market as the house price is lower. However, as stated earlier, the conditional mean based ARDL-M model shows no evidence of cointegration, which assumes that the non-cointegrating relationship holds for the whole conditional distribution. A valuable lesson from this example is that the cointegration relationship may be absent at the conditional mean-based cointegration model (e.g., ARDL-M model) while it may be established at the conditional quantile-based cointegration model (e.g., QARDL model).

The asymmetric version of the NARDL-Q model presents some interesting results. It seems that house prices are asymmetrically cointegrated with fundamentals at the 10% quantile only based on the NARDL-Q model while asymmetric cointegration is established based on two testing procedures for the NARDL-M model. Interestingly, the NARDL-Q model shows that house prices reflect fundamentals in the lower quantile, indicating that a long-run stable relationship seems to hold for a bear market. Although an asymmetric cointegrating relationship is identified based on the conditional mean model of NARDL, no overwhelming evidence supports asymmetric cointegration at the conditional quantile. This is an important example as it shows that the cointegration may be established based on the conditional mean model (e.g., NARDL-M model) but there could be no overwhelming evidence of cointegration based on its relevant quantile model (e.g., NARDL-Q model). This is to say that the assumption of relationship does not hold at the entire conditional distribution.

The lower house prices for a bear market in the District of Columbia are cointegrated with economic fundamentals between the 10% and 30% quantiles based on the QARDL model while house prices are asymmetrically cointegrated with economic fundamentals at the 10% quantile only based on the NARDL-Q model. These results indicate that lower house prices are driven by economic fundamentals. A rise or drop in low house prices is likely to associated with changes in fundamentals to catch up as there is long-run equilibrium relationship. Therefore, changes in fundamentals are likely to be effective in maintaining lower house price inflation.

8.5.3.4 Florida (FL)

Both conditional mean models indicate no long-run equilibrium relationship in Table 8.11, where the t-statistics of Banerjee et al. (1998) for the two models are not significant from zero (t_{BDM} statistic:-2.878 and -3.185 for ARDL-M and NARDL-M, respectively). Furthermore, the F_{PSS} statistic for the two models is not significant. As the ARDL-M and NARDL-M models are not dynamically stable with non-significant F_{PSS} statistics, we may conclude that house prices in Florida and the relevant economic fundamentals are neither symmetrically nor asymmetrically cointegrated.

The QARDL model reveals no evidence of a long-run equilibrium relationship as house prices and fundamentals are not cointegrated across the whole conditional distribution. For either low, average or high house prices, there is no long-run cointegrating relationship. This is an additional insight offered from the QARDL model.

The NARDL-Q model shows some interesting findings, which indicates that house prices and economic fundamentals in Florida are only asymmetrically cointegrated at the lower tail (e.g., 10% and 20% quantile) in the long-run due to the significant t- and Ward statistic at the 1% level. House prices in Florida is asymmetrically cointegrated with personal income, constriction costs and mortgage rates only at the low quantile. This example also demonstrates that the conditional mean based cointegration test may be unable to identify a long-run relationship while the quantile-based cointegration may uncover some evidence at specific quantiles.

The above results likely mean that a low house price in Florida is asymmetrically cointegrated with those economic fundamentals for low-income earners in the long-run as these people with low income are more likely to purchase low priced houses. For those people who earn a reasonable salary or can afford to retire in Florida with a certain amount of wealth, they are more likely to purchase an average house or a more expensive house than those low incomes. These high-income earners' or wealthy retirees' buy relatively more expensive houses which are not simply cointegrated with the relevant economic fundamentals in the long-run as t- and Ward statistic are not simultaneously significant from the 30% to 90% quantile. An average and high house price in Florida does not seem to reflect fundamentals including income, population growth, construction costs and mortgage rates. Those houses owned by high-income earners or wealthy retirees are unlikely to be significantly affected by changes in the economic fundamentals, or at least, fundamentals will not catch up with the house price.

These results have strong policy implications to policy-makers. As our key selection of economic fundamentals are not cointegrated with higher house prices, changes in these economic fundamentals are less likely to be effective in controlling higher house prices in Florida. However, changes in fundamentals will be more likely to be effective in controlling those lower price houses as cointegration is established in low price houses only.

8.5.3.5 California (CA)

Results for California are presented as Table 8.10. Based on two testing procedures for the existence of a symmetric or an asymmetric long-run relationship at the conditional mean, both the ARDL-M and NARDL-M models suggest no cointegrating relationship as the t-statistic of Banerjee et al. (1998) is not significant for the two models. The t_{BDM} statistic is -2.823 and -1.787 for ARDL-M and NARDL-M models while the 10% critical value is -3.46 and -4.04, respectively. Neither the ARDL-M model nor NARDL-M model is dynamically stable as the relevant t_{BDM} statistic is not significant at the 10% level, indicating no long-run relationship. Thus house prices in California are not cointegrated with the relevant economic fundamentals (e.g., personal income (a combination of income and population), construction costs and mortgage rates) at the conditional mean based on linear and non-linear ARDL models.

The same conclusion can be drawn for the two quantile ARDL models. The t-statistic test fails to identify the existence of a symmetric or an asymmetric long-run relationship for the whole conditional distribution for the QARDL and NARDL-Q models, indicating house prices are not aligned with economic fundamentals in California. For lower, average or higher house prices, these house prices do not cointegrate with economic fundamentals. Based on the above, the results show substantial evidence of no long-run cointegrating relationship between house prices and relevant economic fundamentals in California based on both conditional mean and quantile models.

Housing in California has long been more expensive than most of the rest of the country. The critical conclusion perhaps suggests that Californian housing market is not driven by economic fundamentals. For either, a bear, normal or boom market, house price in California does not reflect fundamentals. Hence, the Californian housing market will probably not respond much in the changes of economic fundamentals. In other words, policy regarding these fundamentals is unlikely to be effective in fighting high house prices in California.

8.5.3.6 Hawaii (HI)

The two conditional mean models of ARDL-M and NARDL-M clearly indicate no long-run relationship between house prices and their economic fundamentals as the F_{PSS} statistic and t_{BDM} statistic are not significant under both models. The t_{BDM} statistic is -0.854 and -1.688 for the ARDL-M and NARDL-M models, which are not significant at the 10% critical value of -3.46 and -4.04. Due to consistent results of two testing procedures, it is reasonable to conclude no long-run relationship between house prices and the relevant economic fundamentals in Hawaii. Both conditional mean models seem to suggest that, for those beachgoers in Hawaii, house prices are neither symmetrically nor asymmetrically cointegrated income, population, construction costs and mortgage rates.

Neither the QARDL model nor NARDL-Q model offers overwhelming evidence of cointegration across the whole conditional distribution as suggested by the two testing procedures. We, therefore, have firm evidence to conclude that there is no stable long-run relationship between house prices and fundamentals in Hawaii based on both conditional mean- and quantile-based models.

8.5.3.7 Massachusetts (MA)

It is clear from Table 8.12 that house prices in Massachusetts are not cointegrated with economic fundamentals based on the two conditional mean models. The t_{BDM} statistic for ARDL-M and NARDL-M models is -0.115 and -1.313, which are not significantly different from zero, suggesting no long-run relationship. Despite a significant F-statistic of the joint null, we conclude no long-run equilibrium relationship between house prices and fundamentals for the ARDL-M and NARDL-M models.

The same conclusions can be drawn from the two quantile models. The t_{BDM} statistic is not significant across the whole distribution for the two quantile-based models. By applying the QARDL and NARDL-Q models, our results highlight that for a low-, an average or high-value house in Massachusetts, house prices are not cointegrated with the relevant economic fundamentals. This is the key finding from the two quantile models. Based on the above results, there is firm evidence to conclude no stable long-run relationship between house prices and these fundamentals in Massachusetts based on several ARDL cointegration models.

8.5.3.8 New York (NY)

Based on the two conditional mean models, the t_{BDM} statistic is -0.173 and -0.861 for the ARDL-M and NARDL-M models and is not significantly different from zero, indicating no long-run equilibrium relationship. At the conditional mean level, house prices in New York are neither symmetrically nor asymmetrically cointegrated with economic fundamentals. Whatever the house price level is, there is no stable long-run relationship.

Both the QARDL and NARDL-Q models confirm no long-run relationship between house prices and economic fundamentals across the whole conditional distribution as neither t_{BDM} statistic nor F_{PSS} statistic are simultaneously significant. Results from New York are similar with those of Hawaii and Massachusetts, where no cointegration is established for both the symmetric and non-symmetric conditional mean- and quantile-based ARDL cointegration models.

Whatever the house price level is, house prices in New York state are neither symmetrically nor asymmetrically cointegrated at the conditional quantile. This indicates that, for a bear, normal or boom market, house prices in New York state do not reflect economic fundamentals. Changes in the fundamentals are unlikely to be effective in influencing house price inflation. Policy relating to economic fundamentals will not impact much on house prices. A drift between house prices and fundamentals in New York is unlikely to come back to an equilibrium level. For example, economic fundamentals are unlikely to catch-up the rise in house prices.

8.5.3.9 Nevada (NV)

As shown as Table 8.12, the ARDL-M model identifies a symmetric long-run relationship between house prices in Nevada and economic fundamentals as the t_{BDM} and F_{PSS} statistic with values of -3.967 and 5.466 are significant at the 10% and 1% level, respectively. The NARDL-M model also supports an asymmetric long-run relationship due to the significant t_{BDM} and F_{PSS} statistic at the 5% level. At the conditional mean level, the above results based on the two ARDL models clearly show that house prices are cointegrated with fundamentals.

The QARDL model offers some interesting results. It can be seen from Table 8.13 that the t_{BDM} and F_{PSS} statistic are both significant for the 60% to 90% quantiles. A higher house price or above-average price are cointegrated with economic fundamentals as the cointegration is established between the 60% and 90% quantile. At the conditional mean level, house prices are symmetrically cointegrated with economic fundamentals as suggested by the ARDL-M model. Bear in mind that the assumption of the ARDL-M and NARDL-M models is that the relationship not only holds at the mean but in other parts of the conditional distribution of the dependent variable (e.g., lower and upper tail quantiles). The cointegration established by the ARDL-M or NARDL-M model assumes that the long-run relationship holds not only at the lower tail but also the upper tail. However, it is clear that the cointegration does not hold at all quantiles as a symmetric long-run cointegration relationship is absent between the 10% and 50% quantile based on QARDL model. This example demonstrates the importance of considering quantile cointegration models as the estimated cointegration relationship from the conditional mean-based cointegration models may not hold across the entire conditional distribution.

The NARDL-Q model confirms the existence of an asymmetric long-run relationship at the 50%, 60% and 90% quantile due to the significant t_{BDM} and F_{PSS} statistic. It should be pointed out that the t_{BDM} statistic is -3.571 and -3.646 at the 70% and 80% quantile. Although they are not significant at the 10% level if $k = 6$ critical value of -4.04 is selected, they are significant at the 10% level if $k = 3$ critical value of -3.46 is selected. Therefore, we have some weak evidence of cointegration at the 70% and 80% quantile.

Based on the two quantile models, it is reasonable to conclude that house prices in Nevada do not align with the relevant economic fundamentals in the lower quantiles (e.g., 10%-40% quantiles). This is to say, there is no long-run relationship between lower-value house prices and economic fundamentals. These results are interesting. Nevada is famous for its tourism and gaming industry. Las Vegas is an internationally renowned major resort city, known primarily for its gambling, shopping, fine dining, entertainment, and nightlife. A large number of people pursue careers in hospitality and tourism there. They are more likely to own lower or average price houses than those high-income earners who probably own more luxury houses. Results from two quantile models suggest that, for those people who pursue careers in hospitality and tourism industry, their relatively lower price houses are not cointegrated with economic fundamentals. Those people who work in tourism and hospitality industries earn a relatively low salary, and their income can be more sensitive to house prices. Their job is not as secure as those highly trained/educated professionals as they could lose jobs for some unforeseen circumstances. The two quantile ARDL models indicate that these house prices in the lower price level are not driven by economic fundamentals and are less likely to be affected by changes in the fundamentals.

8.5.3.10 Pennsylvania (PA)

Cointegration results for Pennsylvania are presented in Table 8.13. Based on the t_{BDM} and F_{PSS} statistic with values of -2.476 and 2.348, the ARDL-M model identifies no cointegration relationship at the conditional mean of the house prices in

Pennsylvania. On the other hand, the NARDL-M model provides some interesting results as it finds evidence of an asymmetric long-run equilibrium relationship based on two testing procedures based on the t_{BDM} and F_{PSS} statistic at the 1% level.

Cointegration seems to be established in the lower tail as suggested by the QARDL model. As can be seen from Table 8.13, house prices are symmetrically cointegrated with economic fundamentals at the 10% and 30% quantiles only. At the 20% quantile, the t_{BDM} statistic for the ECM parameter is -3.414 and is not significantly different from zero. However, the t_{BDM} statistic of -3.414 is very close to the 10% critical value of -3.46. Therefore, it is probably acceptable to conclude that cointegration is present in the lower quantiles.

Results from the NARDL-Q model indicates that house prices in Pennsylvania are asymmetrically cointegrated with economic fundamentals at all quantiles across the entire conditional distribution, which is a key feature. The NARDL-M model finds asymmetric long-run cointegration, which is assumed to hold in the whole conditional distribution. Our results from the NARDL-Q model further confirm that the asymmetric long-run relationship holds in the entire conditional distribution. For either lower or higher house prices, these prices are asymmetrically cointegrated with fundamentals.

As suggested above, the linear models of ARDL-M and QARDL show no overwhelming evidence of a symmetric cointegration relationship. Overall, as suggested by the non-linear models of NARDL-M and NARDL-Q, house prices in Pennsylvania are asymmetrically cointegrated with economic fundamentals based on two testing procedures. This may indicate that, either house prices and fundamentals may drift apart temporarily from each other, but they will return to the equilibrium level in the long run based on two non-linear models.

Table 8.9 – Cointegration between US house prices and economic fundamentals.

y-x		$\tau = 0.1$	$\tau = 0.2$	$\tau = 0.3$	$\tau = 0.4$	$\tau = 0.5$	$\tau = 0.6$	$\tau = 0.7$	$\tau = 0.8$	$\tau = 0.9$
National level										
QARDL	T-stat$_{\rho_*(\tau)=0}$	-1.071	-1.662	-2.852	-2.758	-1.144	-1.069	-2.568	-1.880	-1.448
model	W-stat$_{\theta_*(\tau)=\rho_*(\tau)=0}$	1.985	22.114***	23.887***	29.434***	64.072***	21.178***	32.929***	14.996***	19.953***
	Cointegration									
NARDL-Q	T-stat$_{\rho_*(\tau)=0}$	-2.114	-1.942	-1.286	0.470	-0.625	0.453	-0.500	-3.035	-2.268
model	W-stat$_{\rho_*(\tau)=\theta_*^+(\tau)=\theta_*^-(\tau)=0}$	253.580***	68.187***	86.816***	31.926***	40.904***	28.205***	25.965***	27.419***	36.644***
	Cointegration									
ARDL-M	T-stat$_{\rho=0}$	-1.746								
model	F-stat$_{\theta=\rho=0}$	5.012**								
	Cointegration									
NARDL-M	T-stat$_{\rho=0}$	-1.664								
model	F-stat$_{\rho=\theta^+=\theta^-=0}$	4.121**								
	Cointegration									
Arizona(AZ)										
QARDL	T-stat$_{\rho_*(\tau)=0}$	-2.004	-4.095***	-8.306***	-10.710***	-8.163***	-8.756***	-5.929***	-4.658***	-4.014***
model	W-stat$_{\theta_*(\tau)=\rho_*(\tau)=0}$	18.646***	19.886***	76.595***	146.349***	83.818***	113.417***	38.562***	25.281***	18.570***
	Cointegration		**Y**	**Y**	**Y**	**Y**	**Y**	**Y**	**Y**	**Y**
NARDL-Q	T-stat$_{\rho_*(\tau)=0}$	-3.740	-5.143***	-9.091***	-7.851***	-6.684***	-4.445**	-2.585	-4.697**	-5.985***
model	W-stat$_{\rho_*(\tau)=\theta_*^+(\tau)=\theta_*^-(\tau)=0}$	114.035***	42.472***	136.622***	79.286***	59.447***	38.554***	31.604***	40.401***	1011.750***
	Cointegration		**Y**	**Y**	**Y**	**Y**	**Y**		**Y**	**Y**
ARDL-M	T-stat$_{\rho=0}$	-5.933***								
model	F-stat$_{\theta=\rho=0}$	9.051***								
	Cointegration	**Y**								
NARDL-M	T-stat$_{\rho=0}$	-5.821***								
model	F-stat$_{\rho=\theta^+=\theta^-=0}$	5.211***								
	Cointegration	**Y**								

Table 8.10 – Cointegration between US house prices and economic fundamentals.

y-x		$\tau = 0.1$	$\tau = 0.2$	$\tau = 0.3$	$\tau = 0.4$	$\tau = 0.5$	$\tau = 0.6$	$\tau = 0.7$	$\tau = 0.8$	$\tau = 0.9$
California(CA)										
QARDL	T-stat$_{\rho_*(\tau)=0}$	-1.069	-3.274	-2.545	-3.766*	-2.389	-3.090	-2.461	-2.618	-2.742
model	W-stat$_{\theta_*(\tau)=\rho_*(\tau)=0}$	5.190	30.609***	22.338***	28.282***	13.443*	21.056***	9.408*	7.760	13.400***
	Cointegration									
NARDL-Q	T-stat$_{\rho_*(\tau)=0}$	-1.247	-4.724	-2.963	-1.267	-2.827	-1.494	-1.764	-1.230	-1.176
model	W-stat$_{\rho_*(\tau)=\theta_*^+(\tau)=\theta_*^-(\tau)=0}$	63.950***	287.030***	42.325**	20.011**	28.999***	20.994*	23.432**	11.094	126.512***
	Cointegration									
ARDL-M	T-stat$_{\rho=0}$	-2.823								
model	F-stat$_{\theta=\rho=0}$	3.752								
	Cointegration									
NARDL-M	T-stat$_{\rho=0}$	-1.787								
model	F-stat$_{\rho=\theta^+=\theta^-=0}$	3.477*								
	Cointegration									
District of	**Columbia(DC)**									
QARDL	T-stat$_{\rho_*(\tau)=0}$	-5.350***	-3.454	-4.180*	-3.114	-3.172	-3.316	-3.633*	-2.342	-0.817
model	W-stat$_{\theta_*(\tau)=\rho_*(\tau)=0}$	54.292***	13.142***	21.983***	10.048**	24.037***	20.255	47.917***	29.400***	9.677**
	Cointegration	**Y**		**Y**						
NARDL-Q	T-stat$_{\rho_*(\tau)=0}$	-6.503***	-3.723	-2.670	-2.855	-3.601	-3.358	-3.651	-2.840	-1.999
model	W-stat$_{\rho_*(\tau)=\theta_*^+(\tau)=\theta_*^-(\tau)=0}$	474.931***	47.601***	30.722***	55.990***	41.547***	25.733***	81.858***	24.842***	61.239***
	Cointegration	**Y**								
ARDL-M	T-stat$_{\rho=0}$	-3.122								
model	F-stat$_{\theta=\rho=0}$	2.982								
	Cointegration									
NARDL-M	T-stat$_{\rho=0}$	-4.303*								
model	F-stat$_{\rho=\theta^+=\theta^-=0}$	3.484*								
	Cointegration	**Y**								

Table 8.11 – Cointegration between US house prices and economic fundamentals.

y-x		$\tau = 0.1$	$\tau = 0.2$	$\tau = 0.3$	$\tau = 0.4$	$\tau = 0.5$	$\tau = 0.6$	$\tau = 0.7$	$\tau = 0.8$	$\tau = 0.9$
Florida(FL)										
QARDL	T-stat$_{\rho_*(\tau)=0}$	-2.864	-3.041	-3.082	2.163	-2.427	-2.589	-1.727	-0.862	-0.512
model	W-stat$_{\theta_*(\tau)=\rho_*(\tau)=0}$	119.079***	41.574***	17.730***	10.679**	10.515**	7.187	5.027	9.978**	97.482***
	Cointegration									
NARDL-Q	T-stat$_{\rho_*(\tau)=0}$	-7.671***	-4.432***	-3.599	-2.792	-1.666	-2.383	-1.106	-0.244	-0.775
model	W-stat$_{\rho_*(\tau)=\theta_*^+(\tau)=\theta_*^-(\tau)=0}$	161.329***	56.351***	38.881***	19.969***	10.916	20.656***	10.314	51.599***	53.825***
	Cointegration	**Y**	**Y**							
ARDL-M	T-stat$_{\rho=0}$	-2.878								
model	F-stat$_{\theta=\rho=0}$	3.288								
	Cointegration									
NARDL-M	T-stat$_{\rho=0}$	-3.185								
model	F-stat$_{\rho=\theta^+=\theta^-=0}$	2.993								
	Cointegration									
Hawaii(HI)										
QARDL	T-stat$_{\rho_*(\tau)=0}$	1.623	-0.654	-1.428	-0.850	-1.969	-2.885	-3.474*	-1.579	-0.809
model	W-stat$_{\theta_*(\tau)=\rho_*(\tau)=0}$	16.065***	16.779**	12.735**	3.366	7.455	19.329***	43.422***	36.632***	1.009
	Cointegration									
NARDL-Q	T-stat$_{\rho_*(\tau)=0}$	0.783	0.134	-0.282	0.188	0.003	-2.844	-2.322	-1.823	-3.739
model	W-stat$_{\rho_*(\tau)=\theta_*^+(\tau)=\theta_*^-(\tau)=0}$	10.404	62.188***	20.713***	29.256***	44.461***	37.630***	33.578***	18.597***	82.679***
	Cointegration									
ARDL-M	T-stat$_{\rho=0}$	-0.854								
model	F-stat$_{\theta=\rho=0}$	1.623								
	Cointegration									
NARDL-M	T-stat$_{\rho=0}$	-1.688								
model	F-stat$_{\rho=\theta^+=\theta^-=0}$	1.479								
	Cointegration									

Table 8.12 – Cointegration between US house prices and economic fundamentals.

y-x		$\tau = 0.1$	$\tau = 0.2$	$\tau = 0.3$	$\tau = 0.4$	$\tau = 0.5$	$\tau = 0.6$	$\tau = 0.7$	$\tau = 0.8$	$\tau = 0.9$
Massachusetts(MA)										
QARDL	T-stat$_{\rho_*(\tau)=0}$	-1.131	0.149	0.786	0.093	0.403	-0.839	-1.574	-0.800	-1.391
model	W-stat$_{\theta_*(\tau)=\rho_*(\tau)=0}$	50.725***	30.344***	14.938***	36.239***	20.340***	63.659***	53.510***	46.618***	52.655***
	Cointegration									
NARDL-Q	T-stat$_{\rho_*(\tau)=0}$	-1.714	-1.221	0.001	-0.290	-0.398	0.085	-0.416	-0.871	-2.203
model	W-stat$_{\rho_*(\tau)=\theta_*^+(\tau)=\theta_*^-(\tau)=0}$	130.213***	152.969***	37.501***	28.474***	32.677***	35.362***	59.840***	59.801***	92.607***
	Cointegration									
ARDL-M	T-stat$_{\rho=0}$	-0.115								
model	F-stat$_{\theta=\rho=0}$	6.171***								
	Cointegration									
NARDL-M	T-stat$_{\rho=0}$	-1.313								
model	F-stat$_{\rho=\theta^+=\theta^-=0}$	4.654***								
	Cointegration									
Nevada(NV)										
QARDL	T-stat$_{\rho_*(\tau)=0}$	-3.657	-1.765	-2.785	-1.820	-1.621	-4.209***	-4.249***	-5.731***	-3.631*
model	W-stat$_{\theta_*(\tau)=\rho_*(\tau)=0}$	244.461***	9.002*	13.083**	20.280***	19.941***	38.503***	31.092***	34.651***	14.213***
	Cointegration						**Y**	**Y**	**Y**	**Y**
NARDL-Q	T-stat$_{\rho_*(\tau)=0}$	-2.213	2.652	-2.321	-3.440	-4.326*	-4.344*	-3.571	-3.646	-8.153***
model	W-stat$_{\rho_*(\tau)=\theta_*^+(\tau)=\theta_*^-(\tau)=0}$	106.990***	27.196***	54.306***	27.524***	40.733***	33.615***	41.460***	35.857***	439.670***
	Cointegration					**Y**	**Y**			**Y**
ARDL-M	T-stat$_{\rho=0}$	-3.967*								
model	F-stat$_{\theta=\rho=0}$	5.466***								
	Cointegration	**Y**								
NARDL-M	T-stat$_{\rho=0}$	-4.579**								
model	F-stat$_{\rho=\theta^+=\theta^-=0}$	4.147**								
	Cointegration	**Y**								

Table 8.13 – Cointegration between US house prices and economic fundamentals.

y-x		$\tau=0.1$	$\tau=0.2$	$\tau=0.3$	$\tau=0.4$	$\tau=0.5$	$\tau=0.6$	$\tau=0.7$	$\tau=0.8$	$\tau=0.9$
New York(NY)										
QARDL	T-stat$_{\rho_*(\tau)=0}$	-0.127	-0.659	-1.281	-1.149	-2.588	-0.940	-0.407	1.005	0.450
model	W-stat$_{\theta_*(\tau)=\rho_*(\tau)=0}$	6.402	5.603	12.235**	32.069***	61.264***	22.425***	36.478***	30.001***	161.196***
	Cointegration									
NARDL-Q	T-stat$_{\rho_*(\tau)=0}$	-1.330	-0.857	-3.351	-2.618	-2.169	-1.596	0.488	1.202	-0.239
model	W-stat$_{\rho_*(\tau)=\theta_*^+(\tau)=\theta_*^-(\tau)=0}$	387.542***	20.082***	81.377***	17.505**	28.690***	33.737***	60.608***	25.762***	221.424***
	Cointegration									
ARDL-M	T-stat$_{\rho=0}$	-0.173								
model	F-stat$_{\theta=\rho=0}$	3.826*								
	Cointegration									
NARDL-M	T-stat$_{\rho=0}$	-0.861								
model	F-stat$_{\rho=\theta^+=\theta^-=0}$	3.337*								
	Cointegration									
Pennsylvania(PA)										
QARDL	T-stat$_{\rho_*(\tau)=0}$	-4.070***	-3.414	-3.980***	-2.778	-2.305	-1.566	-0.785	-1.672	-3.022
model	W-stat$_{\theta_*(\tau)=\rho_*(\tau)=0}$	23.781***	16.617***	17.625***	8.172*	6.776	6.048	1.342	5.761	21.711***
	Cointegration	**Y**		**Y**						
NARDL-Q	T-stat$_{\rho_*(\tau)=0}$	-14.650***	-5.793***	-7.054***	-7.724***	-6.827***	-8.485***	-7.171***	-5.008***	-9.232***
model	W-stat$_{\rho_*(\tau)=\theta_*^+(\tau)=\theta_*^-(\tau)=0}$	1416.190***	68.166***	79.911***	99.741***	55.376***	98.132***	90.309***	72.031***	325.159***
	Cointegration	**Y**	**Y**	**Y**	**Y**	**Y**	**Y**	**Y**	**Y**	**Y**
ARDL-M	T-stat$_{\rho=0}$	-2.476								
model	F-stat$_{\theta=\rho=0}$	2.348								
	Cointegration									
NARDL-M	T-stat$_{\rho=0}$	-5.791***								
model	F-stat$_{\rho=\theta^+=\theta^-=0}$	5.766***								
	Cointegration	**Y**								

8.6 Conclusions

We investigate the long-run equilibrium relationship between US house prices and economic fundamentals at the both national and selected State levels from 1978Q1 to 2016Q4 using several ARDL based cointegration tests. The presence of cointegration between house prices and fundamentals suggests a stable equilibrium relationship, indicating that a temporary increase or drop in house prices would eventually come back to an equilibrium. If house prices and fundamentals are not cointegrated, then the error-correction specifications are inappropriate. In this case, house prices do not have to stagnate or fall to maintain an equilibrium relationship even if house prices outpace fundamentals.

Both the conditional mean- and quantile-based cointegration test of ARDL-M of Pesaran et al. (2001) and NARDL-M of Shin et al. (2014) are applied to examine the potential symmetric or asymmetric cointegration relationship. A crucial contribution to the literature is that two recently proposed quantile-based ARDL models are employed to test for cointegration across the whole conditional distribution. Based on the symmetric quantile ARDL model (QARDL) and the non-asymmetric quantile ARDL model (NARDL-Q), two testing procedures for examining a long-run cointegration relationship are utilised in this study as suggested by Prof Yongcheol Shin. These two testing procedures are the t-statistic of Banerjee et al. (1998) and a Wald-test for the joint null.

Some key findings have been presented in this study. The results above show no cointegration between house prices and fundamentals at the national level based on both conditional mean- and quantile-based cointegration tests. This finding is consistent with the early view of Gallin (2006), who also identify non-cointegrating relationship at the aggregate level. As argued by Gallin (2006), it would be inappropriate to model house price dynamics using an error-correction specification. Hence those studies on house price dynamics of Abraham and Hendershott (1996), Capozza et al. (2002) and Malpezzi (1999) based on error correction specifications are likely misspecified. Empirical results show, under certain cases, that cointegration may be established based on the conditional mean-based model (e.g., ARDL-M or NARDL-M models) but there could be no overwhelming evidence of cointegration based on the quantile-based model (e.g., QARDL or NARLD-Q models). This particular phenomenon has been observed in the District of Columbia and Nevada. The OLS estimation procedure assumes that the estimated relationship holds not only at the mean but in other parts of the conditional distribution as well. On the other hand, the conditional mean based cointegration model may be unable to identify a stable long-run relationship while the quantile-based model may uncover some evidence of cointegration at certain quantiles. This phenomenon has been observed in District of Columbia and Florida. Based on both conditional mean and quantile models, several examples demonstrate that there is no stable long-run relationship between house prices and fundamentals at the national and state levels including California, Hawaii, Massachusetts and New York. This finding is of great interest as it provides strong evidence of a non-cointegrating relationship. The corresponding results for Pennsylvania show that house prices and fundamentals are asymmetrically cointegrated based on the NARDL-M model. Such asymmetric cointegration relationship is also confirmed by the NARDL-Q model, where the relationship holds at the whole conditional distribution. It is quite a remarkable finding for Pennsylvania.

We re-visit the hypothesis between US house prices and the relevant economic fundamentals using advanced econometric methodologies. As mentioned above, this research reports several interesting findings. The findings from this paper indeed offer new meaningful insights from the point view of methodology development and empirical evidence. Apart from two conditional-mean based cointegration models, two quantile-based ARDL cointegration models essentially allow policy-makers to understand house price dynamics under different phases (e.g., a bear, normal and boom market). There have been no studies modelling a long-run relationship between house prices and fundamentals based on a quantile cointegration approach in the prior literature. The conclusion provides further evidence of a cointegration or non-cointegration relationship for the national and several key

States.

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8.A Appendix

Table 8.14 – Unit root tests for the national level.

Variables/test	ADF (P-value)		PP (P-value)		KPSS (Stat)		QKS(Stat)
US	C	C+T	C	C+T	C	C+T	
HP (level)	0.2907	0.0224**	0.7013	0.5947	0.9667***	0.0958	2.3724
(First Difference)	0.0991*	0.2776	0.0000***	0.0001***	0.0699	0.0652	⇓
	I(1)	I(0)	I(1)	I(1)	I(1)	I(0)	I(1)
IC (Level)	0.8665	0.8402	0.8707	0.7269	1.5089***	0.2265***	0.6523
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.0940	0.0693	⇓
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
CC (Level)	0.3685	0.2296	0.6106	0.6834	0.4851**	0.1095	2.6817
(First Difference)	0.0005***	0.0018***	0.0000***	0.0000***	0.1311	0.0602	⇓
	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)	I(1)
MR (Level)	0.7108	0.0191**	0.3703	0.0001***	1.2397***	0.0701	1.3482
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.1131	0.0589	⇓
	I(1)	I(0)	I(1)	I(0)	I(1)	I(0)	I(1)
ICP (Level)	0.9634	0.5368	0.9612	0.3757	1.5119***	0.1033	0.8897
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.0518	0.0530	⇓
	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)	I(1)
ICN (Level)	0.9923	0.9749	0.9858	0.9532	1.2609***	0.2737***	3.6952
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.2481	0.1258	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)
CCP (Level)	1.0000	0.9998	1.0000	0.9999	1.4667***	0.3352**	0.6048
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.7364**	0.1064	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(2)	I(1)
CCN (Level)	0.8596	0.2927	0.9856	0.8492	1.3970***	0.2153**	0.3290
(First Difference)	0.0578*	0.2129	0.0000***	0.0000***	0.1643	0.0762	
	I(1)	I(2)	I(1)	I(1)	I(1)	I(1)	I(1)
MRP (Level)	0.9918	0.9546	0.9922	0.9529	1.4693***	0.3002***	0.5757
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.2437	0.1334*	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(2)	I(1)
MRN (Level)	0.9990	0.8876	0.9995	0.9091	1.4769***	0.2973***	0.9725
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.3865*	0.0815	
	I(1)	I(1)	I(1)	I(1)	I(2)	I(1)	I(1)

Table 8.15 – Unit root tests for Arizona (AZ).

Variables	ADF (P-value)		PP (P-value)		KPSS (Stat)		QKS (Stat)
Arizona (AZ)	C	C+T	C	C+T	C	C+T	
HP (level)	0.1087	0.1753	0.3228	0.5163	0.3512*	0.1149	4.4617
(First Difference)	0.0027***	0.0147**	0.0027***	0.0150**	0.0609	0.0513	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)	I(0)
IC (Level)	0.7752	0.6788	0.4531	0.8612	1.5032***	0.2270***	1.4039
(First Difference)	0.0081***	0.0347**	0.0000***	0.0000***	0.2371	0.0565	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
CC (Level)	0.5358	0.1617	0.7791	0.5983	0.8720***	0.1291*	4.4367
(First Difference)	0.0089***	0.0315**	0.0000***	0.0002***	0.1245	0.0597	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)
MR (Level)	0.8575	0.0126**	0.3709	0.0001***	1.2439***	0.069592	1.6025
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.1308	0.093022	
	I(1)	I(2)	I(1)	I(2)	I(1)	I(0)	I(1)
ICP (Level)	0.2828	0.4659	0.5328	0.6652	1.5108***	0.1731**	1.9726
(First Difference)	0.0046***	0.0000***	0.0000***	0.0000***	0.2049	0.0441	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
ICN (Level)	0.9501	0.8357	0.9867	0.9330	1.2299***	0.2522***	4.0128
(First Difference)	0.0019***	0.0092***	0.0000***	0.0000***	0.2202	0.0982	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)
CCP (Level)	0.9993	0.9117	0.9999	0.9627	1.4575***	0.3167***	2.1131
(First Difference)	0.0000***	0.0001***	0.0000***	0.0000***	0.4845	0.0534	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
CCN (Level)	0.8814	0.5761	0.9583	0.8124	1.3293***	0.2125	0.2858
(First Difference)	0.0036***	0.0199**	0.0000***	0.0000***	0.1247	0.0771	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)	I(1)
MRP (Level)	0.9788	0.8914	0.9908	0.9617	1.4649***	0.3075***	1.3141
(First Difference)	0.0011***	0.0062***	0.0000***	0.0000***	0.2521	0.1565**	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(2)	I(1)
MRN (Level)	0.9994	0.9516	0.9998	0.9695	1.4774***	0.3248	1.0784
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.4723**	0.0927	
	I(1)	I(1)	I(1)	I(1)	I(2)	I(0)	I(1)

Table 8.16 – Unit root tests for California (CA).

Variables	ADF (P-value)		PP (P-value)		KPSS (Stat)		QKS (Stat)
California (CA)	C	C+T	C	C+T	C	C+T	
HP (level)	0.1800	0.0094***	0.6423	0.5337	0.9522***	0.0760	3.8098
(First Difference)	0.0391**	0.1428	0.0068***	0.0337**	0.0586	0.0556	
	I(1)	I(0)	I(1)	I(1)	I(1)	I(0)	I(0)
IC (Level)	0.8883	0.7784	0.8868	0.5624	1.5026***	0.1403*	1.6149
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.0702	0.0610	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
CC (Level)	0.5358	0.1617	0.7791	0.5983	0.8720***	0.1291	4.4367
(First Difference)	0.0089***	0.0315**	0.0000***	0.0002***	0.1245	0.0597	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)	I(0)
MR (Level)	0.8575	0.0126**	0.3709	0.0001***	1.2439***	0.0696	1.6025
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.1308	0.0930	
	I(1)	I(0)	I(1)	I(0)	I(1)	I(0)	I(1)
ICP (Level)	0.9904	0.0028***	0.9831	0.6713	1.5041***	0.0769	1.4845
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.0977	0.0742	
	I(1)	I(0)	I(1)	I(1)	I(1)	I(0)	I(1)
ICN (Level)	0.9988	0.9692	0.9984	0.9631	1.3099***	0.3016***	0.3025
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.4007*	0.1000	
	I(1)	I(1)	I(1)	I(1)	I(2)	I(1)	I(1)
CCP (Level)	0.9993	0.9117	0.9999	0.9627	1.4575***	0.3167***	2.1131
(First Difference)	0.0000***	0.0001***	0.0000***	0.0000***	0.4845**	0.0534	
	I(1)	I(1)	I(1)	I(1)	I(2)	I(1)	I(1)
CCN (Level)	0.8814	0.5761	0.9583	0.8124	1.3293***	0.2125**	0.2858
(First Difference)	0.0036***	0.0199**	0.0000***	0.0000***	0.1247	0.0771	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
MRP (Level)	0.9788	0.8914	0.9908	0.9617	1.4649***	0.3075***	1.3141
(First Difference)	0.0011***	0.0062***	0.0000***	0.0000***	0.2521	0.1565**	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(2)	I(1)
MRN (Level)	0.9994	0.9516	0.9998	0.9695	1.4774***	0.3248***	1.0784
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.4723**	0.0927	
	I(1)	I(1)	I(1)	I(1)	I(2)	I(1)	I(1)

Table 8.17 – Unit root tests for District of Columbia (DC).

Variables	ADF (P-value)		PP (P-value)		KPSS (Stat)		QKS (Stat)
District of Columbia (DC)	C	C+T	C	C+T	C	C+T	
HP (level)	0.8746	0.0772*	0.9586	0.7521	1.2819***	0.2117**	1.5591
(First Difference)	0.0631*	0.1532	0.0002***	0.0007***	0.1623	0.0719	
	I(1)	I(0)	I(1)	I(1)	I(1)	I(1)	I(1)
IC (Level)	0.9999	0.1831	0.9998	0.1500	1.4702***	0.2296***	1.5436
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.4846**	0.0729	
	I(1)	I(1)	I(1)	I(1)	I(2)	I(1)	I(1)
CC (Level)	0.2432	0.1685	0.5597	0.5446	0.9019***	0.0803	1.4077
(First Difference)	0.0000***	0.0002***	0.0000***	0.0002***	0.0651	0.0643	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)	I(1)
MR (Level)	0.6892	0.0003***	0.5176	0.0006***	1.2494***	0.0721	2.5637
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.0716	0.0474	
	I(1)	I(0)	I(1)	I(0)	I(1)	I(0)	I(1)
ICP (Level)	0.9995	0.9274	0.9996	0.9125	1.5062***	0.2934***	94.2994
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.3227	0.0502	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)
ICN (Level)	0.0025***	0.0003***	0.0477**	0.0010***	1.4990***	0.0626	1.8124
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.3246	0.1238*	
	I(0)	I(0)	I(0)	I(0)	I(1)	I(0)	I(1)
CCP (Level)	0.8205	0.2917	0.8533	0.5260	1.4728***	0.1843**	0.6075
(First Difference)	0.0000***	0.0001***	0.0000***	0.0001***	0.0681	0.0514	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
CCN (Level)	0.8231	0.1075	0.8812	0.7648	1.4347***	0.1157	0.3290
(First Difference)	0.0001***	0.0005***	0.0000***	0.0000***	0.1020	0.0984	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)	I(1)
MRP (Level)	0.8574	0.5150	0.8583	0.4207	1.4939***	0.2123**	0.6778
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.1111	0.1147	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
MRN (Level)	0.9790	0.5644	0.9829	0.6225	1.4981***	0.2253***	0.8863
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.0995	0.0562	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)

Table 8.18 – Unit root tests for Florida (FL).

Variables	ADF (P-value)		PP (P-value)		KPSS (Stat)		QKS (Stat)
Florida (FL)	C	C+T	C	C+T	C	C+T	
HP (level)	0.3511	0.0402**	0.4652	0.5866	0.4704**	0.1050	3.5686
(First Difference)	0.0008***	0.0048***	0.0004***	0.0024***	0.0731	0.0591	
	I(1)	I(0)	I(1)	I(1)	I(1)	I(0)	I(0)
IC (Level)	0.0589*	0.8726	0.1824	0.8075	1.4856***	0.2969***	2.6044
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.4048**	0.0516	
	I(0)	I(1)	I(1)	I(1)	I(2)	I(1)	I(1)
CC (Level)	0.3267	0.6976	0.5289	0.8624	0.1717	0.1727**	3.0878
(First Difference)	0.0001***	0.0005***	0.0000***	0.0000***	0.1880	0.0669	
	I(1)	I(1)	I(1)	I(1)	I(0)	I(1)	I(0)
MR (Level)	0.6949	0.0053***	0.3784	0.0000***	1.2521***	0.0698	2.3276
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.1182	0.0723	
	I(1)	I(0)	I(1)	I(0)	I(1)	I(0)	I(1)
ICP (Level)	0.3935	0.4136	0.4779	0.3654	1.5069***	0.2200***	3.6517
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.2653	0.0784	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)
ICN (Level)	1.0000	0.9899	0.9986	0.9672	1.0940***	0.3068***	39.6263
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.4890**	0.0869	
	I(1)	I(1)	I(1)	I(1)	I(2)	I(1)	I(0)
CCP (Level)	1.0000	0.9998	1.0000	0.9999	1.4261***	0.3611***	3.0128
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.7274**	0.1271*	
	I(1)	I(1)	I(1)	I(1)	I(2)	I(2)	I(0)
CCN (Level)	0.9584	0.3506	0.9631	0.7759	1.3817***	0.1754**	9.7954
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.1162	0.0781	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)
MRP (Level)	0.9905	0.9740	0.9872	0.9484	1.4609***	0.3021***	4.3629
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.2379	0.1591**	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)
MRN (Level)	0.9991	0.9607	0.9997	0.9766	1.4756***	0.3200***	1.0162
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.4380*	0.0993	
	I(1)	I(1)	I(1)	I(1)	I(2)	I(1)	I(1)

Table 8.19 – Unit root tests for Hawaii (HI).

Variables	ADF (P-value)		PP (P-value)		KPSS (Stat)		QKS (Stat)
Hawaii (HI)	C	C+T	C	C+T	C	C+T	
HP (level)	0.6892	0.3416	0.7044	0.5401	1.0676***	0.0636	1.7834
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.0588	0.0574	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)	I(1)
IC (Level)	0.9083	0.8169	0.9033	0.6992	1.4692***	0.1374*	0.7391
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.0838	0.0808	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
CC (Level)	0.5358	0.1617	0.7791	0.5983	0.8720***	0.1291*	1.6025
(First Difference)	0.0089***	0.0315**	0.0000***	0.0002***	0.1245	0.0597	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
MR (Level)	0.8575	0.0126**	0.3709	0.0001***	1.2439***	0.0696	4.4367
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.1308	0.0930	
	I(1)	I(0)	I(1)	I(0)	I(1)	I(0)	I(0)
ICP (Level)	0.8388	0.7986	0.8483	0.7290	1.4856***	0.1610**	0.7358
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.1487	0.1214*	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(2)	I(0)
ICN (Level)	0.7492	0.3922	0.7806	0.5304	1.4404***	0.1189	114.5816
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.1023	0.0980	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)	I(0)
CCP (Level)	0.9993	0.9117	0.9999	0.9627	1.4575***	0.3167***	2.1105
(First Difference)	0.0000***	0.0001***	0.0000***	0.0000***	0.4845**	0.0534	
	I(1)	I(1)	I(1)	I(1)	I(2)	I(1)	I(1)
CCN (Level)	0.8814	0.5761	0.9583	0.8124	1.3293***	0.2125**	0.2792
(First Difference)	0.0036***	0.0199**	0.0000***	0.0000***	0.1247	0.0771	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
MRP (Level)	0.9788	0.8914	0.9908	0.9617	1.4649***	0.3075***	1.3141
(First Difference)	0.0011***	0.0062***	0.0000***	0.0000***	0.2521	0.1565**	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(2)	I(1)
MRN (Level)	0.9994	0.9516	0.9998	0.9695	1.4774***	0.3248***	1.0784
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.4723**	0.0927	
	I(1)	I(1)	I(1)	I(1)	I(2)	I(1)	I(1)

Table 8.20 – Unit root tests for Massachusetts (MA).

Variables	ADF (P-value)		PP (P-value)		KPSS (Stat)		QKS (Stat)
Massachusetts (MA)	C	C+T	C	C+T	C	C+T	
HP (level)	0.1111	0.0004***	0.3412	0.5927	0.9805***	0.1167	2.7552
(First Difference)	0.0338**	0.1000	0.0000***	0.0000***	0.1232	0.0626	
	I(1)	I(0)	I(1)	I(1)	I(1)	I(0)	I(0)
IC (Level)	0.6425	0.8706	0.7270	0.6459	1.4747***	0.1826**	1.9646
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.1138	0.0527	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
CC (Level)	0.2432	0.1685	0.5597	0.5446	0.9019***	0.0803	1.4077
(First Difference)	0.0000***	0.0002***	0.0000***	0.0002***	0.0651	0.0643	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)	I(1)
MR (Level)	0.6892	0.0003***	0.5176	0.0006***	1.2494***	0.0721	2.5637
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.0715	0.0474	
	I(1)	I(0)	I(1)	I(0)	I(1)	I(0)	I(1)
ICP (Level)	0.6936	0.7699	0.7791	0.6418	1.4954***	0.1938**	1.8824
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.1292	0.0671	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
ICN (Level)	0.9809	0.7661	0.9785	0.6611	1.4554***	0.1403*	3.9730
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.1102	0.0594	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)
CCP (Level)	0.8205	0.2917	0.8533	0.5260	1.4728***	0.1843**	0.6075
(First Difference)	0.0000***	0.0001***	0.0000***	0.0001***	0.0681	0.0514	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
CCN (Level)	0.8231	0.1075	0.8812	0.7648	1.4347***	0.1157	0.3290
(First Difference)	0.0001***	0.0005***	0.0000***	0.0000***	0.1020	0.0984	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)	I(1)
MRP (Level)	0.8574	0.5150	0.8583	0.4207	1.4939***	0.2123**	0.6778
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.1111	0.1147	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
MRN (Level)	0.9790	0.5644	0.9829	0.6225	1.4981***	0.2254***	0.8863
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.0995	0.0562	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)

Table 8.21 – Unit root tests for Nevada (NV).

Variables	ADF (P-value)		PP (P-value)		KPSS (Stat)		QKS (Stat)
Nevada (NV)	C	C+T	C	C+T	C	C+T	
HP (level)	0.0006***	0.0039***	0.1957	0.4733	0.0825	0.0844	3.1655
(First Difference)	0.0333**	0.1207	0.0103**	0.0469**	0.0551	0.0527	
	I(0)	I(0)	I(1)	I(1)	I(0)	I(0)	I(0)
IC (Level)	0.6426	0.9643	0.4171	0.9886	1.4815***	0.2938***	2.2456
(First Difference)	0.0151**	0.0411**	0.0000***	0.0000***	0.4074*	0.1525**	
	I(1)	I(1)	I(1)	I(1)	I(2)	I(2)	I(1)
CC (Level)	0.5358	0.1617	0.7791	0.5983	0.8720***	0.1291*	4.4367
(First Difference)	0.0089***	0.0315**	0.0000***	0.0002***	0.1245	0.0597	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)
MR (Level)	0.8575	0.0126**	0.3709	0.0001***	1.2439***	0.0696	1.6025
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.1308	0.0930	
	I(1)	I(0)	I(1)	I(0)	I(1)	I(0)	I(0)
ICP (Level)	0.6987	0.9464	0.6837	0.9782	1.5056***	0.2706***	1.7532
(First Difference)	0.0058***	0.0208**	0.0000***	0.0000***	0.3093	0.1677**	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
ICN (Level)	0.9998	0.9915	0.9988	0.9789	1.0857***	0.3124***	3.7955
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.4911**	0.0934	
	I(1)	I(1)	I(1)	I(1)	I(2)	I(1)	I(0)
CCP (Level)	0.9993	0.9117	0.9999	0.9627	1.4575***	0.3167***	2.1131
(First Difference)	0.0000***	0.0001***	0.0000***	0.0000***	0.4845**	0.0533	
	I(1)	I(1)	I(1)	I(1)	I(2)	I(1)	I(1)
CCN (Level)	0.8814	0.5761	0.9583	0.8124	1.3293***	0.2125**	0.2858
(First Difference)	0.0036***	0.0199**	0.0000***	0.0000***	0.1247	0.0771	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
MRP (Level)	0.9788	0.8914	0.9908	0.9617	1.4649***	0.3075***	1.3141
(First Difference)	0.0011***	0.0062***	0.0000***	0.0000***	0.2521	0.1565**	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(2)	I(1)
MRN (Level)	0.9994	0.9516	0.9998	0.9695	1.4774***	0.3248***	1.0784
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.4723**	0.0927	
	I(1)	I(1)	I(1)	I(1)	I(2)	I(1)	I(1)

Table 8.22 – Unit root tests for New York (NY).

Variables	ADF (P-value)		PP (P-value)		KPSS (Stat)		QKS (Stat)
New York (NY)	C	C+T	C	C+T	C	C+T	
HP (level)	0.0448**	0.0001***	0.4909	0.6933	0.9671***	0.1100	2.2059
(First Difference)	0.0376**	0.1095	0.0000***	0.0000***	0.0986	0.0639	
	I(0)	I(0)	I(1)	I(1)	I(1)	I(0)	I(1)
IC (Level)	0.8008	0.7124	0.8016	0.6299	1.4655***	0.2179***	1.6610
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.0958	0.0577	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
CC (Level)	0.2432	0.1685	0.5597	0.5446	0.9019***	0.0803	1.4077
(First Difference)	0.0000***	0.0002***	0.0000***	0.0002***	0.0651	0.0643	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)	I(1)
MR (Level)	0.6892	0.0003***	0.5176	0.0006***	1.2494***	0.0721	2.5637
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.0716	0.0474	
	I(1)	I(0)	I(1)	I(0)	I(1)	I(0)	I(1)
ICP (Level)	0.9373	0.4185	0.9378	0.3557	1.5040***	0.1616**	2.0556
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.0509	0.0501	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
ICN (Level)	0.9888	0.7140	0.9890	0.6842	1.4657***	0.2423***	62.119
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.1706	0.0620	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)
CCP (Level)	0.8205	0.2917	0.8533	0.5260	1.4728***	0.1843**	0.6075
(First Difference)	0.0000***	0.0001***	0.0000***	0.0001***	0.0681	0.0514	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
CCN (Level)	0.8231	0.1075	0.8812	0.7648	1.4347***	0.1157	0.3290
(First Difference)	0.0001***	0.0005***	0.0000***	0.0000***	0.1020	0.0984	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)	I(1)
MRP (Level)	0.8574	0.5150	0.8583	0.4207	1.4939***	0.2123**	0.6778
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.1111	0.1147	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
MRN (Level)	0.9790	0.5644	0.9829	0.6225	1.4981***	0.2254***	0.8863
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.0995	0.0562	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)

Table 8.23 – Unit root tests for Pennsylvania (PA).

Variables	ADF (P-value)		PP (P-value)		KPSS (Stat)		QKS (State)
Pennsylvania (PA)	C	C+T	C	C+T	C	C+T	
HP (level)	0.0983*	0.0016***	0.5988	0.5197	0.9289***	0.0803	2.7819
(First Difference)	0.0388**	0.1437	0.0000***	0.0000***	0.0649	0.0655	
	I(0)	I(0)	I(1)	I(1)	I(1)	I(0)	I(0)
IC (Level)	0.9329	0.2565	0.9383	0.3114	1.5068***	0.1695**	1.2298
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.0556	0.0582	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
CC (Level)	0.2432	0.1685	0.5597	0.5446	0.9019***	0.0803	1.4077
(First Difference)	0.0000***	0.0002***	0.0000***	0.0002***	0.0651	0.0643	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)	I(1)
MR (Level)	0.6892	0.0003***	0.5176	0.0006***	1.2494***	0.0721	2.5637
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.0716	0.0474	
	I(1)	I(0)	I(1)	I(0)	I(1)	I(0)	I(1)
ICP (Level)	0.8898	0.7845	0.8890	0.7814	1.4883***	0.2252***	1.7496
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.1211	0.0990	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
ICN (Level)	0.5695	0.6795	0.5818	0.6582	1.3656***	0.1944**	3.1675
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.2491	0.1859**	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)	I(0)
CCP (Level)	0.8205	0.2917	0.8533	0.5260	1.4728***	0.1843**	0.6075
(First Difference)	0.0000***	0.0001***	0.0000***	0.0001***	0.0681	0.0514	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
CCN (Level)	0.8231	0.1075	0.8812	0.7648	1.4347***	0.1157	0.3290
(First Difference)	0.0001***	0.0005***	0.0000***	0.0000***	0.1020	0.0984	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)	I(1)
MRP (Level)	0.8574	0.5150	0.8583	0.4207	1.4939***	0.2123**	0.6778
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.1111	0.1147	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)
MRN (Level)	0.9790	0.5644	0.9829	0.6225	1.4981***	0.2254***	0.8863
(First Difference)	0.0000***	0.0000***	0.0000***	0.0000***	0.0995	0.0562	
	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)

Chapter 9

Conclusion

9.1 Overview

The notion of bubbles has attracted considerable interest over time with a resurgence of following the recent global financial crisis. This thesis has investigated the phenomenon of financial bubbles in different markets using extended period of time from the early 18th century to the 21st century. The purpose of this thesis is to examine a range of famous bubble episodes in history using some recently proposed bubble detection tests. In particular, I consider the historical house price series in Amsterdam, Norway and France, the Mississippi and South Sea Bubbles in 1719-20, Japan's asset price bubble during the late 1980s and the US real estate bubble in the 2000s. The existing literature mainly focuses on detecting some recent financial bubbles while few studies attempts to take into account historical bubbles. This thesis, therefore, seeks to fill the existing gap in the literature by testing for the empirical and robust evidence of some widely recognised bubbles which have typically not been subjected to modern econometric testing approaches. Empirical results for the above episodes are presented as Chapters 3 through to 7. This thesis also considers the long-run relationship between US house prices and the relevant economic fundamentals using advanced quantile cointegration approaches with the results presented in Chapter 8.

9.1.1 Key Results

Section 1.2 and particularly between pages 10 and 12 identified six key questions to be addressed in the thesis. Below, I consider what the work presented in the thesis concludes in relation to these, and other questions related to bubbles. Since the excellent work on exchange rate bubbles by, for example, Huang (1981), Evans (1986), Meese (1986), West (1987) and Wu (1995), there have been many empirical studies which focus on this topic. Chapter 3 explored the presence of bubbles in some G10 and a range of emerging markets countries (including some Asian and the BRICS). This chapter applied the popular and well-researched method of PSY under different regression models to the widest and most extensive range of exchange rates currently undertaken. Empirical results from Chapter 3 seem to suggest that emerging market economies are more likely to exhibit bubble-like behavior. Moreover, date-stamping results are sensitive to different model specifications.

Unlike most existing studies on housing bubbles using a short period time series for few decades, Chapter 4 investigated the bubble-like behavior in several well-constructed and regarded house price series; the Herengracht index for Amsterdam (1649-2010), Norway (1819-2014) and Paris (1650-2012). These historical series comprise the few long-term house price series in the literature and represent different types of real estate markets in terms of their coverage. Chapter 4 confirms no explosive behavior in the Herengracht index for Amsterdam (1649-2010). This finding is perhaps not surprising as Shiller (2006) concludes that Herengracht index doesn't offer very positive real returns, suggesting an annual average price increase of only 0.2% a year. Chapter 4 also presents evidence of explosive behavior in historical Norwegian house price at the aggregate and disaggregate levels. Perhaps more importantly and interestingly, the identified episodes generally coincide or overlap with several major financial crises in Norway as discussed in Grytten and Hunnes (2010), who conclude in favour of nine 'devastating financial crises' in almost 200 years of Norwegian history. Moreover, empirical results also suggest that the house price episodes identified from four Norwegian main cities (Oslo, Bergen, Trondheim and Kristiansand) are generally in line with those episodes obtained from the real house price index at the national level. The real house price index of Paris exhibits significant evidence of exuberance under a particular model specification only. Bubble detection results from PSY are generally in line with those obtained from long memory models based on the mean value of d.

US real estate prices have drawn a lot of attention as it is widely regarded as the most speculative housing market. Chapter 5 finds evidence of housing bubbles in several US States in the 1980s (i.e., California, Hawaii, Massachusetts, New York etc.), which coincide with some existing studies that investigate housing bubbles or booms and busts using a range of alternative approaches. Moreover, Chapter 5 also identifies, for the first time, the existence of a housing bubble that originates in the early 2000s and collapses in the mid-2000s in more than 20 States and the District of Columbia (DC). This finding seems to be in agreement with the talk given by Alan Greenspan in 2005, who suggests that there was no sign of a nationwide housing bubble. Moreover, the bubbles of the 2000s were more widespread than the 1980s, which is of special interest and importance.

The Mississippi Bubble and South Sea Bubbles are widely believed to be most speculative episodes and earliest examples in stock markets. Chapter 6 focuses on the famous Mississippi Bubble in French and the South Sea Bubble in England in the early 18th century. This chapter investigates bubble-like behaviour in the Mississippi Company and South Sea Company, and a selection of other financial organisations in England during the period 1719-20 using the PSY approach. In particular, evidence of explosive behavior (exuberance) in share prices is confirmed for the South Sea Company for the first time, and also established for a number of other 18th century British financial organisations, for example, East India Company, London Assurance, Million Bank, the Royal African Company and the Royal Exchange Assurance. However, there is little evidence of exuberance for the Bank of England. The South Sea Company was not the only company experiencing exuberance in its share price for this particular period in the British share market. The timings of these bubble episodes show signs of possible contagion during the South Sea episode.

Japan's asset price bubble of the late 1980s to early 1990s is widely regarded as one of the most famous bubble episodes in the history. Chapter 7 provides empirical evidence to support Japan's recent asset price bubble in its most inflated stock and real estate markets during the 1980s-90s. For both markets, this chapter reports the origination and collapse dates of so-called 'double bubbles', which coincide with the traditional view of the well-known Japan's asset price bubble. By utilising a recently developed time-varying regression based contagion methodology of Greenaway-McGrevy and Phillips (2016), such a new contagion procedure clearly demonstrates signs of bubble migration from the stock market (core market) to the real estate market for the first time in the literature. The greatest contribution of this chapter is that the bubble-like behaviour from the stock market not only migrates to, but also fuels the booming real estate market.

An important question has been researched extensively in the literature is whether US house prices reflect economic fundamentals. Chapter 8 tests for a long-run cointegration relationship between house prices and economic fundamentals (e.g., personal income, construction costs, mortgage rates) using the US aggregate and disaggregate data employ by both conventional cointegration tests based on conditional mean models of Pesaran et al. (2001) and Shin et al. (2014) and quantile cointegration models of Cho et al. (2015) and Greenwood-Nimmo et al. (2011). Firstly, this chapter reports no cointegration between house prices and fundamentals at the national level based on both conditional mean- and quantile-based cointegration tests. Secondly and more importantly, empirical results show, under certain cases, that cointegration may be established based on conditional mean-based model (e.g., ARDL-M model of Pesaran et al. (2001) or NARDLM model of Shin et al. (2014)) but there is no overwhelming evidence of cointegration based on the quantile-based model (e.g., QARDL model of Cho et al. (2015) or NARLD-Q model of Greenwood-Nimmo et al. (2011)), and vice versa. Thirdly, several examples show that there is no long-run relationship based on both conditional mean- and quantile-based cointegration tests. The main contribution of this chapter is to provide further evidence of a cointegration or non-cointegration relationship for the US and several selected states based on new econometric approaches.

9.2 Future Work

There are several possible avenues for future research related to this thesis.

First, an alternative approach to exploring the presence of bubble-like behaviour for historical bubble episodes is of particularly interest and importance. This thesis examines the presence of historical financial bubbles in the 18th and 19th centuries using the right-tailed unit root tests of Phillips, Shi, and Yu (2015, PSY). Conclusions from Chapter 6 provide econometric-based evidence to support the existence of the Mississippi Bubble, the South Sea Bubble and the British Railway Mania. Despite these novel contributions to the literature, there have been few studies to examine those well-documented bubble episodes. The PSY testing procedure is often applied to a price-fundamental ratio to assess explosive behaviour. However, the fundamental series is rarely available for these historical episodes. Markov-switching unit root test for detecting historical bubbles will be a great addition to the growing literature. For example, the Markov-switching unit root tests of Hall et al. (1999) van Norden and Vigfusson (1998) can be used for identifying bubbles. The two-regime Markov-switching augmented Dickey-Fuller (ADF) test of Hall et al. (1999) assumes that the error variances of the two regimes are identical. On the other hand, the model of van Norden and Vigfusson (1998) assumes time-varying error variance. Shi (2013) suggests some empirical guidelines for the practical implementation of the Markov-switching unit root test proposed by Hall et al. (1999) for detecting explosive bubble behavior. Interested readers may refer to this article for details.

Second, research on bubble contagion across markets has been relatively scarce. Chapter 7 documents overwhelming evidence in Japan's stock and real estate markets between the 1980s and the early 1990s and signs of spillovers from stock market to its real estate market for the first time in the literature using the right-tailed unit root test of Phillips et al. (2015) and time-varying regression methodology of Greenaway-McGrevy and Phillips (2016). A potential future topic is to apply Greenaway-McGrevy and Phillips's (2016) approach to other applications. China's real estate market is widely believed to contain a bubble. House prices in China have achieved rapid and remarkable growth. For example, average real housing prices have grown at an annual rate of around 17% over the past decade, which was higher than the average income growth rate of 11% across the thirty-five major cities and the nation's 10% average GDP growth (Chen and Wen, 2017). One may consider identifying real estate bubbles in China at the city level and look for evidence of bubble migration from four first-tier cities (Beijing, Shanghai, Guangzhou and Shenzhen) to the surrounding areas. The price-income ratio and price-rent ratio in these cities are extremely high.

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9.A Appendix: Co-Authorship Form

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Co-Authorship Form

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Chapter 3

Hu, Y., & Oxley, L. (2017). Are there bubbles in exchange rates? Some new evidence from G10 and emerging market economies. Economic Modelling, 64, 419-442.

Nature of contribution by PhD candidate	Performing empirical analysis, data collection, writing the initial draft.
Extent of contribution by PhD candidate (%)	80%

CO-AUTHORS

Name	Nature of Contribution
Les Oxley	DISCUSSIONS ON ECONOMETRIC METHODS TO USE &
	INTERPRETATIONS OF INITIAL + FINAL RESULTS.
	COMPLETION OF FINAL DRAFT. SUBMISSION &
	CONTRIBUTIONS TO RESPONSES TO REFEREES
	& EDITOR

Certification by Co-Authors

The undersigned hereby certify that:

- the above statement correctly reflects the nature and extent of the PhD candidate's contribution to this work, and the nature of the contribution of each of the co-authors; and

Name	Signature	Date
Yang Hu	[signature]	28/08/18.
Les Oxley	[signature]	22/8/18

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Chapter 4

Hu, Y., & Oxley, L. (2017). Exuberance, Bubbles or Froth? Some results using very long run historical house price data for Amsterdam, Norway and Paris. (Under review at Empirical Economics) Revise + resubmit LO

Nature of contribution by PhD candidate	Performing empirical analysis and writing the initial draft.
Extent of contribution by PhD candidate (%)	80%

CO-AUTHORS

Name	Nature of Contribution
Les Oxley	SUGGESTED APPLICATION & IDENTIFIED WHICH DATA NEEDED TO BE COLLECTED. CONTRIBUTED TO FINAL DRAFT. SUBMITTED PAPER & CONTRIBUTED TO RESPONSES TO REFEREES & EDITOR

Certification by Co-Authors

The undersigned hereby certify that:

- the above statement correctly reflects the nature and extent of the PhD candidate's contribution to this work, and the nature of the contribution of each of the co-authors; and

Name	Signature	Date
Yang Hu	[signature]	28/08/18
Les Oxley	[signature]	22/8/18

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Chapter 5

Hu, Y., & Oxley, L. (2018). Bubbles in US regional house prices: evidence from house price–income ratios at the State level. Applied Economics, 50(29), 3196-3229.

Nature of contribution by PhD candidate	Performing empirical analysis, data collection, writing the initial draft.
Extent of contribution by PhD candidate (%)	80%

CO-AUTHORS

Name	Nature of Contribution
Les Oxley	DISCUSSIONS ON ECONOMETRIC METHODS TO APPLY
	& INTERPRETATION OF RESULTS, SUGGESTIONS ON
	DATA TO COLLECT & APPLY METHODS TO, FINAL
	DRAFT, SUBMISSION + CONTRIBUTIONS TO
	RESPONSES TO REFEREES & EDITOR

Certification by Co-Authors

The undersigned hereby certify that:

- the above statement correctly reflects the nature and extent of the PhD candidate's contribution to this work, and the nature of the contribution of each of the co-authors; and

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Chapter 6

Hu, Y., & Oxley, L. (2018). Do 18th century 'bubbles' survive the scrutiny of 21st century time series econometrics?. Economics Letters, 162, 131-134.

Nature of contribution by PhD candidate	Performing empirical analysis, data collection, writing the initial draft.
Extent of contribution by PhD candidate (%)	80%

CO-AUTHORS

Name	Nature of Contribution
Les Oxley	IDEA BEHIND APPLICATION TO 18TH BUBBLES. SUGGESTION & IDENTIFICATION OF DATA TO COLLECT. DISCUSSIONS ON ECONOMETRIC METHODS + TESTS. COMPLETION OF FINAL DRAFT. SUBMISSION. CONTRIBUTED TO RESPONSES TO REFEREE ST EDITOR.

Certification by Co-Authors

The undersigned hereby certify that:

- the above statement correctly reflects the nature and extent of the PhD candidate's contribution to this work, and the nature of the contribution of each of the co-authors; and

Name	Signature	Date
Yang Hu	[signature]	28/08/18
Les Oxley	[signature]	22/8/18

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Chapter 7

Hu, Y., & Oxley, L. (2017). Bubble contagion: evidence from Japan's asset price bubble of the 1980s-90s. (~~Under review~~ Revise + resubmit LTO at Journal of the Japanese and International Economies)

Nature of contribution by PhD candidate	Performing empirical analysis and writing the initial draft.
Extent of contribution by PhD candidate (%)	80%

CO-AUTHORS

Name	Nature of Contribution
Les Oxley	DISCUSSIONS OF ECONOMETRIC TESTS & INITIAL
	RESPONSES TO RESULTS ~~RECEIVE~~ PRODUCED.
	SUGGESTIONS ON ADDITIONAL TESTING.
	COMPLETION OF FINAL DRAFT, SUBMISSION
	CONTRIBUTIONS TO REFEREE & EDITOR COMMENTS

Certification by Co-Authors

The undersigned hereby certify that:

- the above statement correctly reflects the nature and extent of the PhD candidate's contribution to this work, and the nature of the contribution of each of the co-authors; and

Name	Signature	Date
Yang Hu	[signature]	28/08/18
Les Oxley	[signature]	22/8/18

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Chapter 8

Hu, Y., & Oxley, L. (2017). Do US house prices reflect economic fundamentals? Evidence from quantile cointegration tests

Nature of contribution by PhD candidate	Performing empirical analysis, data collection, writing the initial draft.
Extent of contribution by PhD candidate (%)	80%

CO-AUTHORS

Name	Nature of Contribution
Les Oxley	DISCUSSION ON ECONOMETRICS & SUGGESTIONS ON
	LITERATURE TO CONSULT. SUGGESTIONS ON DATA TO
	APPLY METHODS TO. COMPLETION OF FINAL DRAFT,
	~~SUBMISSION & RESPONSES TO REFEREES & EDITOR~~
	SUPPORT FOR YANG HU TO VISIT YONGCHEOL
	SHIN AT YORK UNIVERSITY, UK

Certification by Co-Authors

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