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By

**Ralf Becker, Denise R Osborn, Dilem Yildirim**

Centre for Growth and Business Cycle Research, Economic Studies,
University of Manchester, Manchester, M13 9PL, UK

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# A threshold cointegration analysis of interest rate pass-through to UK mortgage rates

**Ralf Becker[1], Denise R Osborn[1] and Dilem Yildirim[2]**

[1]Centre for Growth and Business Cycle Research
Economics, School of Social Sciences, The University of Manchester, UK

[2]Economics, Middle East Technical University, Ankara, Turkey

## Abstract

This paper empirically analyses the interest rate transmission mechanism in the United Kingdom by exploring the pass-through of the official rate to the money market rate and of the market rate to the mortgage rate. Potential asymmetries, due to financial market conditions and monetary policy, lead to the use of a nonlinear threshold error-correction model, with hypothesis tests based on non-standard bootstrap procedures that take into account the discrete nature of changes in the official rate. The empirical results indicate the presence of substantial asymmetries in both steps of the process, with these asymmetries depending on past changes in the money market rate and whether these are motivated by official rate changes. Generalized impulse response function analysis shows that adjustments differ with regard to the sign and magnitude of interest rate changes in a way that is consistent with conditions in the interbank and mortgage markets over the recent period.

JEL classification: C51, C52, G21

Key Words: Interest Rate Transmission, Mortgage Rates, Nonlinear Cointegration

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## 1. Introduction

The principal tool of monetary policy, as conducted by many central banks in developed and developing countries around the world, is the official short-term interest rate. By varying its official rate, the central bank aims to influence the retail loan and deposit rates offered by commercial banks to non-financial institutions and individuals, in order to achieve its aims for inflation and output. However, as has become clear during the recent credit crunch, the "pass-through" from official to commercial interest rates is neither necessarily immediate nor one-to-one. Indeed, it is now evident that the money market itself plays a key role in the interest rate transmission process, with the rates at which commercial banks provide short-term loans to each other in this market reflecting demand and supply considerations, as well as the current official interest rate.

In relation to the key role it plays in determining the effectiveness of monetary policy, there is a surprisingly scant literature on the pass-through from official to retail interest rates[1]. Nevertheless, recent empirical contributions to this literature (including Hofmann and Mizen, 2004, Sander and Kleimeier, 2004, Payne, 2007, Fuertes, Heffernan and Kalotychou, 2009) find strong evidence of nonlinearities, with retail rates responding asymmetrically to disequilibrium in relation to the official rate or its proxy. However, these studies typically ignore the role of the money market (Hofmann and Mizen, 2004, Payne, 2007, Fuertes *et al.*, 2009). Although Sander and Kleimeier (2004) examine retail rates in relation to both official and money market rates, they treat these as providing competing explanations for observed retail rates, rather than examining the interactions between them.

The recent credit crunch has, however, focused attention on the role of money markets. Indeed, the historically high spread for money market rates over official rates at certain times during this period has highlighted the crucial role played by these markets for the determination of both the level of retail interest rates and the availability of funds. While the operation of the money market has undeniably been affected by the abnormal conditions of the credit crunch, nevertheless this has also served to emphasise the lack of research to date about the nature of the pass-through from official interest rates to money market rates and how these, in turn, affect retail rates.

Introducing money market considerations points to a two-stage transmission process, namely from official rates to money market rates and from money market rates to retain interest rates. The only study that considers such a two-stage process is de Bondt (2005), who examines this through a three equation linear system. However, in addition to the evidence noted above for the pass-through to retail rates, Sarno and Thornton (2003) find that the transmission from the

[1] See de Bondt (2005), who provides a useful summary of the literature relating to individual euro area countries.

federal funds to the US Treasury bill rate is nonlinear, underlining the need to consider both stages of the transmission for an adequate understanding of this process.

This paper analyses the interest rate transmission mechanism in the United Kingdom by exploring both the pass-through of the official rate to the money market rate and subsequently the money market rate to the retail mortgage rate. The mortgage rate is selected for study since it is the key interest rate in terms of household expenditure and, consequently, is the “headline” rate used by the press for interpreting the impact of monetary policy changes by the Bank of England. Two sample periods are used in our analysis, namely one ending early in 2006 that does not include the credit crunch and may be considered a “normal” period and an extended sample to August 2008 that includes a period in which the credit market was under considerable stress (credit crunch).

Methodologically, a threshold cointegration relationship is employed, in line with other pass-through studies. Unlike previous studies, however, both pass-through stages are analysed in this context. Further, our methodology relies on nonstandard tests for nonlinearity that recognise the inherent unidentified parameter problem, in the spirit of Balke and Fomby (1997), Enders and Siklos (2001) and Hansen and Seo (2002). Indeed, the application of such tests in our context requires the development of a new bootstrap testing procedure, due to the discrete nature of changes in the official Bank of England rate. In contrast, previous UK studies (including Heffernan, 1997, Hofmann and Mizen, 2004, and Fuertes *et al.* 2009) ignore this important characteristic of the data and apply tests that assume continuous variables.

The remainder of the paper is organised as follows. Section 2 reviews the background literature, while Section 3 describes our data. The econometric methodology, including our new testing procedure, is discussed in Section 4. Substantive empirical results are presented in Sections 5 and 6, with the former showing the estimation results and the latter providing discussion, including generalized impulse response functions (Koop, Pesaran and Potter, 1996) that facilitate interpretation. Section 7 contains some concluding remarks, while an Appendix includes additional results.

## 2. Previous Literature

Money market rates are the marginal costs of funds faced by banks. However, due to adjustment costs (namely the costs to banks of changing mortgage rates), banks may not adjust their mortgage rates in response to very small market rate changes and/or changes that are expected to be temporary. Consequently, when they have some monopolistic power, banks may wait for large changes and/or a sequence of small changes to accumulate, leading to asymmetry and state-dependence in the pass-through to retail rates. Although discussed in the context of base rate changes, a theoretical model of this type is developed by Hofmann and Mizen (2004).

A different perspective is given by the industrial organization literature, which examines the setting of retail rates in the context of increasing market deregulation. For example, Corvoisier and Gropp (2001) develop a simple theoretical model to explore the role of competition and test this using data from euro area countries. These theoretical perspectives are, however, linked, since a competitive market will lead to a more complete and symmetric pass-through by increasing the cost of not adjusting.

Empirical analysis of asymmetries in the interest rate pass-through dates back to Neumark and Sharpe (1992), who apply a partial adjustment model with differing adjustment speeds depending on the sign of the past disequilibrium. Most subsequent contributions follow this broad approach, although it is now standard to represent the longrun equilibrium in terms of cointegration between official or market interest rates and the retail rate. Consequently nonlinear threshold error-correction models (ECMs) provide an appropriate modelling framework[2]. Burgstaller (2005) and de Bondt, Mojon and Valla (2005) examine mortgage rates in Austria and the euro zone, respectively, and find different responses to positive and negative disequilibrium deviations. However, both studies make the untested assumption that the threshold value giving rise to nonlinearity is zero.

Sander and Kleimeier (2004) and Payne (2007), however, apply the testing methodology of Enders and Siklos (2001) in order to allow for an endogenously determined threshold. Although these studies also find asymmetric pass-through for variable mortgage rates in the euro area and US, respectively, it is interesting that other work (Payne, 2006a, 2006b) concludes that adjustments for US fixed and 30-year mortgage rates are symmetric.

Initial analyses for retail mortgage rates in the UK by Heffernan (1993) and Paisley (1994) are in a linear context. However, the later studies of Heffernan (1997), Hofmann and Mizen (2004) and Fuertes *et al.* (2009) all find significant asymmetries in the pass-through from the official rate to retail rates. These studies consider possible asymmetries in relation to changes in the official rate, with Heffernan finding that the mortgage rate reacts slower when the official rate is rising than when it is falling. In contrast, using a later sample period and more disaggregated data, Fuertes *et al.* (2009) find quicker responses to rising official rates. This latter paper also uncovers faster adjustment for larger changes in the official rate, while Hofmann and Mizen (2004) detect faster adjustment when the deviation from equilibrium is widening or expected to widen.

One implication of these UK studies is that the appropriate nonlinear driver for the pass-through to mortgage interest rates may not be simply the disequilibrium, as assumed in applications based directly on Enders and Siklos (2001). In particular, possible nonlinearities associated with size effects should also be examined.

A common finding of the above studies, both for the UK and for other countries, is that the pass-through from official or money market rates to mortgage rates is incomplete in the longrun, and hence mortgage rates do not fully reflect the effects of monetary policy as conducted by the central bank. However, none of these papers recognise the two steps of the pass-through. As discussed in the Introduction, the pass-through from official rates to the money market rate also needs to be considered for an adequate understanding of the behaviour of retail interest rates. The single paper considering both steps is de Bondt (2005), who finds a complete pass-through for the first step but incomplete pass-through for the second.

Although there is an extensive literature on the relationship between interest rate series at different maturities, especially in the context of the expectations hypothesis of the term structure, the literature on the dynamics of the pass-through from official rates to money market rates is relatively thin. However, the findings of Kuttner (2001) emphasize the different impacts on money market rates of anticipated versus unanticipated monetary policy actions by the Federal Reserve, while Sarno and Thornton (2003) uncover strong evidence of a nonlinear adjustment between the federal funds rate and the 3 month Treasury bill rate.

In order to take account of the above findings, the present study incorporates both stages of the pass-through for the UK in a threshold ECM framework that permits the possibility that any asymmetry may be due to size effects. Nevertheless, this raises methodological issues for

[2] However, Humala (2005) employs a Markov switching vector autoregressive (MSVAR) model.

hypothesis testing, due to the discrete nature of interest rate changes implemented by the Bank of England[3]. As discussed in Section 4, we confront these issues by developing a new simulation-based procedure to test for cointegration and asymmetry that explicitly recognises the discrete nature of this variable. However, before detailing the methodology we employ, the data are examined in the next section.

## 3. Data

This study employs interest rate series measured at the end-of-month from January 1995 to August 2008[4]. These data are described in the first subsection, followed by a preliminary linear cointegration analysis.

3.1 Data Description and Sample Period

The starting point of 1995 for our analysis is selected in the light of changes during the 1980s and early 1990s in both the structure of mortgage lending and also in UK monetary policy. Prior to the 1980s, the UK mortgage market was dominated by building societies, who effectively operated an interest rate cartel. Although this cartel was broken down in the 1980s by deregulation and the large-scale entry of banks into the mortgage market, further legislation was passed in the mid-1990s to ensure that building societies were able to compete within a relatively equitable competitive environment. Stephens (2007) provides a detailed discussion and analysis of these changes, which took place alongside substantial shifts in UK monetary policy. However, monetary policy has been essentially stable since the UK adopted inflation targeting in October 1992. Although interest rates were initially set at monthly meetings of the Chancellor of the Exchequer with the Governor of the Bank of England, with full independence given to the Bank of England in May 1997, researchers interested in the nature and impact of UK monetary policy typically find the period from around 1992 to be a single regime (for example, Benati, 2004, Kesriyeli, Osborn and Sensier, 2006).

The sample used in this paper ends in August 2008. This ensures that at the back end of this sample we observe some severely stressed market conditions. The placement of Fannie Mae and Freddie Mac in US federal conservatorship and the collapse of Lehman Brothers in September 2008 triggered unprecedented intervention of governments and central banks into the banking sector and hence into money and credit markets. We therefore exclude the period from September 2008 from our analysis, due to these extraordinary events.

In the light of the abnormal circumstances prevalent in money and credit markets in the period running up to September 2008, the empirical analysis that follows is undertaken for two sample periods, namely January 1995 to January 2006, which we judge to be a relatively unstressed period, and January 1995 to August 2008, which clearly includes the beginning of the credit crunch. The former of these is our reference sample and some results are presented only for this sample. Whenever appropriate, we also comment on the robustness of results to the extension of this sample period and, when the extended sample results shed light on recent developments, we comment more thoroughly on these. Nevertheless, it should be noted that the primary aim of the paper is to investigate the interest rate pass through mechanism and recent events will merely

[3] The empirical analysis of de Bondt (2005) replaces the official rate with an overnight rate, apparently to avoid the discrete and infrequent changes exhibited by the official series (de Bondt, 2005, p.48).
[4] All data are from the Bank of England database, at www.bankofengland.co.uk/statistics/index.htm.

be used to provide sensible robustness checks. It is not the purpose of this paper to provide a detailed empirical analysis of the events of the credit crunch.

Figure 1 shows the official Bank of England interest rate (or base rate)[5], together with the one month London inter-bank offer rate (denoted as LIBOR) and the spread between these rates (LSPREAD). Overall, Figure 1 gives the impression of a complete or near complete pass-through, of official rates to the money market. Only during the final part of this period did the money market rate trade at a persistent positive premium relative to the base rate. The discrete nature of the official rate is also evident from Figure 1, with monetary policy typically implemented by the Bank through one quarter (or 25 basis points) changes in the rate. Indeed, there are relatively long sequences where the rate is unchanged, notably the fifteen months from November 2001. On a relatively small number of occasions, the official rate changes by ±0.50, but no larger shifts are observed over this period.

This study uses the LIBOR rate as the market rate, since this is the reference rate for (sterling) borrowing and lending in the London interbank market; it is also the market rate employed in the UK studies of Paisley (1994) and Heffernan (1993). Although the LIBOR is calculated for a range of maturities, from overnight to 12 months, the one-month maturity rate is selected for analysis since this shows the highest correlation (for both levels and changes) with the mortgage rate. Our use of correlation analysis to select the market rate that provides the appropriate marginal cost measure follows de Bondt (2005).

The mortgage rate is the average standard variable mortgage rate (SVR) of banks[6], which reflects the general rate of interest paid by borrowers. Miles (2004) indicates that at the end of 2003 around 35% of mortgage loans were at standard variable rate, while fixed and discounted variable mortgage made up around 25% and 18% of total loans, respectively. As seen in Figure 2, LIBOR and the mortgage rate are highly correlated, although with a significant mark-up (MSPREAD) that generally fluctuates between 1 and 2 percentage points. The data seem to suggest that this spread widens in the latter part of the period. Despite the high correlation of 0.971 between the mortgage rate and the LIBOR, the former is a little less volatile with a standard deviation of 0.986 compared with 1.180 for LIBOR.

Although closely correlated in levels, the correlation between changes in the mortgage and LIBOR rates is substantially lower at 0.457, than the correlation between changes in the base rate and LIBOR, which is 0.665. This suggests that the pass-through from LIBOR to the mortgage rate may be imperfect[7].

[5] According to the Bank of England database, this series is the average of four major clearing banks' base rates. However, from May 1997 this series is identical to the official rate, except for the specific days when interest rates change. The series given in the Bank's database as the official rate prior to May 1997 is typically 0.12 percentage points lower than this rate. However, over this earlier period the series used is identical to the interest rate values reported in the Bank of England's *Quarterly Bulletin* in relation to monetary policy decisions taken by the Chancellor and the Govenor of the Bank.

[6] The data series for the average SVR of banks was discontinued after December 2007. In order to extend the series to the entire sample period, we regressed the average bank SVR on a constant and the available series for the combined mortgage rates of banks and building societies over January 1995 to December 2007. This estimated relationship was then used to generate fitted values for the banks' average SVR for January to August 2008. A check of this methodology applied to a shortened sample (and checked against actual SVR) confirmed that it delivers satisfactory results. These results are available on request.

[7] If the degree and/or speed of the pass-through is not complete, an increase (decrease) in LIBOR will result in a decrease (increase) in MSPREAD because of small and/or slow responses of the mortgage rate.

3.2 Preliminary Cointegration Analysis

As is frequently the case for interest rates, a conventional unit root analysis (based on ADF and Phillips-Perron tests[8]) does not reject the null hypothesis of a unit root in each of our three series at the 10 percent level. Although macroeconomic arguments may point to the stationarity of interest rates, these data have statistical properties that are associated with nonstationary, or near-nonstationary, $I(1)$ series. Consequently, and following earlier studies, we proceed to a cointegration analysis of the pass-through.

A linear cointegration analysis provides insight into the relationships between these variables. Denoting observations on the base rate, LIBOR and mortgage rate as $brate_t$, $libor_t$ and $mrate_t$, respectively, Table 1 presents Johansen cointegration test results for a three variable system, as well as for each of the two bivariate subsystems, namely ($brate_t$, $libor_t$) and ($libor_t$, $mrate_t$), over the sample periods ending in both January 2006 and August 2008. The evidence for a cointegrating relationship between *libor* and *mrate* is robust to the period considered; indeed, the test statistics for this vary little when the longer sample is considered. Further, there is clear evidence of at least one cointegrating relation in the trivariate system. However, the apparent linear cointegration between the base rate and the money market rate that exists up to January 2006 in the bivariate system breaks down over the longer sample. Further, the evidence of a second cointegrating relation in the trivariate system is not compelling over the shorter sample and apparently disappears completely over the longer period.

The conclusion from this preliminary analysis is that UK mortgage rates are linked with money market rates over the longrun, with the apparent breakdown of the relationship between base rates and money market rates (and, by implication, between the base and mortgage rates) pointing to potential pitfalls in omitting the money market from an analysis of the pass-through to mortgage rates.

Nevertheless, the analysis of Table 1 ignores the possibility of nonlinearity. The next section discusses the econometric methodology we develop for dealing with possible nonlinear adjustment for interest rate changes, before results are considered in Section 5.

## 4. Econometric Methodology

In common with many other studies, our pass-through analysis employs single equation modelling under the assumption of weak exogeneity. More specifically, it is assumed that the official rate determined by the Bank of England is weakly exogenous to the market rate, which in turn is weakly exogenous to the mortgage rate, since banks' retail rates are not expected to affect market rate movements (de Bondt *et al.*, 2005).

This section first discusses the (linear and nonlinear) ECM models on which our empirical analysis is based (subsection 4.1), with subsection 4.2 then outlining the approach we take to model specification and estimation. Our bootstrap testing methodology, which explicitly allows for the characteristics of the interest rate data, is detailed in subsection 4.3. A final subsection outlines the nature of the generalized impulse response functions later used to aid the interpretation of our estimated models.

[8] These results are standard and not presented here. They are available upon request.

4.1 Error-Correction Models

Assuming that all interest rates are nonstationary $I(1)$ variables with long run cointegrating relationships existing between the interest rate pairs, and with the exogeneity assumptions already noted, then linear ECMs for the two steps of the pass-through imply

$$\Delta libor_t = \sum_{j=1}^{p} \phi_{1j} \Delta libor_{t-j} + \sum_{j=0}^{q} \theta_{1j} \Delta brate_{t-j} + \gamma_1 \left( libor_{t-1} - \alpha_1 - \beta_1 brate_{t-1} \right) + \varepsilon_{1t} \tag{1}$$

$$\Delta mrate_t = \sum_{j=1}^{p} \phi_{2j} \Delta mrate_{t-j} + \sum_{j=0}^{q} \theta_{2j} \Delta libor_{t-j} + \gamma_2 \left( mrate_{t-1} - \alpha_2 - \beta_2 libor_{t-1} \right) + \varepsilon_{2t} \tag{2}$$

where $\varepsilon_{1t}$ and $\varepsilon_{2t}$ are each assumed to be *iid* disturbances with zero mean and constant variance. The specification of (1) and (2) assumes that all effects of the base rate on the mortgage rate operate through the money market rate, an issue to which we return below.

The long-run coefficients $\alpha_i$ and $\beta_i$ ($i$ = 1, 2) measure the mark up (or down) and the degree of the pass-through in the long-run, with complete pass-through indicated by $\beta_i = 1$ and incomplete pass-through by $\beta_i < 1$. Further, $\theta_{i0}$ captures the impact response of an interest rate change, while $\gamma_i$ indicates the speed of adjustment for LIBOR (or the mortgage rate) to its long-run equilibrium. Notice that the intercept enters only through the longrun in (1) and (2), hence constraining this to lie in the cointegration space.

However, previous studies find that the symmetric speed of adjustment embodied in (1) and (2) may not adequately capture the interest rate pass-through. Following these papers, we employ threshold error-correction models, originally proposed by Balke and Fomby (1997), and further developed by Enders and Siklos (2001) and Hansen and Seo (2002). In this case, (1) and (2) are generalized to:

$$\Delta libor_t = \sum_{i=1}^{p} \varphi_{1i} \Delta libor_{t-i} + \sum_{i=0}^{q} \vartheta_{1i} \Delta brate_{t-i} + \gamma_{11} M_{1t} u_{1t-1} + \gamma_{12} \left( 1 - M_{1t} \right) u_{1t-1} + v_{1t} \tag{3}$$

$$\Delta mrate_t = \sum_{i=1}^{p} \varphi_{2i} \Delta mrate_{t-i} + \sum_{i=0}^{q} \vartheta_{2i} \Delta libor_{t-i} + \gamma_{21} M_{2t} u_{2t-1} + \gamma_{22} \left( 1 - M_{2t} \right) u_{2t-1} + v_{2t} \tag{4}$$

where $u_{1t} = libor_t - \alpha_1 - \beta_1 brate_t$ and $u_{2t} = brate_t - \alpha_2 - \beta_2 mrate_t$ are the disequilibria at $t$ in each of the two stages of the pass-through, $M_{it}$ ($i$ = 1, 2) is the regime operating at time $t$ for the $i^{th}$ stage, and $v_{1t}$, $v_{2t}$ are *iid* error terms with zero mean and constant variances. The regime is specified through an indicator variable that is expressed as the Heaviside function, such that

$$M_{it} = \begin{cases} 1 & z_{it} > \tau_i \\ 0 & z_{it} \le \tau_i \end{cases} \qquad i = 1, 2 \tag{5}$$

Even if the threshold variable $z_{it}$ is known, the threshold value $\tau_i$ is typically unknown. This implies that non-standard procedures are required for testing the presence of nonlinear cointegration between the interest rate pairs, which is the subject of the next subsection.

The threshold cointegration literature commonly adopts either the lagged disequilibrium or the changes in this disequilibrium as the threshold variable, corresponding to $z_{it} = u_{i,t-1}$ or $z_{it} = \Delta u_{i,t-1}$ for our case. The latter is referred to as M-TAR (momentum threshold autoregressive) adjustment by Enders and Siklos (2001), who suggest that it is appropriate when policy-makers smooth out large adjustments, and Payne (2007) adopts this specification when modelling the pass-through to retail interest rates in the US. Sander and Kleimeier (2004), on the other hand, consider both possibilities, together with a band-TAR model represented by $z_{it} = |u_{i,t-1}|$, which implies that the speed of adjustment depends on the *size* of the disequilibrium[9].

The only paper that considers using directly observed variables as possible drivers for the nonlinearity is Hofmann and Mizen (2004), who find that (actual or expected) changes in the official rate influence the speed of adjustment. Their specification defines the regime as being dependent on $\Delta brate_t$ ($\Delta brate_t > 0$ versus $\Delta brate_t \leq 0$), which not only assumes a known zero threshold, but also conflates zero with negative changes.

However, consider a disequilibrium value $u_{1,t-1}$ for the first-stage pass-through. From an initial position of equilibrium, this disequilibrium could arise because either (or both) the base rate or LIBOR changes. This suggests that the underlying driver for adjustment may not be the disequilibrium value $u_{1,t-1}$, but rather the change that gives rise to this disequilibrium, namely $\Delta brate_{t-1}$ or $\Delta libor_{t-1}$, and hence we consider each of these as the possible first-stage nonlinear drivers. For analogous reasons, $\Delta brate_{t-1}$, $\Delta libor_{t-1}$ and $\Delta mrate_{t-1}$ are examined as possible drivers for nonlinearity at the second stage of the pass-through, namely from the money market rate to the mortgage rate.

In addition to these changes, the absolute values of the corresponding variables (including $u_{i,t-1}$, $\Delta u_{i,t-1}$ and the relevant $\Delta brate_{t-1}$, $\Delta libor_{t-1}$ and $\Delta mrate_{t-1}$) are examined as possible nonlinear drivers, in order to examine asymmetry arising from size effects[10]. This may be particularly important for the pass-through, since the official rate frequently remains unchanged (as noted in Section 3). Although this constancy is not as evident for either LIBOR or the mortgage rate, nevertheless small changes may be essentially noise and hence generate different adjustment responses compared to large changes.

## 4.2 Model Specification and Estimation

Results are presented in Section 5 for both linear and threshold ECMs. The lag orders required in these models are specified in the linear framework (equations (1) and (2) without the equilibrium error terms). In particular, the Schwarz Bayesian criterion (SBC) is used in order to determine (separately) the lag orders $p$ and $q$ in (1) and (2), up to a maximum lag order of 12 in each case. These lags[11] are then carried over to the threshold ECMs of (3) and (4).

[9] In fact, Sander and Kleimeier (2004) allow three regimes, with different speeds of adjustment for disequilibrium values above the upper threshold $\tau$ and below the lower threshold - $\tau$.

[10] Indeed, contemporaneous $\Delta brate_t$ and its absolute value were also considered as the possible nonlinear driver for both (3) and (4), together with $\Delta libor_t$ and its absolute value for (4). However, stronger evidence of nonlinearity and cointegration were obtained using the lagged values of these variables.

[11] Intermediate lags were allowed to be dropped.

One practical issue in empirical modelling is the handling of "outlier" observations, which can play a particularly important role in a nonlinear context. In order that specific events do not unduly influence the models presented below, a dummy variable is included to account for a residual whose (absolute) value is larger in magnitude than 3 standard errors. To maintain the asymptotic distribution of test statistics, relevant step dummies are also added to the cointegrating relationship; see, for example, Doornik, Hendry and Nielsen (1998) for details. To ensure that linear and nonlinear models presented are comparable, the same dummy variables are included in all models for a specific (first or second) stage of the pass-through, with these dummy variables determined using the initial linear model.

Estimations of the linear ECMs are carried out using nonlinear least squares (NLS), as suggested by Stock (1987). In a preliminary step, the initial values of the long-run coefficients are found using ordinary least squares (OLS), with the initial values of the parameters of the short-term dynamics then obtained by OLS conditional on these. All parameters of (1) and (2) are then estimated simultaneously through NLS by minimizing the sum of squared residuals.

The threshold ECMs are estimated by modifying the sequential least squares approach of Hansen (1997). That is, for each potential threshold value $\tau_i$, which is typically in the middle 70% of the ordered values of the threshold variable, a threshold ECM is estimated through NLS using the same procedure as for a linear ECM. The estimate $\hat{\tau}_i$ is then determined by minimizing the sum of squared residuals over these estimations. Estimates of the cointegrating vector and the remaining parameters are then obtained by NLS conditional on this $\hat{\tau}_i$.

Although the disturbances ($\varepsilon_{1t}$, $\varepsilon_{2t}$) in (1)/(2) or ($\nu_{1t}$, $\nu_{2t}$) in (3)/(4) may be correlated, since each represents a "seemingly unrelated" system of equations, this is not taken into account in model estimation (or the subsequent impulse response calculation), due to the complexity of the non-linear procedure that is our principal focus. It may, however, be noted that the application of nonlinear ECM models reduces this correlation substantially compared to the linear model.

## 4.3 Testing for Threshold Cointegration

Prior to estimation of a threshold model such as (3) or (4), the presence of nonlinearity should be established. Although Balke and Fomby (1997) and Hansen and Seo (2002) undertake such a test based on an initial linear cointegration analysis, Enders and Siklos (2001) argue this is unsatisfactory due to the misspecification and low power of these tests in the presence of asymmetric adjustment. Instead, they develop a cointegration test that allows for a threshold adjustment under the alternative (of cointegration) and, if cointegration is established, test the null hypothesis of symmetric adjustment using a standard *F*-test.

We follow Enders and Siklos (2001) by testing for the presence of cointegration allowing for asymmetric adjustment through the model

$$\Delta u_{it} = \gamma_{i1} M_{it} u_{it-1} + \gamma_{i2} \left(1 - M_{it}\right) u_{it-1} + \sum_{j=1}^{q'} \delta_{ij} \Delta u_{it-j} + \eta_{it} \tag{6}$$

where $u_{it}$ ($i$ = 1 ,2) are as defined for (3) and (4), $q'$ is the required number of lagged changes of $\Delta u_{it}$ that ensures an *iid* structure for the error term, $\eta_{it}$, and the regimes for $M_{it}$ are defined in (5).

The null hypothesis of no cointegration, $\gamma_{i1} = \gamma_{i2} = 0$ in (6), is tested against the alternative of threshold cointegration. As the threshold value, $\tau_i$, defining $M_{it}$ is unidentified under the null hypothesis, the test statistic $\sup LM_T^{nc}$ is obtained by maximization over the range of possible $\tau_i$, defined as the central 70 percent of the distribution of the relevant $z_{it}$. The distribution of this test statistic is nonstandard and must be obtained by simulation.

Although Enders and Siklos (2001) provide critical values for the test applied to (6), these do not consider the possibility of a variable being discrete. Therefore, we develop a procedure that combines the proposal of Enders and Siklos (2001) to base cointegration testing on (6) with a fixed design model-based bootstrap along the lines of that suggested by Hansen and Seo (2002) in order to represent adequately the observed data features.

Specifically, the bootstrap *p*-values for testing the null hypothesis of no cointegration between the LIBOR and the base rate are simulated through the following algorithm:

i) Estimate the long-run equilibrium relationship $libor_t = \alpha_1 + \beta_1 brate_t + u_{1t}$ by OLS; obtain the estimates $\hat{\alpha}_1$ and $\hat{\beta}_1$.

ii) Generate the bootstrap DGP series $libor_t^*$ as

$$libor_t^* = \hat{\alpha}_1 + \hat{\beta}_1\, brate_t + v_{1t}^*,\ t = 1, 2, \ldots, T$$

where $v_{1t}^*$ is a random walk sequence with standard deviation set equal to the empirical residual standard deviation of $u_{1t}$ and $T$ is the sample size.

iii) Re-estimate the longrun relationship using $libor_t^*$ in conjunction with the actual $brate_t$ and obtain the residuals $u_{1t}^*$.

iv) Using the sequence $u_{1t}^*$, estimate the threshold model of (5) and (6), and calculate the bootstrap LM test statistic, $LM_T^{nc*}(\tau_1)$, for the null of $\gamma_{11} = \gamma_{12} = 0$ for each value of $\tau_1$ on the grid set $[\tau_{1L}, \tau_{1U}]$, where $\tau_{1L}$ and $\tau_{1U}$ are the 15th and 85th percentiles of the potential threshold variable $z_{1t}^*$.[12]

v) Obtain $\sup LM_T^{nc*}$ as

$$\sup LM_T^{nc*} = \sup_{\tau_1 \in [\tau_{1L}, \tau_{1U}]} LM_T^{nc*}(\tau_1).$$

vi) By repeating steps ii) to v), generate 50,000 bootstrap replications of $\sup LM_T^{nc*}$, and calculate the bootstrap *p*-value as the percentage of $\sup LM_T^{nc*}$ values that exceed the observed test statistic $\sup LM_T^{nc}$.

When the potential threshold variable in step iv) is endogenous, the corresponding bootstrap series is employed (that is, $u_{1t}^*$, $\Delta u_{1t}^*$, $\Delta libor_{t-1}^*$ or their absolute values). It is straightforward to

[12] For the purpose of the bootstrap tests the $q$' in (6) is estimated from the observed data and depends on the choice of $z_{it}$ in (5).

adapt this algorithm for the cointegration analysis between the mortgage rate and LIBOR, with $libor_t$ treated as exogenous.

The only case where this procedure is not employed for cointegration testing is when the potential threshold variable is the absolute change in the base rate. In this case, given the infrequency with which base rate changes of more than 25 basis points are observed in our sample period, the only feasible threshold to be examined in (6) is zero, which is therefore known and no unobserved parameter problem arises.

When, using the above test, the interest rate pairs are found to be cointegrated, the next step is to test the null hypothesis of symmetric adjustment, namely $\gamma_{i1} = \gamma_{i2} = \gamma_i$. Although Enders and Siklos (2001) employ a standard $F$-test, based on the estimate of $\tau_i$ obtained from the cointegration testing, they note that this could be problematic. In contrast, our approach continues to recognise that $\tau_i$ is unidentified under the null hypothesis being tested and we define a model-based bootstrap procedure similar to that of Balke and Fomby (1997), with one important modification. The approach of Balke and Fomby (1997), which is motivated by the standard Engle-Granger test, ensures stationarity of the disequilibrium term when the null of no cointegration is rejected. In our case, however, since the $\sup LM$ test that we use to establish cointegration has a two-sided alternative, rejection of the null does not guarantee the stationarity of $u_{it}$ in (6). As shown by Petrucelli and Woolford (1984) and Chan *et al.* (1985), necessary and sufficient conditions for stationarity are $\gamma_{i1} < 0$, $\gamma_{i2} < 0$ and $(1+\gamma_{i1})(1+\gamma_{i2}) < 1$.[13]

Therefore, our procedure first checks (for every bootstrap replication) that the estimated coefficients satisfy these stationarity conditions, before testing the symmetry null hypothesis $\gamma_{i1} = \gamma_{i2} = \gamma_i$. A $\sup LM$ test statistic is employed, as above. Again considering the first stage of the interest rate pass-through, the bootstrap $p$-values are obtained through a conditional model-based bootstrap procedure as follows:

i) Estimate the two-regime threshold model (6) under the restriction $\gamma_{11} = \gamma_{12} = \gamma_1$, obtain the coefficient estimates $\left(\hat{\gamma}_1, \hat{\delta}_{11}, .., \hat{\delta}_{1q'}\right)$ and residuals $\hat{\eta}_{1t}$; calculate the centered residuals $\hat{\eta}_{1t}^{\,c}$ as $\hat{\eta}_{1t}^{\,c} = \hat{\eta}_{1t} - \bar{\hat{\eta}}_{1t}$, where $\bar{\hat{\eta}}_{1t}$ is the sample mean of the residuals $\hat{\eta}_{1t}$.

ii) By randomly sampling with replacement from the centered residuals, obtain the sequence $\eta_{1t}^{*}$ for $t = 1, 2, \ldots, T$.

iii) Recursively generate the bootstrap DGP series $u_{1t}^{**}$ as

$$u_{1t}^{**} = (1+\hat{\gamma}_{11})u_{1,t-1}^{**} + \sum_{j=1}^{q'} \hat{\delta}_{1j}\, \Delta u_{1,t-j}^{**} + \eta_{1t}^{*},\ t = 1, \ldots.., T+100.$$

iv) Cut the first 100 observations of $u_{1t}^{**}$ and estimate (5)/(6) for the potential threshold variable $z_{1t}^{**}$ for each value of $\tau_1$, as above. Check the stationarity conditions for $u_{1t}^{**}$ after each estimation and, if the required conditions are satisfied, calculate the bootstrap LM statistic, $LM_T^{sa*}(\tau_1)$, for the symmetric adjustment null of $\gamma_{11} = \gamma_{12} = \gamma_1$. Discard the replication and return to step (ii) if the stationarity conditions are not satisfied.

[13] If the transition variable $z_{it}$ is chosen to be an absolute value we only require $\gamma_{i1} < 0$ for global stationarity.

v) Obtain the bootstrap sup statistic, $\sup LM_T^{sa*}$.

vi) Repeat steps ii) to v) 50,000 times and calculate the bootstrap $p$-value as

$$p = \frac{\#\{\sup LM_T^{sa*} > \sup LM_T^{sa}\}}{N_{stat}}$$

where $\sup LM_T^{sa}$ is the test statistic for symmetric adjustment obtained using the observed data and $N_{stat}$ is the number of replications in which the stationarity conditions are satisfied for $u_{1t}^{**}$.

The same procedure is applied to simulate the bootstrap $p$-values for the second step of the pass-through, from LIBOR to the mortgage rate. For both steps, this bootstrap test procedure is used for all cases except when the base rate change is considered as the threshold variable, with asymptotic test statistics being employed in this case due to the known threshold of zero.

## 4.4 Dynamic Analysis of Threshold Error-Correction Models

In order to provide further insights into the implications of the estimated nonlinear threshold cointegration models, generalized impulse response analysis is performed in relation to each of the two stages encapsulated in (3) and (4), and also for the system consisting of both equations.

Gallant, Rossi and Tauchen (1993) and Koop, Pesaran and Potter (1996) point out that, unlike linear models, the impulse response function of a nonlinear model is not (in general) independent of either the history of the series at the time of the shock or the sign and size of the shock. Further, due to the analytical intractability of these models, the impulse response functions have to be obtained by simulation. In the interest rate pass-through literature, the only study utilizing impulse response analysis of a threshold ECM is Sander and Kleimeier (2004), who do not, however, take account of the history dependent nature of the impulse response functions.

In this study, we follow Koop *et al.* (1996) and define the generalized impulse response functions for the two-regime threshold ECMs in (3) and (4) as

$$GI_Y(h, v_t, W_{t-1}, X_{t+h}) = E(Y_{t+h} | v_t, W_{t-1}, X_{t+h}) - E(Y_{t+h} | W_{t-1}, X_{t+h}), \;\; h = 0, 1, \ldots, H \tag{7}$$

where $GI_Y$ is the generalised impulse response function of the variable $Y$, which is $libor$ or $mrate$ depending on the stage of pass-through under analysis, $v_t$ is an arbitrary shock applied at time $t$, $W_{t-1}$ is the history (information set of all variables up to time $t-1$), $X_{t+h}$ is the information set of weakly exogenous variables to $t+h$ and $H$ is the horizon[14].

More specifically, our threshold ECM models have two regimes, corresponding to $M_{it} = 1$ and $M_{it} = 0$, in (5). To emphasize the nature of the regime-dependent adjustment in (3) and/or (4), we compare the generalized impulse response functions for shocks occurring in each regime. For the interest rate pass-through to the money market, consider a set of $k_1$ occasions for which $M_{1t} = 1$ and define $W_{t-1}$ to be the corresponding set of $k_1$ sequences of initial (lagged) values of *libor* and *brate* required in (3), namely $libor_{t-p-1}, \ldots, libor_{t-1}, brate_{t-q-1}, \ldots, brate_{t-1}$. Similarly, for these same $k_1$ specific periods for which $M_{it} = 1$, $X_{t+h}$ is the corresponding set of $k_1$ sequences of values $brate_t, \ldots, brate_{t+h}$.

[14] Both Gallant *et al.* (1993) and Koop *et al.* (1996) examine impulse response functions of nonlinear autoregressive models. We modify their approach for nonlinear univariate ECMs by assuming that the weakly exogenous variables are known to time $t+h$.

In order to calculate the generalized impulse response function in (7) conditional on $M_{it} = 1$ we simulate $Y$ forward from all $k_1$ histories. This forward simulation uses randomly drawn innovation terms from the empirical distributions of estimated model residuals. The difference between a particular simulation $Y_{t+h}|v_t, W_{t-1}, X_{t+h}$ and $Y_{t+h}|W_{t-1}, X_{t+h}$ is the additional (given) perturbation $v_t$. The generalised impulse response function (conditional on $M_{it} = 1$) is then obtained by first averaging across 10,000 simulations for every particular history and subsequently averaging across all $k_1$ histories for which $M_{1t} = 1$.

Generalized impulse response functions for the regime corresponding to $M_{it} = 0$ and for the regimes in (4) are obtained in an analogous way. Impulse response functions are also presented when the two stages of the pass-through are considered (base rate shocks being transmitted to the mortgage market via the money market). In this case, four regimes are possible for $(M_{1t}, M_{2t})$, since different regimes can apply for each of the stages.

# 5. Estimation Results

After discussing the results of our tests for cointegration (subsection 5.1), the estimated threshold ECM models are presented in the following two subsections. Although no diagnostic test results are presented for any ECM models, it may be noted that none of these models indicate the presence of either residual autocorrelation or conditional heteroscedasticity according to conventional tests applied to lag 12[15]. Indeed, the most marginal significance level for these is around the 10 percent level.

As the focus of this analysis is to shed light on the two different stages of the interest pass through, the discussion primarily relates to the sample (January 1995 to January 2006) which we judge to be largely free from the recent stresses caused by the sub-prime mortgage crisis and the subsequent more general financial crisis. In Section 6 we comment in more detail on the implications of the estimated models, including those arising from generalised impulse response functions.

## 5.1 Cointegration Tests

Using the testing methodology detailed in Section 4.3, based on the specification given in (6), Table 2 presents the results for both steps of the pass-through, estimated over both samples. The potential nonlinear drivers considered are discussed in Section 4.1 above[16].

Consider first the pass-through from LIBOR to the mortgage rate (namely, the second step of our pass-through analysis), for which Table 1 provides clear evidence for cointegration, albeit of a linear form. Cointegration is confirmed by the results in the right-hand panel of Table 2, irrespective of the potential threshold variables considered. Judged by the values and significance of the cointegration and asymmetry test statistics, especially in the "normal" period to January 2006, the strongest candidates for the nonlinear driver are the magnitudes (that is, absolute values) of changes in past rates, rather than the (signed) changes or the cointegration

[15] All models considered (for both sample periods) pass these tests at the usual 5 percent level. Indeed, all pass at 10 percent, except that the threshold ECM for the pass-through to the mortgage rate (for the extended sample) yields a $p$-value of 0.099 for the ARCH test.

[16] The use of contemporaneous $\Delta brate_t$ as the possible nonlinear driver for the first step, or $\Delta brate_t$ and $\Delta libor_t$ for the second step, together with the absolute values of these variables, yields quantitatively similar statistics to those reported in the table for the corresponding lagged value.

residuals. Thus, previous literature analyzing the pass-through that follows Enders and Siklos (2001) in assuming the nonlinear driver to be a cointegration residual or its change (or, as in Hofmann and Mizen, 2004, the observed interest rate changes), appears to have overlooked the potentially most important source of nonlinearity for mortgage rates.

Turning now to the pass-through from the base rate to the LIBOR rate, the results in the left-hand panel of Table 2 for the main sample period also provide evidence for cointegration and, for drivers other than changes in the base rate itself, for asymmetry of adjustment. Thus, in accord with the linear results for this period in Table 1, cointegration appears to be the norm between the base rate and LIBOR. However, when the longer sample to August 2008 is considered, and in contrast to the results for this period in Table 1, there is evidence that nonlinear cointegration continues to apply. This is, of course, associated with asymmetric adjustment, with this nonlinearity highly significant (at 1 percent or less) for the threshold variables $\Delta libor_{t-1}$ and $|\Delta libor_{t-1}|$. Consequently, conditions in the money market itself appear to be crucial for the adjustment process towards longrun equilibrium with the base rate, a finding which resonates with experience during the credit crunch period.

Given the evidence in Table 2, we proceed to estimate threshold ECMs for both stages of the pass-through, with the following two sub-sections discussing the resulting models.

### 5.2 Pass-Through to LIBOR

Table 3 summarises the results for three models that capture the pass-through from base rates to LIBOR. For reference a linear ECM is shown, but only for the shorter sample to January 2006, since Table 1 provides no evidence for linear cointegration over the longer period. Nonlinear ECMs, with threshold variable $|\Delta libor_{t-1}|$, are shown for both samples. Although the results of Table 2 do not clearly indicate the appropriate driver, $|\Delta libor_{t-1}|$ is selected as it yields the best fit (according to SBC, AIC and the residual standard error) for models estimated over both periods[17]. The left-hand panel of Table 3 shows estimates of the longrun equilibrium relationship, while parameters relating to the shortrun dynamics are in the right-hand panel. To conserve space, the shortrun coefficients associated with impulse dummy variables are not shown, but these are always individually significant at levels of significance of 5 percent or (typically) less.

Both the linear and the nonlinear ECM models estimated over the reference sample (to January 2006) are compatible with the pass-through to LIBOR being complete in the longrun. Indeed, the estimated coefficient of $brate_t$ is very close to unity in both models and the respective hypothesis tests have large $p$-values (0.869 and 0.763 for the linear and nonlinear ECM respectively)[18]. In both models the mark-up (captured by the intercept) is not significantly different from zero, which chimes well with the graphical evidence in Figure 1. (Note that values in brackets for coefficients are $p$-values for a null hypothesis of zero.)

[17] Specifically, $|\Delta \hat{u}_{1t-1}|$, $\Delta libor_{t-1}$, $|\Delta brate_t|$ and $|\Delta libor_{t-1}|$ were all investigated over this shorter sample, since these yield similar values, with very high significance, for both the cointegration and asymmetry tests. For the same reason, $\Delta libor_{t-1}$, $\Delta brate_t$ and $|\Delta libor_{t-1}|$ were considered over the extended sample.

[18] It is also notable that none of the step dummy variables included in the longrun specification are significant (at 5 percent), which is compatible with this relationship capturing enduring features of the relationship between money market and base rates. They remain in the specification as their impulse dummy counterparts in the shortrun dynamics are significant (coefficients for these are not shown, to conserve space).

The shortrun dynamics indicate that much of the pass-through is immediate and, further, tell an interesting story. The threshold model implies that when LIBOR changes by more than around ±0.1 percentage points in any month, there is a further adjustment in the next period to remove half of the resulting disequilibrium. As a careful inspection of Figure 1 makes clear, LIBOR sometimes anticipates base rate changes, which provides a rationale for why the occurrence of nontrivial changes in the money market rate is the driver for the nonlinear ECM specification. On the other hand, when LIBOR changes by very small amounts, the adjustment coefficient is not significantly different from zero, with the small changes in LIBOR presumably reflecting very shortrun and minor fluctuations in the money market. The adjustment speed in the linear model is, unsurprisingly, between the adjustment speeds of the two regimes, but this is an unreliable estimate due to the neglected nonlinearity.

Extending the sample to 2008 gives rise to a number of changes. There is now less evidence of complete pass-through and we find two dummy variables that, although not individually significant, seem to drive a lasting wedge between the base and LIBOR rates (consistent with what can be gleaned from Figure 1); this is also indicated by the significant mark-up. The August 2007 hike in the LIBOR rate, which was not mirrored by any increase in the base rate, was the start of a period in which the LIBOR rate persistently exceeded the base rate, and this is also reflected in different dynamic responses to base rate changes from that date[19]. Nevertheless, there is little change in the disequilibrium adjustment.

It is interesting to identify the underlying reasons for the dummy variables which are identified in these specifications. The January 2000 dummy (*D*0001) corresponds to millenium effects which (although details are not shown) are highly significant in all shortrun specifications. The *D*0708 dummy (August 2007) corresponds to the beginning of the Northern Rock crisis, which resulted in the nationalization of Northern Rock in January 2008 (*D*0801).

## 5.3 Pass-Through from LIBOR to Mortgage Rate

Results for the estimated ECM models for the pass-through from LIBOR to mortgage rates are shown in Table 4. These are analogous to those of Table 3 for the pass-through to LIBOR, with the estimated dynamic coefficients for the short-run model shown in right panel and the estimated longrun relationships in the left panel. Since Table 1 indicates linear cointegration for the mortgage rate pass-through irrespective of the time period considered, estimated linear and nonlinear ECMs are presented for both the reference and extended sample periods. The threshold models of Table 4 employ $|\Delta brate_{t-1}|$ as the nonlinear driver, since this leads to the strongest evidence of threshold cointegration and asymmetry in Table 2 over both periods[20]. In the light of the infrequency of base rate changes of more than a quarter of one percent, the only feasible threshold value is zero (predetermined), and hence the regimes separate months where the rate remains constant versus those where the base rate changes.

First we evaluate the results for our reference sample ending in January 2006. While the complete pass-through hypothesis cannot be rejected for the linear model (*p*-value 0.166), the estimated longrun equilibrium relationship in the nonlinear threshold cointegration model is not consistent with a complete pass-through. Not surprisingly, both models provide evidence of a significant mark-up of the mortgage rate over LIBOR.

[19] Tests of stability did not indicate a change in the coefficients for $\Delta libor_{t-i}$ ($i$ = 1, 2) at this date.

[20] For the reference sample, an ECM model with $|\Delta libor_{t-1}|$ as the threshold variable was also estimated, but that using $|\Delta brate_{t-1}|$ yielded the best fit according to SBC, AIC and the residual standard error.

The nature of the different responses to the longrun relationship, from the linear and nonlinear models, are plausible. With an assumption of linear cointegration, we find a rather sluggish adjustment coefficient of -0.164, whereas the nonlinear analysis reveals that the adjustment speed differs significantly depending on whether $|\Delta brate_{t-1}| > 0$ ($M_{2t} = 1$), in which case we observe very fast disequilibrium correction (-0.702) or $|\Delta brate_{t-1}| = 0$ ($M_{2t} = 0$) which is associated with sluggish, yet statistically significant, adjustment (-0.073). The latter reflects stability in the monetary policy stance. Disequilibria that occur in such an environment are not eliminated as swiftly as equilibria that arise from LIBOR movements that are backed by changes in the monetary policy instrument. This contrasts with the implications of the linear model, which indicates slow adjustment over all months. It is also interesting to note that misspecifying this equilibrium adjustment, as in the linear ECM model, results in richer short-term dynamics than those resulting from the more general nonlinear model. As evident from the results in the table, allowing for the nonlinear equilibrium correction substantially improves the fit compared to a linear specification, with the residual standard error being reduced by more than a quarter.

Overall, these models change little when the sample is extended to August 2008 to cover the beginning of the credit crunch period. In fact all the above findings regarding the presence (or not) of a complete 2nd stage pass-through and the equilibrium adjustment mechanisms remain valid for the longer sample, indicating a remarkable stability even in the presence of some severely stressed market conditions.

Extending the sample to August 2008 does, however, require the addition of a number of dummy variables. Most notably it requires the inclusion of the step dummy *D*0612 (December 2006) in the longrun equation and its corresponding impulse dummy (together with changed dynamics) in the shortrun specification. It is difficult to pin this dummy variable to a particular event, but it roughly coincides with the market's realisation that the decline in house price inflation, which began in the summer 2006, would be long lasting. The dummy for June 1997 (*D*9706) coincides with the conversion of Alliance & Leicester and Halifax from building societies to banks. This may well have had the effect of decreasing the competition between building societies and banks[21], resulting in an increase in the mark-up of banks' mortgage rates.

## 6. Discussion and Interpretation

In this section we will present and discuss the generalized impulse response functions (GIRF, see Section 4.4 for a short summary of the methodology) for our pass-through models and use these to comment on the nature of the two stage mechanism. We present GIRFs for a horizon (*h*) of up to 14 months. In the context of nonlinear models, impulse responses can vary with the sign and size of the shock and therefore impulse responses for four different shocks, $\pm 1\hat{\sigma}$ and $\pm 2\hat{\sigma}$, are presented, where $\hat{\sigma}$ is the relevant estimated residual standard deviation.

First we examine the response of LIBOR rates to shocks in the base rate for the nonlinear model, using the estimates of the reference sample period. The type of shock considered is a permanent monetary policy shock, i.e. a lasting increase or decrease in the policy rate. The impulse response functions shown in Figure 3 differentiate between the two regimes, in which (at the time of the initial shock) $|\Delta libor_{t-1}| > 0.090$ in the upper panel and $|\Delta libor_{t-1}| \leq 0.090$ in the lower panel; the vertical axis shows the percentage of the initial shock that is adjusted at a given

[21] See Heffernan (2005) for evidence suggesting that post conversion building societies adjusted their price setting behaviour to look more like that of a normal bank.

horizon. It is apparent that, in both regimes, the adjustment is asymmetric with respect to the sign of the base rate change. While, in general, negative shocks are not fully transmitted, positive shocks have a stronger than one-to-one effect on the LIBOR rate. Sander and Kleimeier (2004) and de Bondt (2005) explain overshooting to positive base rate shocks as a potential indication that banks increase their risk premium as a response to potentially increased default risk. It is interesting to note that the incomplete adjustment to negative shocks is most obvious when the money market is characterized by small changes in the previous period (Regime 2), indicating that base rate reductions do not fully translate to the money market when they do not occur in a trending market. On the other hand, base rate changes occurring in a moving money market (Regime 1) trigger strong overshooting. It should also be noted that there is a tendency for the asymmetries to be stronger for larger shocks.

For the second step of the interest rate pass-through mechanism, GIRFs are used to analyse the impact of permanent changes to both LIBOR and the base rate on the mortgage rate. Figure 4 shows the response to shocks in the LIBOR rate, where 100 again represents a full transmission of a LIBOR rate shock to the mortgage rate. It is immediately apparent that the mortgage rate reacts equally to positive and negative LIBOR rate shocks, which occurs because regimes are governed by the base rate. In contrast to the first stage of the pass-through, the second-stage adjustment is sluggish especially in the 2nd regime which is characterised by an unchanged base rate in *t*-1 (after one year less than 80 percent of the change has been passed through). Adjustment is stronger and nearly complete after one year if a LIBOR change was predated by a change in the base rate (1st regime). The size of the shock makes no discernable difference, again since the response is linear given the (base rate) regime.

The impulse response functions in Figure 5 illustrate how the mortgage rate responds to unanticipated changes in the base rate. As the interest rate pass-through is modelled in two stages, these impulse response functions also follow a two stage process. In the first instance, base rate shocks are transmitted to the LIBOR rate according to the threshold cointegration model in Table 3. These responses can be characterised as in Figure 3, but are merely of intermediate interest in this context. The simulated LIBOR rates are then used as the histories of money market rates relevant for the mortgage rate in the threshold cointegration specification of Table 4[22].

As this involves two threshold models with different threshold variables, four different regimes may be implied. Of these four, three regimes are analysed; first the *policy change* regime (Regime 1) in which period *t*-1 is characterized by a change in the base rate and a non-trivial change in the LIBOR rate (greater than 0.09 percentage points in magnitude). The second regime is a *LIBOR only change* regime, which represents the case in which the LIBOR rate changed non-trivially ($|\Delta libor_{t-1}| > 0.090$), despite there not being any change in the base rate. The third regime is the *stable* regime which is defined by no (non-trivial) change in either the base rate or the LIBOR. The potential fourth regime (base rate change but no corresponding change in the LIBOR) yields only three empirical observations over the sample period, and hence is not considered to be empirically plausible.

Comparing responses across regimes in Figure 5, it is clear that the mortgage rate adjusts more quickly after changes in the monetary policy stance (Regime 1) than otherwise, we also observe asymmetries with respect to the sign of the monetary policy shock. In general, positive base rate shocks are fully transmitted within between 6 months (*policy change* Regime) and 10

[22] The correlation between the residuals of the threshold ECMs given in Tables 3 (Step1) and 4 (Step 2) is assumed to be zero. Given an empirical correlation of -0.230 between the residuals, this is a reasonable assumption.

months (Regimes 2 and 3). Negative base rate shocks, however, fail to be fully transmitted to mortgage rates. These asymmetries are strongest in Regimes 2 and 3, which represent cases in which a monetary shock is not preceded by another change in the base rate. Following a similar finding for the effect of base rate shocks in the first-step pass-through (Figure 3), the asymmetries are somewhat stronger for larger shocks.

These results shed important light on the previous finding of asymmetries in the mortgage market in Fuertes *et al.* (2009). There the asymmetries are attributed to the structure of the mortgage market. From the analysis presented here, however, it transpires that the asymmetries arise primarily through asymmetries in the first-step of our pass-through process and hence ought to be explained in the interbank rather than the mortgage market. Without splitting the pass-through process into two steps this result would have been impossible to obtain.

A corresponding GIRF analysis was undertaken for the extended sample ending in August 2008. The results are qualitatively extremely similar and are therefore not reported here.

## 7. Conclusions

This paper investigates the transmission of interest rate shocks to the mortgage market, either induced by monetary policy or originating in the money market. In order to dissect interest shocks appropriately, the transition from monetary policy rates to mortgage rates is separated into two steps (from the base rate to LIBOR and from LIBOR to mortgage rates), allowing for the possibility of asymmetries in both steps. This reveals that asymmetries which appear to be in the mortgage market (namely, incomplete pass-through of base rate reductions to the mortgage rate, but complete pass-through of base rate increases), are really a feature of the money market rather than the mortgage market itself.

It transpires that nonlinearities play an important role in our analysis, as adjustment speeds to long-run equilibria typically vary significantly depending on some underlying state variable. In general we find that adjustment speeds are significantly greater when interest rate movements are motivated by clear monetary policy signals. The nonlinear analysis further reveals that the interest rate pass-through between the policy rate and the money market is complete, but that the pass-through from the money market to the mortgage market is short of being complete. An extended sample, reaching into the beginning of the recent credit crunch period, provides a robustness check on this analysis.

The modelling approach adopted in this paper, in addition to allowing for nonlinear cointegration between the different interest rates, also includes a novel approach to statistical inference by explicitly allowing for the discrete nature of base rate changes. Indeed, a general feature of our approach is the extensive use made of bootstrap inference, which is employed for testing the presence of both cointegration and nonlinearity.

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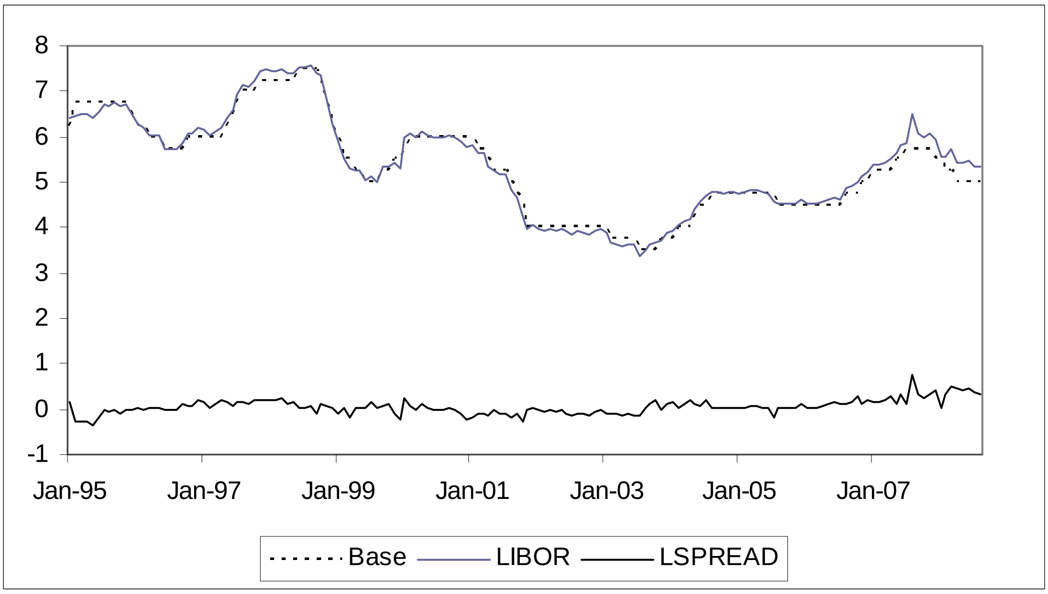


**Figure 1**: The Base Rate and one month London inter-bank offer rate (LIBOR), together with the difference between Libor and the Base Rate (LSPREAD), are shown over the extended sample from January 1995 to August 2008.

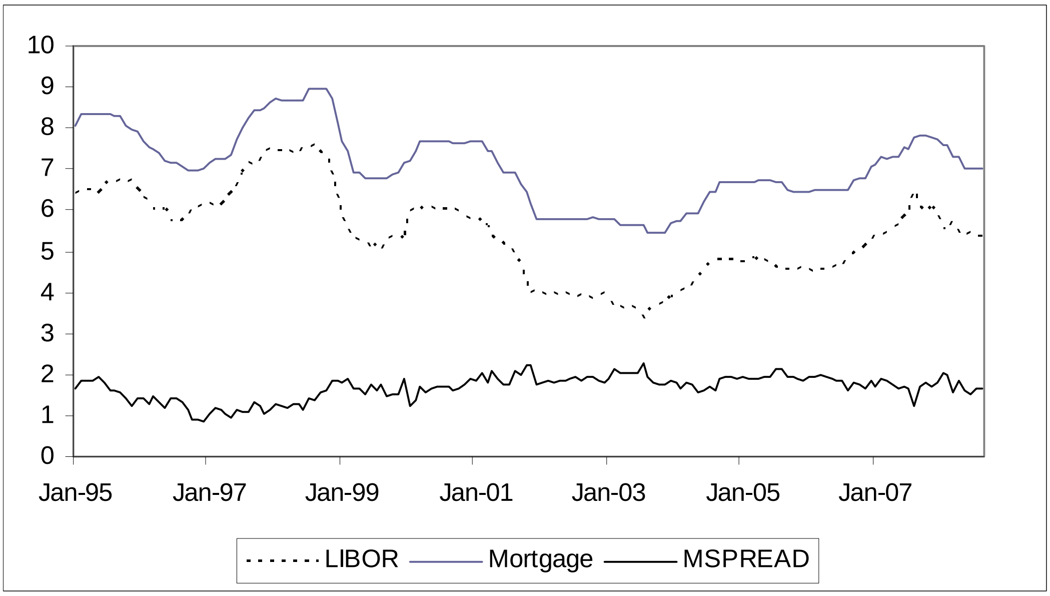


**Figure 2: O**ne month London inter-bank offer rate (LIBOR) and average standard variable mortgage rate of banks (Mortgage Rate), together with the difference between the Mortgage Rate and Libor (MSPREAD), are shown over the extended sample period from January 1995 to August 2008.

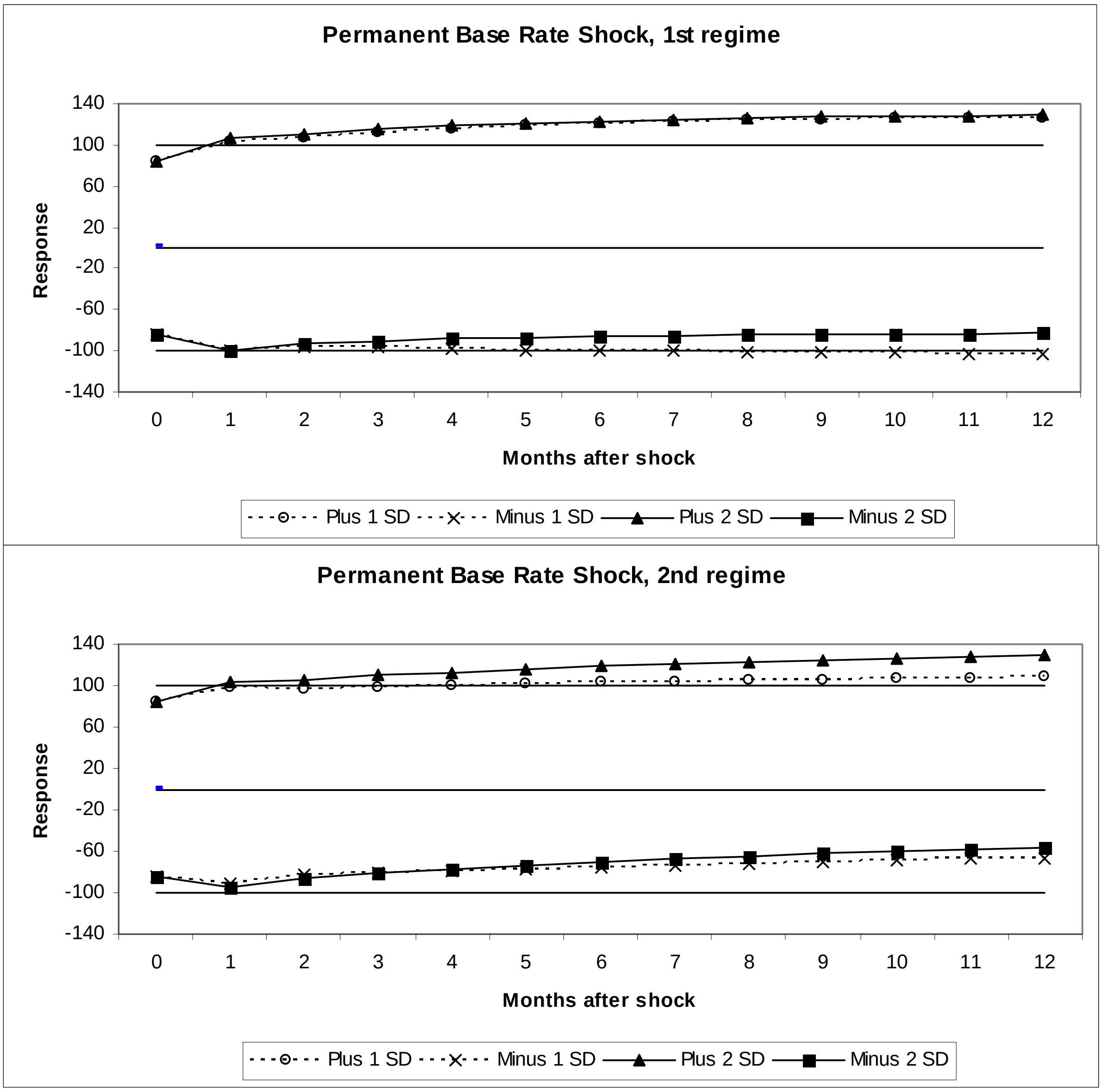


**Figure 3**: Dynamic Responses of LIBOR to Base Rate Shocks. Responses are given in percentage terms and are obtained through stochastic simulations of the two-regime threshold ECM using 10,000 replications. 1st and 2nd regimes refer to cases where $\left|\Delta libor_{t-1}\right| > 0.090$ and $\left|\Delta libor_{t-1}\right| \leq 0.090$, respectively.

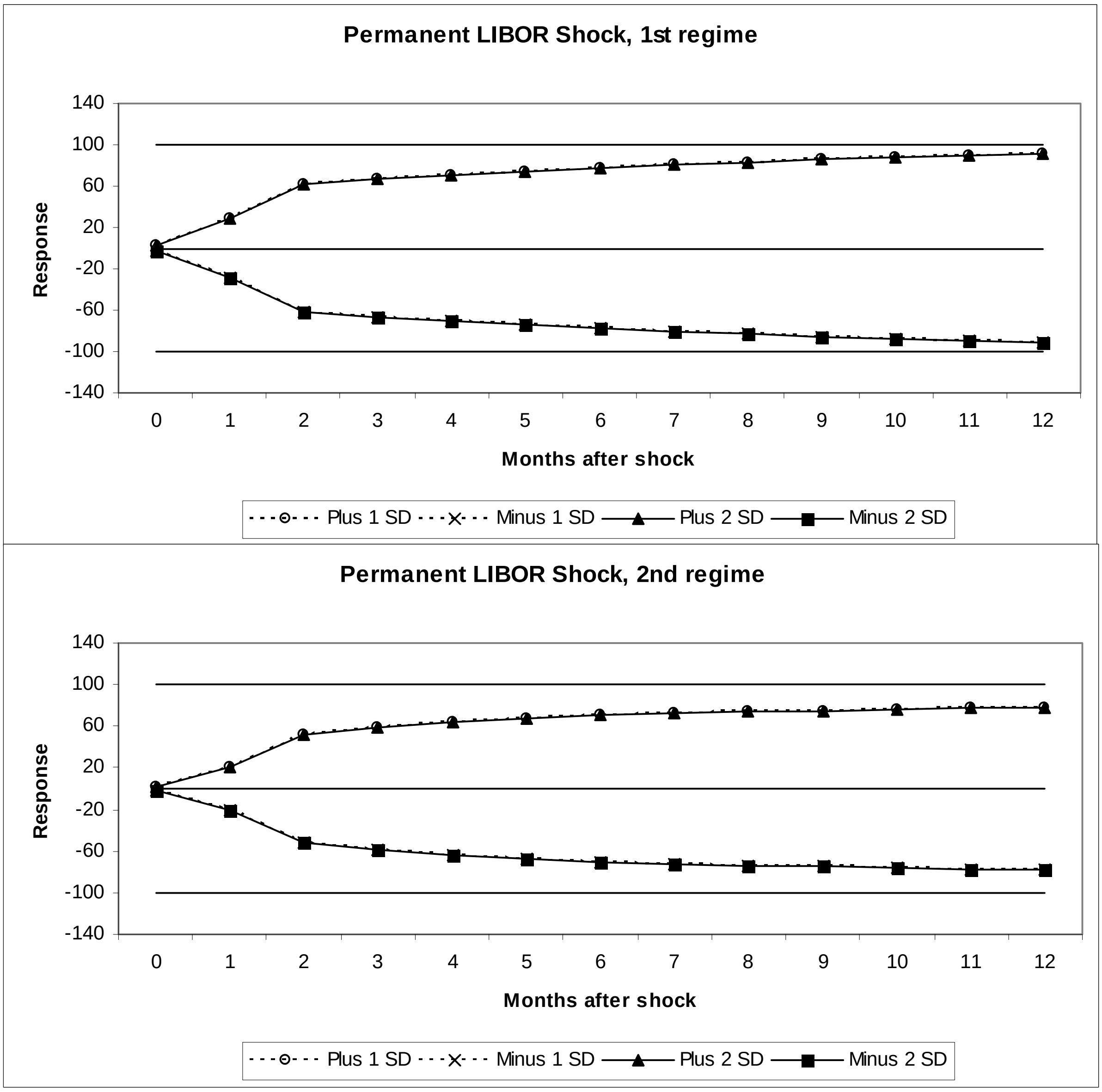


**Figure 4**: Dynamic Responses of the mortgage rate to LIBOR Shocks. Responses are given in percentage terms and are obtained through stochastic simulations of the two-regime threshold ECM using 10,000 replications. 1st and 2nd regimes refer to the cases where $|\Delta brate_{t-1}| > 0$ and $|\Delta brate_{t-1}| \leq 0$, respectively.

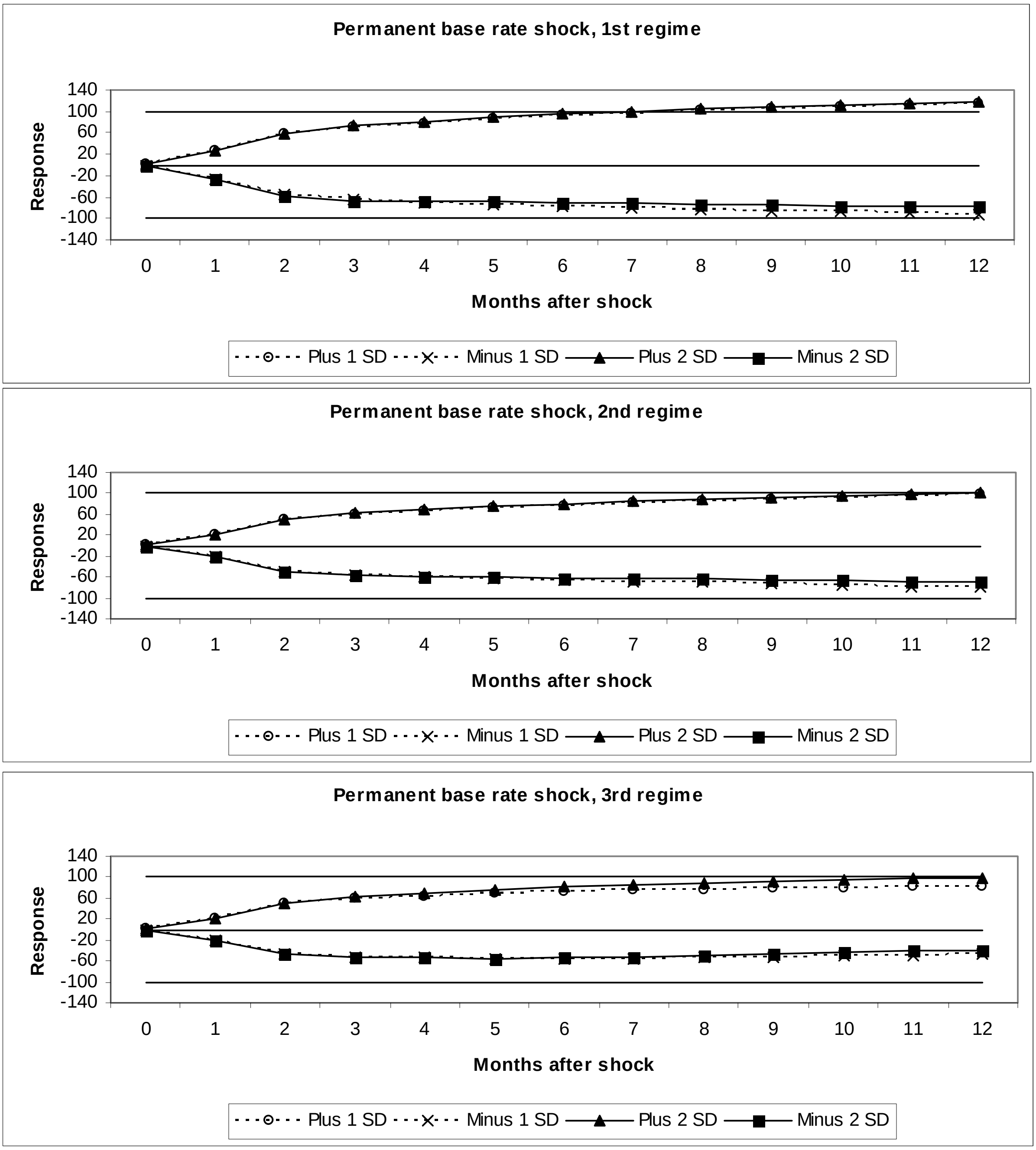


**Figure 5:** Dynamic Responses of the mortgage rate to Base Rate Shocks. Responses are given in percentage terms and are obtained through stochastic simulations of the two-regime threshold ECM using 10,000 replications. Regimes refer to cases where: 1st regime: $|\Delta libor_{t-1}| > 0.090$ and $|\Delta brate_{t-1}| > 0$; 2nd regime: $|\Delta libor_{t-1}| > 0.090$ and $|\Delta brate_{t-1}| \leq 0$; 3rd regime: $|\Delta libor_{t-1}| \leq 0.090$ and $|\Delta brate_{t-1}| \leq 0$. The regime where $|\Delta libor_{t-1}| \leq 0.090$ and $|\Delta brate_{t-1}| > 0$ is not considered due to an insufficient number of initial values.

**Table 1: Johansen Linear Cointegration Test Results**

| | $(brate, libor, mrate)$ | $(brate, libor)$ | $(libor, mrate)$ |
|---|---|---|---|
| Sample period January 1995 to January 2006 | | | |
| Trace | | | |
| $r = 0$ | 45.146** | 21.565** | 19.399* |
| $r \leq 1$ | 16.933 | 3.353 | 2.027 |
| $r \leq 2$ | 1.346 | - | - |
| Max | | | |
| $r = 0$ | 28.021** | 18.212** | 17.356** |
| $r \leq 1$ | 15.588* | 3.353 | 2.027 |
| $r \leq 2$ | 1.346 | - | - |
| Sample period January 1995 to August 2008 | | | |
| Trace test | | | |
| $r = 0$ | 40.142** | 14.670 | 20.299** |
| $r \leq 1$ | 8.689 | 3.870 | 2.438 |
| $r \leq 2$ | 1.878 | - | - |
| Max test | | | |
| $r = 0$ | 31.452** | 10.800 | 17.861** |
| $r \leq 1$ | 6.881 | 3.870 | 2.438 |
| $r \leq 2$ | 1.878 | - | - |

Notes: The results employ the assumption of no deterministic trend, with intercepts restricted to the cointegration space. The. lag orders of the vector autoregressive (VAR) models are determined using the Schwarz criterion to a maximum of 12, which leads to 2 lags for the three-variable VAR for ($brate_t$, $libor_t$, $mrate_t$) and 3 lags for each bivariate VAR. ** and * denote significance at 5% and 10% levels, respectively.

**Table 2: Threshold Cointegration Test Results**

| Nonlinear Driver | Pass-Through to LIBOR ($i$ = 1) | | | | Pass-Through to Mortgage Rate ($i$ = 2) | | | |
|---|---|---|---|---|---|---|---|---|
| | To January 2006 | | To August 2008 | | To January 2006 | | To August 2008 | |
| | Cointegration | Asymmetry | Cointegration | Asymmetry | Cointegration | Asymmetry | Cointegration | Asymmetry |
| $\hat{u}_{i,t-1}$ | 13.623 [0.002]** | 4.263 [0.128] | 11.493 [0.069]* | 4.919 [0.092]* | 14.511 [0.013]** | 2.549 [0.343] | 20.079 [0.001]** | 2.967 [0.268] |
| $\Delta\hat{u}_{i,t-1}$ | 16.427 [0.018]** | 7.346 [0.035]** | 11.024 [0.172] | NA | 17.977 [0.008]** | 5.419 [0.118] | 22.466 [0.001]** | 4.646 [0.189] |
| $\Delta libor_{t-1}$ | 27.238 [0.000]** | 19.231 [0.000]** | 16.313 [0.021]** | 10.095 [0.009]** | 19.998 [0.004]** | 7.665 [0.033]** | 21.598 [0.001]** | 3.667 [0.298] |
| $\Delta brate_{t-1}$ | 12.017 [0.032]** | 2.498 [0.138] | 15.078 [0.004]** | 9.733 [0.052]* | 18.133 [0.002]** | 5.593 [0.019]** | 17.923 [0.000]** | 6.649 [0.011]** |
| $\Delta mrate_{t-1}$ | | | | | 18.019 [0.006]** | 5.465 [0.073]* | 22.365 [0.001]** | 4.532 [0.163] |
| $\lvert\hat{u}_{i,t-1}\rvert$ | 18.990 [0.006]** | 10.164 [0.028]** | 17.811 [0.018]** | 11.704 [0.017]** | 15.202 [0.022]** | 3.311 [0.594] | 20.078 [0.003]** | 2.965 [0.591] |
| $\lvert\Delta\hat{u}_{i,t-1}\rvert$ | 25.173 [0.000]** | 16.961 [0.000]** | 11.860 [0.120] | NA | 15.934 [0.016]** | 3.148 [0.373] | 21.421 [0.001]** | 3.466 [0.373] |
| $\lvert\Delta libor_{t-1}\rvert$ | 22.476 [0.001]** | 13.995 [0.002]** | 17.919 [0.004]** | 11.819 [0.005]** | 23.039 [0.000]** | 11.043 [0.006]** | 25.605 [0.000]** | 8.185 [0.028]** |
| $\lvert\Delta brate_{t-1}\rvert$ | 9.793 [0.007]** | 0.053 [0.818] | 8.468 [0.014]** | 1.672 [0.195] | 25.169 [0.000]** | 13.410 [0.000]** | 30.806 [0.000]** | 14.051 [0.000]** |
| $\lvert\Delta mrate_{t-1}\rvert$ | | | | | 20.074 [0.004]** | 7.750 [0.022]** | 24.577 [0.001]** | 7.026 [0.045]** |

Notes: The tests for (possibly nonlinear) cointegration and asymmetry are described in subsections 4.4 and 4.3, respectively. The required number of lagged changes to ensure *iid* residuals in (6) is two for all cases where the pass-through is to LIBOR ($i$ = 1) and zero for all cases of the pass-through to the mortgage rate ($i$ = 2). The values in brackets in the table are $p$-values. Except when $|\Delta brate_{t-1}|$ is considered as the threshold variable, the $p$-values are obtained using the bootstrap algorithms (with 50,000 replications) described in subsection 4.3. For the threshold variable $|\Delta brate_{t-1}|$, the threshold value is set to 0 and chi-square $p$-values are reported. NA indicates that asymmetry test is not reported due to lack of evidence for cointegration.

## Table 3: Estimated Models for Pass-Through to LIBOR

| Cointegrating relation (Dependent variable $libor_t$) | Estimated to January 2006 | | Estimated to August 2008 |
|---|---|---|---|
| | Linear | Threshold | Threshold |
| Constant | 0.021<br>[0.945] | 0.139<br>[0.496] | 0.424<br>[0.021] |
| $brate_t$ | 1.008<br>[0.000] | 0.990<br>[0.000] | 0.948<br>[0.000] |
| $D$0001 | -0.148<br>[0.088] | -0.080<br>[0.150] | -0.087<br>[0.133] |
| $D$0110 | 0.158<br>[0.158] | 0.132<br>[0.116] | |
| $D$0708 | | | 0.120<br>[0.284] |
| $D$0801 | | | 0.025<br>[0.867] |
| Complete pass-through | 0.027<br>[0.869] | 0.092<br>[0.763] | 3.395<br>[0.065] |
| Model statistics | | | |
| $\hat{\sigma}$ | 0.078 | 0.069 | 0.067 |
| SBC | -2.040 | -2.191 | -2.060 |
| AIC | -2.193 | -2.378 | -2.345 |

| Shortrun adjustment (Dependent variable $\Delta libor_t$) | Estimated to January 2006 | | Estimated to August 2008 |
|---|---|---|---|
| | Linear | Threshold | Threshold |
| $M_{1t} \times \hat{u}_{1,t-1}$ | -0.267<br>[0.000] | -0.513<br>[0.001] | -0.442<br>[0.001] |
| $(1 - M_{1t}) \times \hat{u}_{1,t-1}$ | | 0.044<br>[0.392] | 0.056<br>[0.281] |
| $\Delta brate_t$ | 0.841<br>[0.000] | 0.847<br>[0.000] | 0.844<br>[0.000] |
| $\Delta brate_{t-1}$ | 0.082<br>[0.095] | 0.174<br>[0.000] | 0.369<br>[0.000] |
| $\Delta brate_{t-2}$ | | | 0.216<br>[0.000] |
| $\Delta brate_t \times d0708_t$ | | | -0.269<br>[0.238] |
| $\Delta brate_{t-1} \times d0708_t$ | | | -0.708<br>[0.001] |
| $\Delta brate_{t-2} \times d0708_t$ | | | -0.252<br>[0.199] |
| $\Delta libor_{t-1}$ | | | -0.182<br>[0.000] |
| $\Delta libor_{t-2}$ | | | -0.162<br>[0.005] |
| $\tau$ | | 0.090 | 0.090 |

Notes: *D*yymm indicates a step dummy for month mm of year yy; the short-run adjustment equation includes the corresponding impulse dummy variables, *d*yymm (coefficients not shown). No linear model is presented for the longer period due to lack of evidence for linear cointegration over this period (Table 1). The threshold models use $|\Delta libor_{t-1}|$ as the nonlinear driver (see text). All values in brackets are *p*-values; for coefficients these test the null hypothesis of zero while the complete pass-through test is a Wald test of the null hypothesis that the coefficient on $brate_t$ in the long-run model is unity. SBC and AIC are normalised for sample size and the threshold parameter is included for the threshold ECM specifications.

## Table 4: Estimated Models for Pass-Through to Mortgage Rate

| | Estimated to January 2006 | | Estimated to August 2008 | |
|---|---|---|---|---|
| | Linear | Threshold | Linear | Threshold |
| Cointegrating relation (Dependent variable $mrate_t$) | | | | |
| Constant | 1.682<br>[0.000] | 2.227<br>[0.000] | 1.728<br>[0.000] | 2.239<br>[0.000] |
| $libor_t$ | 0.916<br>[0.000] | 0.829<br>[0.000] | 0.910<br>[0.000] | 0.828<br>[0.000] |
| $D9706$ | 0.324<br>[0.054] | 0.376<br>[0.005] | 0.313<br>[0.043] | 0.364<br>[0.000] |
| $D9812$ | 0.202<br>[0.618] | -0.102<br>[0.192] | 0.166<br>[0.657] | -0.093<br>[0.187] |
| $D9903$ | -0.027<br>[0.943] | 0.115<br>[0.128] | 0.058<br>[0.871] | 0.119<br>[0.083] |
| $D0612$ | | | -0.003<br>[0.984] | 0.134<br>[0.001] |
| $D0805$ | | | -0.061<br>[0.826] | -0.074<br>[0.865] |
| Complete pass-through | 1.918<br>[0.166] | 54.343<br>[0.000] | 2.552<br>[0.110] | 67.388<br>[0.000] |
| Model statistics | | | | |
| $\hat{\sigma}$ | 0.073 | 0.054 | 0.072 | 0.053 |
| SBC | -2.136 | -2.676 | -2.029 | -2.631 |
| AIC | -2.342 | -2.899 | -2.314 | -2.931 |

| | Estimated to January 2006 | | Estimated to August 2008 | |
|---|---|---|---|---|
| | Linear | Threshold | Linear | Threshold |
| Shortrun adjustment (Dependent variable $\Delta mrate_t$) | | | | |
| $M_{2t} \times \hat{u}_{2,t-1}$ | -0.164<br>[0.000] | -0.702<br>[0.000] | -0.174<br>[0.000] | -0.751<br>[0.000] |
| $(1 - M_{2t}) \times \hat{u}_{2,t-1}$ | | -0.073<br>[0.000] | | -0.072<br>[0.005] |
| $\Delta libor_t$ | 0.096<br>[0.042] | 0.025<br>[0.498] | 0.089<br>[0.052] | 0.019<br>[0.580] |
| $\Delta libor_{t-1}$ | 0.343<br>[0.000] | -0.006<br>[0.919] | 0.345<br>[0.000] | -0.020<br>[0.713] |
| $\Delta libor_{t-2}$ | 0.231<br>[0.000] | 0.160<br>[0.000] | 0.217<br>[0.000] | 0.148<br>[0.000] |
| $\Delta libor_t \times d0612_t$ | | | 0.194<br>[0.027] | 0.179<br>[0.008] |
| $\Delta libor_{t-1} \times d0612_t$ | | | -0.261<br>[0.005] | 0.096<br>[0.209] |
| $\Delta libor_{t-2} \times d0612_t$ | | | -0.031<br>[0.727] | -0.036<br>[0.583] |
| $\tau$ | | 0 | | 0 |

Notes: Notes: *D*yymm indicates a step dummy for month mm of year yy; the short-run adjustment equation includes the corresponding impulse dummy variables, *d*yymm (coefficients not shown). Both threshold models use $|\Delta brate_{t-1}|$ as the nonlinear driver (see text). All values in brackets are *p*-values; for coefficients these test the null hypothesis of zero while the complete pass-through test is a Wald test of the null hypothesis that the coefficient on $libor_t$ in the long-run model is unity. SBC and AIC are normalised for sample size and the threshold parameter is included for the threshold ECM specifications.