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Abstract: We present a new approach to evaluate the full extent of price stickiness in credit card interest rates by modifying the existing asymmetric models so that they can be adopted for testing both the amount and adjustment asymmetries as well as the lagged dynamic inertia. Consistent with similar studies, banks behave asymmetrically in response to changes in the Reserve Bank of Australia's (RBA) target interest rate. Rate rises are passed onto the consumer faster than rate cuts and the credit card interest rate showed a very significant degree of downward rigidity. Based on the magnitude of the pass-through parameters obtained from short-run dynamic models, rate rises had a full one-to-one and instantaneous impact on credit card interest rates. However, in absolute terms the short-run effects of rate cuts were not only less than half of the rate rises but also were delayed on average by three months.

JEL classification: E43; E58; G21

Keywords: Interest rates; Asymmetric behaviour; Credit cards; Australia

1. Introduction

Analysing credit card interest rates has been one of the growing strands of research in the banking and finance literature. Credit card suppliers generally operate within a competitive market with many financial institutions offering their credit cards with limited barriers to entry. However, in such a market it is puzzling as to why credit card interest rates (compared to other lending rates) are so high and so sticky downwards, while banks enjoy high profits at the same time. There are a large number of recent studies that have analysed the interest rate pass-through of the central bank's policy rate to various types of deposit and lending rates (see *inter alia* Hofmann and Mizen, 2004; Chong, Liu and Shrestha, 2006; Payne, 2006, 2007; De Graeve, De Jonghe and Vennet, 2007; Liu, Margaritis and Tourani-Rad, 2008; Payne and Waters, 2008; Chong, 2010; De Haan and Sterken, 2011). However, banks' asymmetric behaviour in setting the credit card interest rate has received considerably less attention in the literature, especially in the context of the Australian banking industry.

Previous studies in other countries have found that credit card interest rates are relatively higher than other lending rates and show a very significant degree of inertia and downward rigidity. This has resulted in the rising spread between credit card rates and other lending rates. All major explanations for the asymmetric behaviour of credit card interest rates in the literature can be grouped into six categories: (1) search costs, switching costs, and adverse selection (Ausubel, 1991; Lowe and Rohling, 1992; Lowe, 1995; Calem and Mester, 1995; Stango, 2002; Berlin and Mester, 2004; Calem, Gordy and Mester, 2006; Agarwal, Chomsisengphet and Liu, 2010); (2) rational or irrational borrowers (Brito and Hartley, 1995; Della Vigna and Malmendier, 2004); (3) fixed and variable interest rates (Stango, 2000); (4) tacit collusion between banks (Knittel and Stango, 2003); (5) option value of credit lines (Park, 2004); (6) credit card penalty fees and risk (Massoud, Saunders and Scholnick, 2007).

Providing a comprehensive review of the existing literature is beyond the scope of this paper. For a concise account of these studies see Scholnick et al. (2008).

Ausubel (1991), Lowe and Rohling (1992), Lowe (1995) and Calem and Mester (1995) can be considered as four pioneering studies posing a combination of three explanations for this puzzle, namely: search costs, switching costs, and adverse selection. In their views, to change credit card providers, consumers may incur two possible types of costs: search costs (associated with finding the information on alternative providers) and switching costs, (the expenses incurring in the process of changing from one provider to another). High risk borrowers are usually more inclined to search for lower interest rate cards because they think it is more likely that they may use that debt in the future (Ausubel, 1991). Due to this likely phenomenon, financial institutions hesitate to reduce their credit card interest rates because this action may in fact encourage high risk borrowers to apply for their products. According to Calem, Gordy and Mester (2005), borrowers who possess higher credit card balances also have higher search costs. Thus, if a bank unilaterally reduces its credit card interest rate, it will be appealing to mainly less profitable borrowers who have lower balances. Both Ausubel (1991) and Calem and Mester (1995) therefore conclude that banks have no incentive to unilaterally lower their interest rate.

In a competitive equilibrium and based on completely rational individuals, Brito and Hartley (1995) suggested an alternative explanation to that of Ausubel (1991), Lowe and Rohling (1992) and Calem and Mester (1995). In their view, for two reasons borrowers continue to use credit card debt even if its interest rate is substantially higher than other alternative forms of borrowing: (a) transactions costs associated with the use of non-credit card sources for short periods of time are usually high; (b) borrowers try to smooth their income and consumption streams through time. Brito and Hartley (1995) also provide an explanation for the rising spread between credit card rates and other lending rates. Unlike

other bank loans, credit cards enable the flexibility to borrow more even if consumption exceeds income, especially when other loans are more expensive to set up or the loan periods are either relatively short or unpredictable. However, if the spread between credit card and other forms of finance becomes significantly wider, the credit card benefits will diminish and eventually low-risk borrowers, who have access to other loans, will stop using credit card debt.

Credit card providers usually offer two types of interest rate: fixed rates which do not change for long periods and variable rates which move concurrently with market interest rates. Stango (2000) believes that such a pricing structure influences the competitive structure of the market. Adopting a game theory approach, Stango (2000) demonstrates that the size of firms can also exert a strong influence on the pricing structure chosen (i.e. fixed or variable rates). Knittel and Stango (2003) offer another alternative explanation for credit card stickiness based on the possibility of tacit collusion between banks. Numerous States in the US introduced the State level price ceilings based on Usury Laws in the 1980s while in other States there were no binding ceilings. Knittel and Stango (2003, p.1471) argue that “credit card providers resort to tacit collusion by using the interest rate level of the binding ceiling in some states as the focal point for interest rates in those states that did not have binding ceilings”. The main conclusion of a comprehensive review of the literature by Scholnick et al. (2008, p.1469) is “that while a large amount of research has already been undertaken on credit cards, debit cards and ATMs, we believe that there are still a great many issues and puzzles that remain to be resolved”.

In the contemporary literature the asymmetric behaviour of credit card interest rates could also be referred to as “rockets-and-feathers hypothesis”. In a similar context, Bacon (1991) argued that gasoline prices “shoot up like rockets” in response to a positive rise in oil prices and “float down like feathers” in response to a fall. Hannan and Berger (1991) and

Neumark and Sharpe (1992) are among earlier studies that have tested the rockets and feathers hypothesis in the context of the banking industry. Their thorough investigation suggests that, as compared to negative shocks, consumer deposit interest rates respond much slower to positive shocks. Mojon (2000) examined the pass-through parameters in a multi-country setting (i.e. Belgium, France, Germany, Italy, Netherlands and Spain) by reporting only short-run multipliers for a panel of 25 credit market rates and 17 deposit rates and for the period 1979–1998. He found that the volatility of the money market rate lowers the pass-through parameters for both credit and deposit rates, whereas increasing banking competition can increase it.

This paper aims to develop a general modelling framework that allows one to examine the full asymmetric effects of changes in the cost of funding on the variable-interest rate charged by Australian banks and non-bank financial institutions for credit cards. We consider two specific issues. First, in overall terms, does the credit card interest rate respond asymmetrically to changes in the funding cost? If the funding cost changes, will Bacon's (1991) "rockets-and-feathers hypothesis" be applicable in the context of Australia's retail credit card rates? Second, when the cash rate, which is known as the federal funds rate in the US, increases by one per cent, on average, by how much and how quickly do the standard-variable rate for credit cards rise? One should note that significant asymmetric rate adjustments can also adversely affect the efficacy of the RBA's (Reserve Bank of Australia) expansionary monetary policy.¹

The cash rate is not the only factor that affects the lenders' behaviour. The decision of individual lenders to change their interest rates is also influenced to varying degrees by a number of other factors such as the extent of securitisation and the individual bank's exposures to different types of external and internal sources of borrowing. However, the cash

¹ The RBA is Australia's central bank.

rate has become increasingly politicised and the focus of much attention by media commentators and the public alike. This is not hard to understand as for some families interest payments constitute a substantial part of their income and interest rate changes have a direct and appreciable effect on their consumer decisions. During periods of both increasing and decreasing rates, the media not only focuses on which lenders are raising or lowering their rates more quickly in response to the direction provided by the cash rate, but also the extent to which the change in the cash rate is passed on to borrowers. In the context of mortgage rates banks have responded in a variety of ways, with some justifying apparently excessive rate rises outside the RBA cycle as being due to the increased cost of funding, while others adopt a strategy of rate rise restraint to capture market share. Particular groups of lenders (especially, credit unions and building societies, mortgage originators, and so on) have also used their varying responses to the cash rate for publicity purposes aimed at leveraging their 'consumer-friendly' credentials, especially against the major banks.

Most of the previous studies have focused on the asymmetric pass-through of funding costs into mortgage-interest rates, an issue which is by no means Australia specific. Other developed countries are also facing the same dilemma. For instance, Corvoisier and Gropp (2002) and Bikker and Haaf (2002) have highlighted substantial differences across European countries in terms of the pass-through of monetary policy interest rate changes into money market rates. Bikker and Haaf (2002) examined the banking sector competition in 23 countries and discovered that competition is much weaker in local markets.

Payne and Waters (2008) thoroughly analysed the long-run interest rate pass through of the federal funds rate to the prime rate over the period February 1987 to October 2005. They state that the response of the prime rate to changes in the federal funds rate appears asymmetric. Payne (2007, 2006) has also proposed several useful momentum threshold autoregressive models in the literature on interest rate pass-through from the federal funds

rate to mortgage rates. He has found that the mortgage rates are co-integrated with the federal funds rate in the long run but with incomplete interest rate pass through in the short run. A recent study found that the pass-through of official interest rate changes into mortgage rates in the Netherlands is only about half of the Euro-zone average (De Haan and Sterken, 2011).

Little empirical work has been conducted regarding the dynamic effects of positive and negative changes in the central bank's policy rate on credit card rates. However, there are numerous studies in the literature that found evidence of asymmetric pricing for other lending rates (e.g., Haney, 1988; Allen, Rutherford and Wiley, 1999; Hofmann and Mizen, 2004; and De Haan and Sterken, 2011). Liu, Margaritis and Tourani-Rad (2008) conducted a comprehensive study by using New Zealand's monthly data on various mortgage, deposit and lending rates for the period 1994–2004. They quantified the extent of pass-through and adjustment speed of retail interest rates in response to changes in benchmark market rates by using the Phillips and Loretan (1991) methodology. They found that not only the long-term pass-through of retail rates do vary significantly across different rates but similar to Kim and Nguyen (2008) they also argue that changes in the policy rate can exert more influence on short-term interest rates. With some evidence of asymmetry in the adjustment of retail rates, they conclude that since 1999 the Reserve Bank of New Zealand's transparency in setting its target policy rate has reduced the instrument volatility and enhanced the efficacy of monetary policy.

Lowe and Rohling (1992) and Lowe (1995) investigated the degree of stickiness in Australia's various deposit and lending rates including credit cards by comparing them in the pre- (1979–1985) and post- (1986–1991) deregulation periods.² They found that the credit card interest rate was the stickiest rate and there was no sign of observing less rigidity in the

² Prior to 1985 the maximum rate that could be charged on credit cards had been set at 18 per cent per annum by the Reserve Bank of Australia. In April 1985, this rate was deregulated.

post-deregulation era. Lowe and Rohling (1992) stated that switching and search costs can be the most critical reasons behind the interest rate stickiness in the Australian banking system.³

Lim (2001) examined the asymmetric adjustments between three Australian bank interest rates: a bank bill rate, a loan rate and a deposit rate. She uses a multivariate asymmetric error-correction model to capture the long- and short-run relationships between the levels of the rates and short-run relationships between the changes in the rates. Her empirical results indicate that “banks value their borrowing customers and tend to pass on decreases in the loan rates faster than they pass on increases” (Lim, 2001, p.146).⁴ Lim’s results are in sharp contrast with previous Australian studies (Lowe and Rohling, 1992; Lowe, 1995; Kim and Nguyen, 2008) as well as international studies (Hannan and Berger, 1991; Neumark and Sharpe, 1992; Payne, 2006, 2007; Allen and McVanel, 2009; De Haan and Sterken, 2011).

Kim and Nguyen (2008) examine the effects of the RBA and the U.S. Fed’s target interest rate announcement news on the Australian financial markets over the period 1998–2006, including the 90–day bank bills, 3–year and 10–year bond rates. They found evidence of asymmetric news effects on interest rates where they respond more strongly to unexpected rate rises than rate falls. However, “the news effect is stronger at the short-term ends of the interest rates” (Kim and Nguyen, 2008, p.392).

In this paper we address an important policy issue not recently tackled in the Australian context. The majority of previous studies have tested for asymmetric effects of changes in the central bank’s policy rate by using a Wald test in which two coefficients corresponding to positive and negative changes in the lagged error correction terms were assumed to be equal.

³ For a comprehensive and critical review of the RBA’s credit card reforms, see Gans and King (2003).

⁴ Using monthly data, Lim (2001) has tested only for the first order autocorrelation and the first order ARCH term (see Table 5, p.145).

In other words, as it is demonstrated in the next section, they have tested for the adjustment asymmetry (e.g. see Sarno and Thornton, 2003; Chong, Liu and Shrestha, 2006; Liu, Margaritis and Tourani-Rad, 2008; Chong, 2010).

If the policy rate was not co-integrated with a particular lending rate, then the proposed asymmetric models in such studies would have become symmetric. However, in this paper we argue that even if the lending rate is not co-integrated with the policy rate, it is quite possible that short-run positive and negative changes in the central bank's policy rate may still exert different effects on the corresponding lending rate. In other words, the amount asymmetry should also be tested but within a dynamic framework. It should be noted that Lowe and Rohling (1992) proposed a simple asymmetric model which was used for testing the amount asymmetry. However, in their model only positive and negative changes in the cash rate at time t were allowed to impact on the lending rates (see Table 4 in Lowe and Rohling, 1992, p.25), whereas in our proposed models the amount asymmetry is tested allowing up to 12-month lag.

Our analysis is undertaken within a sample period during which the RBA has conducted monetary policy by setting the desired interest rate on overnight loans in the money market. This paper offers an analytical modelling framework to quantify the full extent of asymmetric rate changes (if any), resulting in greater efficiency and transparency of the lending market. The results and policy implications of this paper increases our understanding of the credit card lending market in Australia and are beneficial to all borrowers and various government regulators, which can play an important role in market efficiency and consumers' protection.

The rest of this paper is structured as follows. In Section 2 various theoretical models are postulated which capture two possible forms of asymmetric behaviour (i.e. the amount and adjustment asymmetry) as well as the dynamic asymmetric responses of the lagged dependent variable. Section 3 discusses the choice of our sample period, the descriptive statistics of the

data employed followed by the results. Section 4 presents the empirical results of the long- and short-run credit card interest rate models as well as the policy implications of the study. Section 5 provides some concluding remarks and the final section presents the agenda for future research.

2. Theoretical framework

In order to capture the long-run relationship between the central bank's policy rate and the credit card cash rate, following Rouseas (1985), we utilise equation (1) in which financial institutions set their standard variable rate as a mark-up on the cash rate. That is:

$$i_t = \theta_0 + \theta_1 r_t + \varepsilon_t \quad (1)$$

where:

i_t = the standard variable interest rate at period t (as a percentage) for credit cards.

r_t = the cash rate (as a percentage) prevailing at time t ,

θ_0 and θ_1 are the average of banks' mark-up and pass-through parameters, respectively, and

ε_t = the white noise error term.

Several other studies in the literature have also specified similar long-term relationships between the policy rate and various lending and deposit rates (see *inter alia* Heffernan, 1997; Chong, Liu and Shrestha, 2006; Toolsema and Jacobs, 2007; Liu, Margaritis and Tourani-Rad, 2008; De Haan and Sterken, 2011). Standard unit root tests, such as the Augmented Dickey-Fuller (ADF) test, the DF test with GLS detrending (Elliott, Rothenberg, and Stock, 1996) and the Kwiatkowski-Phillips-Schmidt-Shin (KPSS, 1992) test as well as the Lee and Strazicich (2003) test, which endogenously incorporate two structural breaks in the testing procedure, are used in the next section to ensure that our empirical results are not biased towards the erroneous non-rejection of the non-stationarity hypothesis. Since both variables in equation (1) are I(1), in the next step the aim is to test if the retail credit card rate is co-

integrated with the cash rate. Following the Engle-Granger two-step procedure if the two variables are co-integrated, then the stationary residuals resulting from equation (1) could form an error correction mechanism (*EC*), representing the short-run deviation from the long-run equilibrium.

Standard co-integration tests (Johansen, 1995) implicitly assume a symmetric adjustment process but if the adjustment process is asymmetric or if the loan rates are sticky downwards, these tests can lead to misleading results. In other words, the Engle-Granger type tests with a linear adjustment procedure will be inappropriate when the dynamic adjustment of loan rates in fact exhibits non-linear behaviour. One alternative is to use the threshold co-integration test proposed by Granger and Lee (1989), Enders and Granger (1998) and Enders and Siklos (2001).

With some modifications the majority of previous studies (see inter alia Sarno and Thornton, 2003; Chong, Liu and Shrestha, 2006; Liu, Margaritis and Tourani-Rad, 2008; Chong, 2010) have used the following short-run dynamic model based on the assumption that the corresponding two interest rates are co-integrated:

$$\Delta i_t = \xi_0 + \sum_{j=0}^q \lambda_j \Delta r_{t-j} + \sum_{j=1}^q \eta_j \Delta i_{t-j} + \omega^+ EC_{t-1}^+ + \omega^- EC_{t-1}^- + v_t \quad (2)$$

$EC_t = \hat{\varepsilon}_t$ is the estimated residual obtained from equation (1), λ_j denotes the symmetric short-run effects of changes in r on i at time $t-j$, and η_j is the symmetric effect of changes in the j^{th} -lagged dependent variable, where $j=1,2,\dots,12$). The superscripts $+$ and $-$ denote the positive part and negative part of EC_t , respectively as defined below:

$$\Delta EC_t^+ = \max\{\Delta EC_t, 0\} \Rightarrow \Delta EC_t^+ = \Delta EC_t \text{ if } \Delta EC_t > 0 \text{ and } \Delta EC_t^+ = 0 \text{ if } \Delta EC_t \leq 0 \quad (3)$$

$$\Delta EC_t^- = \min\{\Delta EC_t, 0\} \Rightarrow \Delta EC_t^- = \Delta EC_t \text{ if } \Delta EC_t \leq 0 \text{ and } \Delta EC_t^- = 0 \text{ if } \Delta EC_t > 0 \quad (4)$$

In equation (2) ω^+ and ω^- are the error-correction parameters associated with positive and negative values of the EC_{t-1} part, which are expected to be negative. If the null hypothesis $\omega^+ = \omega^-$ vs. $\omega^+ \neq \omega^-$ is rejected, one can argue that the adjustment process towards the long-run equilibrium is asymmetric and hence there is evidence of adjustment asymmetry. If $|\omega^-| > |\omega^+|$, then a lagged negative disequilibrium between the actual interest rate and its equilibrium path results in a relatively swifter error correction as compared to the case of a lagged positive disequilibrium. This can easily be justified as financial institutions prefer to increase the actual loan rate to its desired path immediately when $EC_{t-1} < 0$. On the other hand, when the lagged-actual retail rate is above the equilibrium path, banks may have less desire to lower their rate to its equilibrium path, causing the speed of adjustment to be more sluggish or sometimes non-existent. Thus, in the case of adjustment asymmetry $|\omega^-| > |\omega^+|$.

In the absence of the adjustment asymmetry (i.e., $\omega^+ = \omega^-$), equation (2) can be rewritten as:

$$\Delta i_t = \xi_0 + \sum_{j=0}^q \lambda_j \Delta r_{t-j} + \sum_{j=1}^q \eta_j \Delta i_{t-j} + \omega EC_{t-1} + v_t \quad (5)$$

If i_t and r_t are not co-integrated, then in equation (2) the error correction term will disappear and one may use equation (6) instead:

$$\Delta i_t = \xi_0 + \sum_{j=0}^q \lambda_j \Delta r_{t-j} + \sum_{j=1}^q \eta_j \Delta i_{t-j} + v_t \quad (6)$$

Therefore, if two interest rates are not co-integrated or if there is no adjustment asymmetry, one may end up using a symmetric model such as (6) and (5), respectively. However, it is quite possible that short-run positive and negative changes in r (with different lags) may still have asymmetric impacts on the dependent variable. Similarly positive and negative changes in Δi_{t-j} can also have different effects on Δi_t . In this paper we propose equation (7) as an

alternative and flexible framework which can be used not only for testing both the adjustment ($\omega^+ \neq \omega^-$) and amount ($\lambda_i^+ \neq \lambda_i^-$) asymmetries in the short run but also for the “path-dependence” (stickiness) asymmetry ($\eta_i^+ \neq \eta_i^-$):

$$\Delta i_t = \xi_0 + \sum_{j=0}^q \lambda_j^+ \Delta r_{t-j}^+ + \sum_{j=0}^q \lambda_j^- \Delta r_{t-j}^- + \sum_{j=1}^q \eta_j^+ \Delta i_{t-j}^+ + \sum_{j=1}^q \eta_j^- \Delta i_{t-j}^- + \omega^+ EC_{t-1}^+ + \omega^- EC_{t-1}^- + v_t \quad (7)$$

where λ_j^+ and λ_j^- are the short-run effects of positive and negative changes in the cash rate on retail interest rate series at time $t-j$, respectively, and η_j^+ and η_j^- are the effects of positive and negative changes of the dependent variable with $t-j$ lag, respectively. Δi_t^+ , Δi_t^- , Δr_t^+ and Δr_t^- are also defined similarly to EC_{t-1}^+ and EC_{t-1}^- .

It should be noted that Allen and McVanel (2009) and De Haan and Sterken (2011) have recently used a model which allows us to test both the amount and adjustment asymmetries in a short-run dynamic model for mortgage rates. However, in their models it is assumed that $\eta_i^+ = \eta_i^-$. To the best of our knowledge, none of the previous studies have utilised a general model as flexible as equation (7).

In equation (7) the extent of the amount asymmetry depends on the difference between $\sum_{j=0}^q \lambda_j^+$ and $\sum_{j=0}^q \lambda_j^-$. In other words, the higher the difference, the more the extent of the amount asymmetry. It is also hypothesised that rate rises have some immediate effects on the retail rates (say for instance only λ_0^+ and λ_1^+ within the first month to be statistically significant), whereas in the case of rate cuts the lagged effects are exhausted within a longer period such as the first 12 months (i.e., $\lambda_0^-, \lambda_1^-, \dots, \lambda_{12}^-$). Therefore, in our model the short-run effects of changes in the cash rate on the credit card rate are allowed to be different in magnitude as well as through time.

Using equation (7), a Wald test can also be employed to test the amount asymmetry. If the null hypothesis $\lambda_j^+ = \lambda_j^-$ vs. $\lambda_j^+ > \lambda_j^-$ (for all j values) is rejected, there will be evidence of amount asymmetry. This means that short-run changes in the cash rate can exert asymmetric effects on the credit card rate. Accordingly, if $\sum_{j=0}^{12} \lambda_j^+ > \sum_{j=0}^{12} \lambda_j^-$, then banks pass through short-run interest rate rises more than rate decreases. For a detailed discussion of the distinction between amount and adjustment asymmetries in the literature see Chen, Finney and Lai (2005), Bachmeier and Griffin (2003), Bettendorf, van der Geest and Kuper (2009), Allen and McVanel (2009) and De Haan and Sterken (2011).

In a similar way, if the null hypothesis $\eta_j^+ = \eta_j^-$ vs. $\eta_j^+ > \eta_j^-$ (for all j values) is rejected, one may argue that positive and negative lagged values of the dependent variable have different impacts in equation (7). For example, if $\eta_j^+ = 0$ and $\eta_j^- \neq 0$, *ceteris paribus* one may then conclude that the retail interest rates are sticky downwards.

3. Data

It should be noted that Australia's approach to monetary policy has undergone significant changes over time. From the mid-1970s until 1985, based on the assumption of a strong and persistent relationship between inflation and the supply of money, monetary policy was conducted by targeting the annual growth of M3. However, in 1985 this policy was abandoned because deregulation of the financial system made M3 a misleading guide to the stance of monetary policy (Grenville, 1990). From 1985 to 1988 a "checklist approach" was adopted, whereby a multitude of indicators such as, monetary aggregates, the GDP growth rate, the shape of the yield curve, exchange rates, and the unemployment rate were considered prior to the implementation of monetary policy. The checklist approach was also

unsuccessful and finally discontinued in 1989 due to the impossibility of monitoring the above indicators which could provide contradictory policy signals.

Since 1990, the RBA has conducted monetary policy by setting the desired (target) interest rate on overnight loans in the money market. This year has been considered as the starting point for our sample period. Through the monetary policy transmission mechanism, changes in the cash rate are ultimately reflected in the rates on all lending instruments in line with the desired policy intent. The cash rate in Australia is now considered the baseline for the various interest rates paid by borrowers. Fig. 1 suggests that during our sample period, which consists of monthly data from January 1990 to October 2011, there is a very close relationship between the cash rate and credit card rate but at the same time the spread between the credit card rate and the cash rate has been gradually widening since 1990.⁵ All our monthly data have been obtained from the RBA (2011, Statistical Tables F1 and F5).

[FIG. 1 ABOUT HERE]

Brännäs and Ohlsson (1999) suggest that the detection of asymmetry depends on the sampling frequency of the series. They argue that the use of aggregated frequencies (i.e., annual) may obscure the nonlinearities or asymmetries that exist in a series. They found that “asymmetric monthly series may become symmetric when aggregated to quarterly or annual frequencies” (Brännäs and Ohlsson, 1999, p.341). Therefore, in this paper, we use monthly data to detect any discernable asymmetric behaviour not easily observable in the raw and aggregated data. As can be seen in our results presented in the next section, the effects of positive and negative changes in the cash rate on the credit card rate appear to be exhausted within the first three months. Therefore, this confirms Brännäs and Ohlsson’s point that the

⁵ The difference between the credit card rate and the cash rate increased from 4.89 per cent in January 1990 to 14.95 per cent in October 2011.

use of annual time series in lieu of monthly could have masked the existing nonlinearities or asymmetries in the series.

Table 1 presents the descriptive statistics and the unit root test results for all of the variables employed in this paper. During the sample period the average credit card rate was 17.7 per cent, ranging from a minimum of 14.4 in May 1994 to a maximum of 23.5 in January 1991. With a standard deviation of 2.44 per cent, the credit card rate showed a slightly less variability than the cash rate (3.20 per cent), particularly after the adoption of the RBA's inflation targeting policy in 1993. Based on the reported Jarque-Bera statistics, the null hypothesis of normality is rejected at any conventional level for all series as their distributions are positively skewed and show a typical leptokurtic pattern with the kurtosis statistic well exceeding 3.0.

An important step before undertaking our empirical investigation is to determine the time series properties of the data. This is an important issue since the use of non-stationary data in the absence of co-integration can result in spurious regression results. To this end, three unit root tests, i.e. the ADF test, the DF test with GLS Detrending (Elliott, Rothenberg, and Stock, 1996) and the Kwiatkowski-Phillips-Schmidt-Shin (KPSS) (Kwiatkowski et al. 1992) test, have been adopted to examine the stationarity, or otherwise, of the time series data. In this paper the lowest value of the Schwarz criterion (SC) has been used as a guide to determine the optimal lag length. Unlike both the ADF test and the DF test with GLS, the KPSS test has the null of stationarity, and the alternative indicates the existence of a unit root. In Table 1 we have also reported the results of the Lee and Strazicich (2003) test, which endogenously incorporates two structural breaks in the testing procedure. The resulting two break dates for individual series are also reported in Table 1. Irrespective of which test is considered, it appears that our two interest rate series are $I(1)$.

[TABLE 1 AND FIG. 2 ABOUT HERE]

4. Empirical results

In this section, we first use monthly data for the period January 1990–October 2011 to estimate the long-run relationship between the cash rate (r_t) and the credit card rate (i_t) by the OLS method. The estimation results are presented in Table 2. The average mark-up coefficient for i_t is 13.6 per cent. Since credit cards are usually riskier than other types of personal loans, the estimated high mark-up is quite justifiable. The long-run pass-through parameter is estimated at 0.67 per cent, which appears to be of consistent sign and order of magnitude and highly significant. This means that on average the long run credit card rate is 7.1 per cent (13.6–6.51) more than the cash rate. We also tested the null hypothesis that the long-run pass-through parameter is equal to 1, i.e. $H_0 : \theta_1 = 1$ vs $H_1 : \theta_1 < 1$. This hypothesis was rejected at the 1 per cent level of significance (see Table 2) and as a result the spread between the credit card rate and the cash rate has been widening (see Fig. 1).

We have also provided the recursive estimates of both the mark-up and the pass-through parameter in Fig. 2. It seems that both the mark-up and pass-through parameters enjoy relative stability around the estimated mean value since 1993, when the RBA adopted the inflation targeting policy. According to the results of four unit-root tests presented in Table 2, it seems that the resulting residuals obtained from the OLS estimation of equation (1) are I(1) despite considering two endogenous breaks in the Lee and Strazicich (2003) test. Therefore, one cannot argue that the credit card rate is co-integrated with the cash rate. The widening spread between the two series presented in Fig. 1 could also partly explain the absence of co-integration. Similar results were obtained by Lowe (1995, Table A1) 17 years ago in his comprehensive study of the linkage between the cash rate and Australia's lending rates including credit card rate.⁶

⁶ This could be the reason why both positive and negative changes in the lagged residuals (as the error correction terms) are not statistically significant in the estimated Model I (see Table 3 in the next section).

As discussed in the previous section, we first estimate equation (7) and present the results in Table 3. In terms of determining the optimal lag length (q) in equation (7), given that we are using monthly data, an upper band of 12 lags was allowed. Following the general-to-specific methodology, insignificant variables in equation (7) were omitted on the basis of a battery of maximum likelihood tests. We imposed joint-zero restrictions on explanatory variables in the unrestricted (general) model to obtain the most parsimonious and robust equation. We found that the cash rate is weakly exogenous with respect to the dependent variable. This justifies the use of the OLS method to estimate our short-run dynamics models based on an asymmetric version of the Engle-Granger two-step procedure. This is not counter-intuitive as the cash rate is directly controlled by the RBA as a policy variable. The weak exogeneity test results are available from the authors upon request.

4.1. Short-run dynamic models for Δi_t

Let us focus on the short-run dynamic asymmetric model for Δi_t . As mentioned earlier the error correction series is I(1) (see Table 2) and at the same time it is not statistically significant in Model I (see Table 3). We have thus excluded the EC_{t-1} term to obtain Model II. It should be noted that the other estimated coefficients in both Models I and II remain very similar in terms of their magnitudes and statistical significance. Since i_t and r_t are not co-integrated, we only focus on the interpretation of Model II below.

[TABLE 2 AND FIG. 3 ABOUT HERE]

According to the results of Model II presented in Table 3, all of the estimated coefficients are statistically significant at the five per cent level or better and have the expected theoretical signs. Model II also performs well in terms of goodness-of-fit statistics. The adjusted R^2 is 0.42 and the overall F test rejects the null hypothesis at the one per cent level of significance. Furthermore, it passes a battery of diagnostic tests and shows no sign of

serial correlation (see the Breusch-Godfrey LM tests), misspecification (see the Ramsey RESET test), heteroskedasticity (see the ARCH, White and Breusch-Pagan-Godfrey tests) and instability (see the Chow forecast tests in three different forecasting periods). The only diagnostic test that Model II could not pass was the Jarque-Bera normality test of the residuals. This was unavoidable given the use of monthly data.

One problem associated with the analysis of the relationship between the cash rate and the credit card rate is non-constancy or instability of the estimated coefficients which can create economic and econometric complications in deriving any inference from the empirical model. Therefore, parameter constancy is pivotal in modelling changes in the credit card rate.

We have evaluated our preferred estimated short-run model (i.e. Model II) by a number of

recursive stability tests which are displayed in Fig. 3 in the following order:
$$\begin{bmatrix} a & b & c & d \\ e & f & g & h \\ i & j & k & l \end{bmatrix}$$

where panel (a) displays the recursive residuals; (b) depicts the CUSUM test; (c) shows the one-step forecast test; (d) is the n-step forecast test; and panels (e) to (l) present the recursively estimated eight coefficients in exactly the same order that these coefficients appear in Model II in Table 3 (from top to bottom). These evaluative tests are useful in assessing stability of a model, as recursive algorithms avoid arbitrary splitting of the sample. Overall, the graphical tests for stability reported in Fig. 3 support the in-sample constancy of all of the estimated coefficients, with the only two exceptions being the coefficients of Δr_t^- and Δr_{t-3}^- which were subject to minor changes around September 2009.

Ceteris paribus, if in the short run the cash rate had increased, say by one per cent in a particular month, this would have immediately led to a rise of 1.032 per cent in the credit card rate. On the other hand, a similar one per cent rate cut would have resulted in only 0.163 per cent fall at time t and a further 0.295 per cent at time $t-3$. The total short-run effect associated with the RBA's rate cut would then be only 0.458 per cent within the first three

months (0.163+0.295), whereas the corresponding effect for a rate rise would be an immediate 1.032 per cent increase. Thus, in absolute terms the sum of short-run effects of rate cuts were less than half of the rate rises.

What about the adjustment asymmetry? Since both ω^- and ω^+ are statistically insignificant (as a result of the lack of co-integration between i_t and r_t), it is clear that the adjustment asymmetry is not applicable. However, it appears that the credit card rate is very

sticky downwards because $\sum_{j=1}^q \hat{\eta}_j^- > \sum_{j=1}^q \hat{\eta}_j^+$. Due to the persistence of interest rates in the post-

1990 era, the estimated coefficients of the lagged dependent variable are also statistically significant, especially those of negative changes at time $t-1$, $t-10$ and $t-12$, further supporting the downward stickiness hypothesis. This proffers support for the short-run applicability of the rockets-and-feathers hypothesis in the context of the credit card market in Australia. The results of Model II in Table 3 show that the effects of Δi_{t-j}^+ on the dependent variable are statistically significant at only one lag $t-1$ that is $\hat{\eta}_1^+ = 0.213$. However, the effects of Δi_{t-j}^- on

the dependent variable were statistically significant at $t-1$, $t-10$ and $t-12$, yielding to the sum of the lagged dependent variable coefficients $\sum_{j=1}^q \eta_j^- = 0.292 + 0.196 + 0.129 = 0.617$. Therefore,

one may conclude that Australian credit card rates face downward rigidity when the RBA reduces its cash rate. We can use equations (8) and (9) to approximate the total effects of a

one per cent rise or fall in the cash rate on the credit card rate, respectively:

$$\frac{\Delta i_t}{\Delta r_t^+} = \frac{\lambda_0^+}{1 - \sum_{j=1}^q \eta_j^+} \Rightarrow \frac{1.032}{1 - 0.295} = 1.46 \quad (8)$$

$$\frac{\Delta i_t}{\Delta r_t^-} = \frac{\lambda_0^-}{1 - \sum_{j=1}^q \eta_j^-} \Rightarrow \frac{0.163}{1 - (0.292 + 0.196 + 0.129)} = 0.43 \quad (9)$$

It should be noted that a one per cent rate cut (compared to a rate rise), not only has a smaller total effect (0.43 vs. 1.46) but also it takes longer to eventuate. We have also formally tested the absence of the amount asymmetry by using a Wald test. Using the estimated Model II presented in Table 3, the null hypothesis $\lambda_j^+ = \lambda_j^-$ vs. $\lambda_j^+ > \lambda_j^-$ is easily rejected at the one per cent level as $F(1,241)=9.87$. This is not surprising given the total long-run effects of positive and negative changes in the cash rate obtained in equations (8) and (9) are so different. One can thus conclude that in the context of credit cards there is strong evidence of the amount asymmetry. As to the stickiness hypothesis, we have tested another null hypothesis in which $\eta_j^+ = \eta_j^-$. Given $F(1,241)=6.70$, the null is once again rejected at the one per cent level of significance. Hence, based on these results, one can argue that there is convincing evidence for the existence of both the amount asymmetry and stickiness hypothesis in the context of Australia's credit card rates.

4.2. Policy implications

Although our results are consistent with previous studies, our proposed model is capable of capturing all possible types of asymmetric movements in the data. An overwhelming majority of previous studies suggest that there is a great deal of asymmetry in the short-run changes in the credit card loan rates. Researchers from other countries which could be exposed to banks' asymmetric behaviour in setting their various interest rates may also find our proposed model and results useful. This paper can assist borrowers and relevant government regulators to quantify the extent of the asymmetric behaviour exhibited by the banking industry as a whole. Borrowers are entitled to know why rate rises have been passed onto them faster than rate cuts and vice versa. This paper benefits all borrowers through a better understanding of credit card rates, the central bank through specific knowledge of the effects of its monetary policy on money market interest rates, and the regulators in enhancing

transparency and competition in consumer lending. The results presented in this paper reveal that the incomplete credit card rate pass-through is relatively large and persistent, requiring in turn a closer government monitoring and scrutiny.

5. Conclusion

Banks' asymmetric behaviour in setting various loan rates has been a major cause of concern in the banking and finance literature. However, little empirical work has recently been conducted regarding the dynamic effects of changes in the cash rate on Australia's credit card interest rate. This paper is among a few studies that have attempted to model these two rates and trace out the lenders' dynamic asymmetric responses to changes in the funding cost over time. Specifically, this paper attempts to rigorously test the robustness of the rockets-and-feathers hypothesis in the Australian banking context.

This paper uses all available monthly time series data (January 1990 to October 2011) to model the intricacies associated with the dynamic interplay between the cash rate and the retail-variable interest rate for credit cards. Our aim is to modify the existing asymmetric models in the literature to capture two forms of asymmetries, namely, the amount and adjustment asymmetries. Since the credit card rate is not co-integrated with the cash rate, one cannot test for any evidence of the adjustment asymmetry. However, it is found that the credit card interest rate rises more quickly than it falls in response to changes in the funding cost. In other words, at an aggregate level, rate rises are passed onto the consumer faster than rate cuts. In terms of the magnitude of the pass-through parameter in our preferred short-run dynamic model (i.e., Model II), rate rises had a full one-to-one and instantaneous impact on the credit card rate. However, the effects of rate cuts were delayed up to three months. Based on our results, we found a very significant degree of downward rigidity and path-dependency in terms of changes in the credit card rate. Similarly to previous studies we found that when

the RBA reduces the cash rate the Australian banks are reluctant to lower their retail credit card rates and do their best to delay such an action for as long as possible. Our proposed models can also be easily adapted to examine the full extent of price stickiness in the context of other types of retail and wholesale interest rates, including those for personal loans, mortgages and business loans.

6. Agenda for future research

In this paper we have established that the rockets-and-feathers hypothesis is applicable at an aggregate level in Australia's credit card market. On our agenda for future research, our aim will be to purchase consistently-defined weekly (credit card) interest rate data for over 100 bank and non-bank financial institutions from CANSTAR (www.canstar.com.au). Such a disaggregated and weekly study can reveal in which bank or non-bank institutions interest rate pass-through is more incomplete and the extent of the stickiness and asymmetric adjustment is greater. We can then describe the main characteristics of such lending institutions. For example, are the leaders of rate rises mainly big banks, and if so, which one? Have such banks maintained a similar position in the market through the sample period? One can identify the bank-specific opportunistic behaviour in the credit card market and reveal excessive profiteering by the parties involved due to market inefficiency or tacit collusion. The results of such a large-scale project will enable us to identify the leaders and followers of retail rate changes (if any). This type of information could provide consumers with more specific forward information about the expected dynamics, the extent of interest rate pass-through and the "best bank to choose".

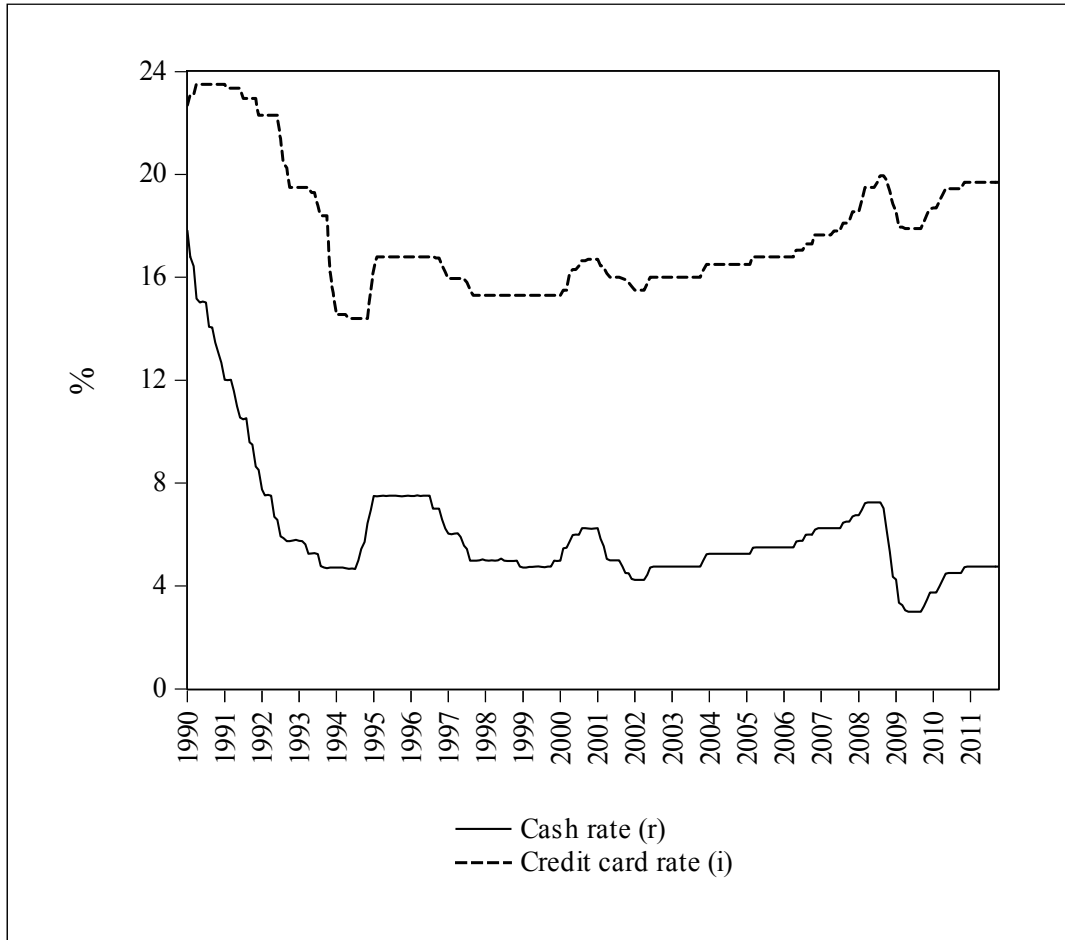
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Source: The Reserve Bank of Australia (2011).

Fig. 1. Cash rate versus credit card rate-January 1990–October 2011

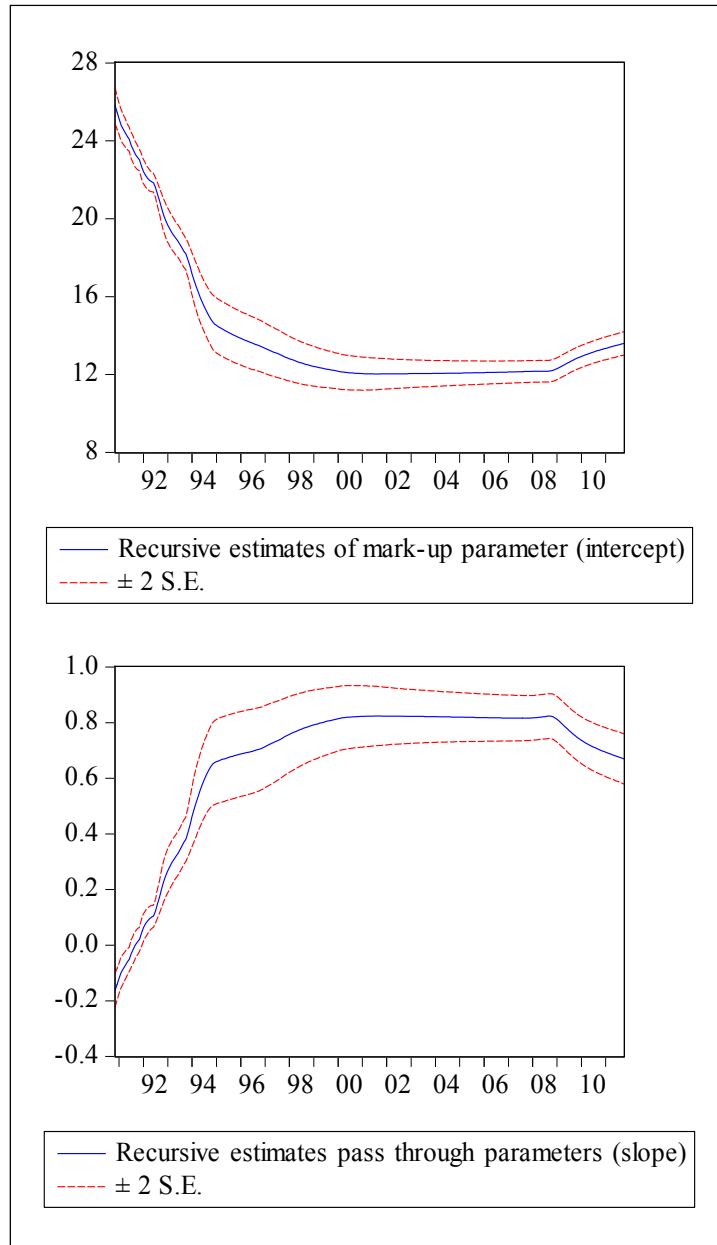


Fig. 2. Recursive estimates of the mark-up and pass-through parameters for $i_t = 13.59 + 0.669r_t$.

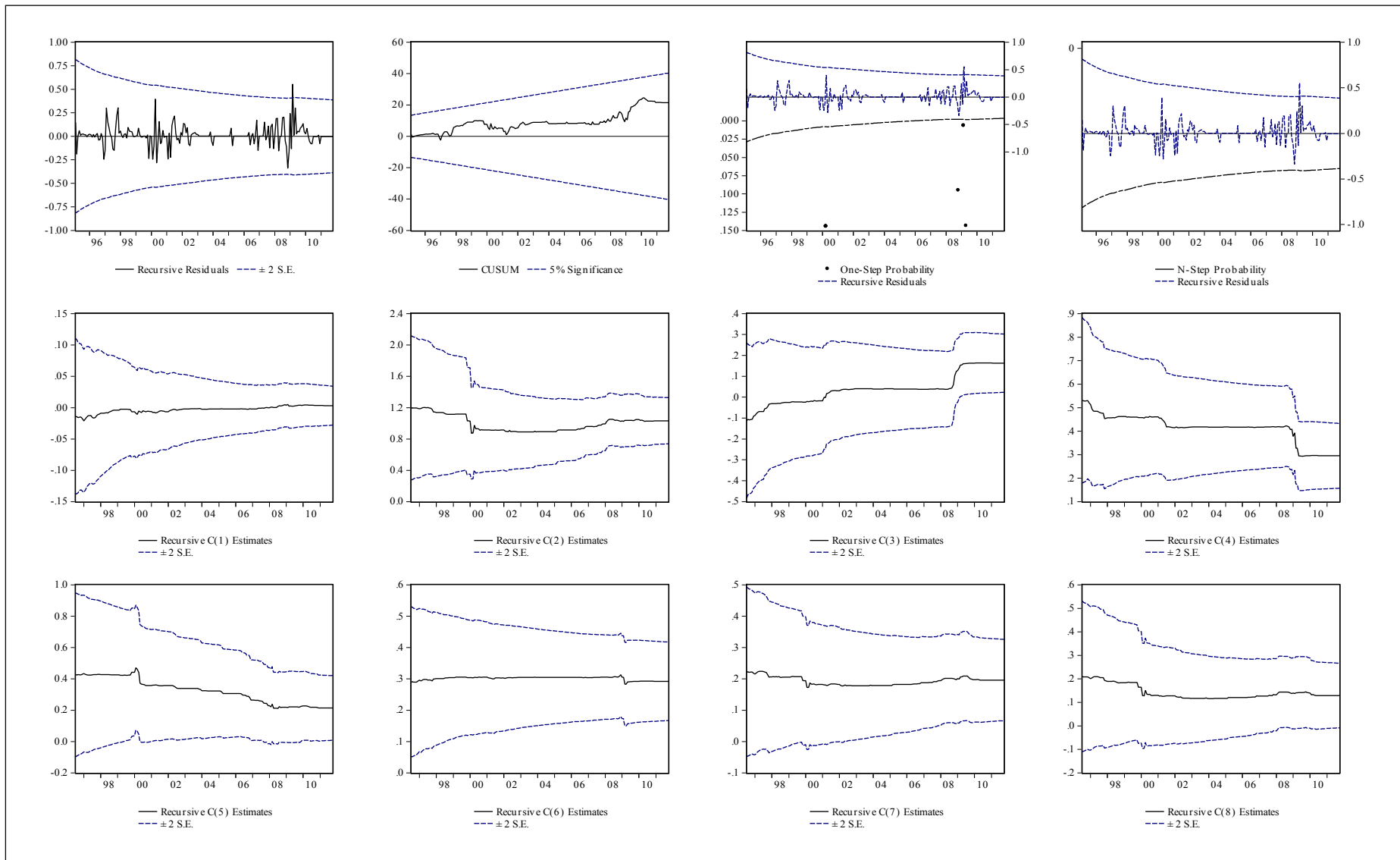


Fig.3.Graphical tests for stability of model II

Table 1

Descriptive statistics and unit root test results (January 1990–October 2011).

Description	r_t	Δr_t	i_t	Δi_t
Mean	6.511	-0.043	17.691	-0.011
Maximum	18.180	0.720	23.500	1.050
Minimum	3.000	-1.260	14.400	-2.200
Std. Dev.	3.199	0.255	2.435	0.247
Skewness	2.330	-1.691	1.049	-2.957
Kurtosis	7.960	8.297	3.216	29.535
Jarque-Bera	523.0**	445.9**	48.5**	8037.2**
ADT test				
t stat.	-2.426	-6.854**	-1.871	-10.443**
Optimal lag	2	1	1	0
DF-GLS test				
t stat.	-0.896	-5.908**	-0.750	-5.16**
Optimal lag	2	1	1	11
KPSS test				
LM stat.	0.274**	0.095	0.413**	0.048
Optimal Bandwidth	12	10	12	8
Lee and Strazicich test				
t ratio	-2.693	-6.856**	-2.2	-11.49**
Optimal lag	6	1	5	0
Break dates	1995:01 2009:03	1992:03 1995:01	1993:10 2009:02	1994:02 1997:09

Variable definition: r_t is the cash rate. i_t denotes the credit card rate.

Notes: (1) The choice between the crash model and the trend break model in the Lee and Strazicich (2003) test was based on the statistical significance of the corresponding parameters. (2) ** indicates that the corresponding null hypothesis is rejected at the 1 per cent level of significance.

Table 2Long-run relationship between the cash rate (r_t) and i_t .

Variable	Dependent variable:	
	i_t	
	Coefficient	t -Stat.
C	13.592**	45.59
r_t	0.669**	14.82
$H_0 : \theta_1 = 1$	$F(1,260)=53.5^{**}$	
$H_1 : \theta_1 < 1$		
R^2	0.458	
Overall F test	219.5**	
Residual unit root tests:		
ADF test		
t stat.	-1.239	
Optimal lag	1	
DF-GLS test		
t stat.	-1.677	
Optimal lag	2	
KPSS test		
LM stat.	0.348**	
Bandwidth	12	
Lee and Strazicich test		
t ratio	-3.878	
Optimal lag	8	
Break dates	1993:09	
	1997:01	

Variable definition: r_t is the cash rate. i_t denotes the credit card rate.

Notes: (1) The choice between the crash model and the trend break model in the Lee and Strazicich (2003) test was based on the statistical significance of the corresponding parameters. (2) ** indicates that the corresponding null hypothesis is rejected at 1 per cent level of significance.

Table 3
Estimated short-run dynamic asymmetric models.

Variable	Model I: Δi_t		Model II: Δi_t	
	Coefficient	<i>t</i> -Stat.	Coefficient	<i>t</i> -Stat.
<i>C</i>	-0.007	-0.26	0.003	0.20
Δr_t^+	1.019**	6.82	1.032**	6.99
Δr_t^-	0.159*	2.21	0.163*	2.34
Δr_{t-3}^-	0.291**	4.01	0.295**	4.28
Δi_{t-1}^+	0.217*	2.09	0.213*	2.07
Δi_{t-1}^-	0.291**	4.65	0.292**	4.67
Δi_{t-10}^-	0.200**	3.03	0.196**	3.02
Δi_{t-12}^-	0.134*	2.00	0.129*	2.01
EC_{t-1}^+	0.002	0.12	-	-
EC_{t-1}^-	-0.012	-0.53	-	-
R^2	0.419		0.418	
\bar{R}^2	0.397		0.401	
Overall <i>F</i> -statistic	19.1**		24.7**	
Akaike info criterion	-0.401		-0.415	
Schwarz criterion	-0.259		-0.302	
<i>DW</i>	1.96		1.97	
Serial correlation LM Test:				
2 lags	$F(2,237)=2.134$		$F(2,239)=1.832$	
4 lags	$F(4,235)=1.370$		$F(4,237)=1.323$	
6 lags	$F(6,233)=1.107$		$F(6,235)=1.108$	
8 lags	$F(8,231)=0.846$		$F(8,233)=0.842$	
10 lags	$F(10,229)=0.901$		$F(10,231)=0.876$	
12 lags	$F(12,227)=1.011$		$F(12,229)=1.013$	
Ramsey RESET	$F(1,238)=0.909$		$F(1,240)=0.776$	
Jarque-Bera	$\chi^2(2)=34763^{**}$		$\chi^2(2)=35339^{**}$	
Heteroskedasticity test:				
ARCH (1)	$F(1,246)=0.007$		$F(1,246)=0.006$	
ARCH (2)	$F(2,244)=0.871$		$F(2,244)=0.819$	
ARCH (3)	$F(3,242)=0.582$		$F(3,242)=0.549$	
ARCH (4)	$F(4,240)=0.434$		$F(4,240)=0.407$	
ARCH (5)	$F(5,238)=0.347$		$F(5,238)=0.326$	
ARCH (6)	$F(6,236)=0.289$		$F(6,236)=0.272$	
White test	$F(9,239)=0.642$		$F(7,241)=0.803$	
Breusch-Pagan-Godfrey	$F(9,239)=1.189$		$F(4,241)=1.482$	
Chow forecast test:				
2007M01-2011M10	$F(58,181)=0.410$		$F(58,183)=0.380$	
2008M01-2011M10	$F(46,193)=0.481$		$F(46,195)=0.456$	
2009M01-2011M10	$F(34,205)=0.404$		$F(34,207)=0.391$	

Notes: * and ** indicate that the corresponding null hypothesis is rejected at 5 and 1 per cent level of significance, respectively.