Are mortgage and capital markets integrated in the USA? A study of time-varying cointegration

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Abstract: Recent research has asserted that before 1980, mortgage and capital markets in the USA were not cointegrated as a result of credit rationing, but deregulation of financial institutions led to cointegrated mortgage and capital markets during the 1980s. In this paper, we investigate the causality and cointegration relationship between the mortgage and treasury yields. Our findings suggest that the mortgage and treasury yields are cointegrated before, but not cointegrated after, 1982. To circumvent the non-cointegration problem, this paper introduces a time-varying cointegration approach to the issue. It extends the cointegration theory in such a way that the cointegrating parameter is allowed to follow a smooth function of time. We apply this technique to study the cointegration relationship between the US mortgage and capital markets and find that they are cointegrated after 1982 with time-varying cointegrating parameter.

Keywords: mortgage markets; capital markets; time-varying cointegration.

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1 Introduction

A large change in capital market rates used to produce a slow and small change in mortgage rates. The financial press often reported the complaints made by home buyers and real estate brokers about the 'stickiness' of mortgage rates and the failure of mortgage rates to follow declines in treasury and corporate yields. The academic literature on mortgage markets has largely focused on the economic effects of credit rationing rather than empirical measurement of mortgage market segmentation.

During the late 1960s, there was a rise in interest rates following a rise in inflation rates. With Regulation Q ceilings, once the market interest rates rose above the deposit rate ceilings, depositors would withdraw their funds from the depository institutions to invest in securities that yield higher returns. As a result, the depository institutions would have fewer funds to issue mortgage loans. Yet they did not raise the mortgage rates sufficiently to reduce the number of loans required. Instead, they chose to ration the available supply of funds among the borrowers. Therefore the mortgage rates merely reflected the demand for and supply of mortgage credit. The mortgage rates failed to follow the changes in capital rates due to credit rationing.

However there was a major financial reform carried out during the early 1980s. The Depository Deregulation and Monetary Control Act of 1980, as well as the Garn-St Ger-main Depository Institutions Act of 1982, were enforced and the Regulation Q ceilings were abolished. In addition, the depository institutions were allowed to expand the range of loans and investments they could make. The belief was that the deregulation of depository institutions and exponential growth in the secondary mortgage market could result in a more cointegrated mortgage and capital market. Klamann (1961) and Guttentag and Beck (1970), in their studies of mortgage market Segmentation, discovered that the speed of adjustment of the mortgage market to changes in the capital market increased over the years. For the period 1947–1956, any changes in the capital market were followed by changes in the mortgage market only after a time lag over four quarters. For 1954–1960, the lag was between four to seven months. Roth (1988) tested the comovement of mortgage and capital markets and found that the degree of comovement increased significantly during the 1980s. Haney (1988) criticised both the methodology and database employed in the early studies on the relationship between mortgage and capital markets.

To overcome the weaknesses, he employed cross-spectral analysis. He found that the primary mortgage market was indeed segmented from the capital market, while the secondary mortgage market was more fully integrated with the capital market for the period 1973–1981. Furthermore, the speed of adjustment was only three weeks.

Some recent research has asserted that before 1980, mortgage and capital markets in the USA were not cointegrated as a result of credit rationing, but deregulation of financial institutions coupled with exponential growth in the secondary mortgage market, led to cointegrated mortgage and capital markets during the 1980s. For example, Devancy et al. (1992) tested the cointegration relationship between mortgage and capital markets for the period 1970–1987. Their data were the monthly yield on the ten-year treasury security and the monthly contract rate reported by the Federal Home Loan Board. Dividing the monthly data into prederegulation (1970–1982) and post-deregulation (1983–1987) testing periods, they discovered that the mortgage market was not cointegrated with the capital market for the period 1970–1982, but fully integrated for the period 1983–1987. They also found that the rate of adjustment in the mortgage market to changes in the capital market increased rapidly towards the late 1980s as compared to the early 1980s.

However, more recent studies have shown the existence of cointegration during the prederegulation as well as the post-deregulation period. For example, Rudolph and Griffith (1997) considered the period from April 1963 to December 1993 and found the mortgage rate to be cointegrated with the capital market rate for the subperiods 1963:04–1968:12, 1969:01–1979:12 and 1980:01–1993:12. They also found cointegration for the whole period. Allen et al. (1999) and Sa-Aadu et al. (2000), also found cointegration during the prederegulation as well as the post-deregulation periods. The difference in results for the prederegulation period can be explained by the difference in methodologies and yield series, as well as the time period used.

Nothaft and Freud (2003) concluded that the integration of the multifamily mortgage into the broader capital markets helped to create new sources of credit, as some traditional portfolio investors reduced their share of loan holdings and thus led to a more stable spread relative to a benchmark security and a lower level of mortgage rates.

In this paper, we first investigate the cointegration relationship between mortgage and capital markets by performing cointegration tests for the periods 1970–1982 and 1983–1987. We then extend our study to the periods 1963–1982 and 1983–2002. Our findings suggest that the mortgage and capital markets were cointegrated for the periods 1970–1982 and 1963–1982 but not cointegrated for the periods 1983–1987 and 1983–2004. We then introduce a new time-varying cointegrating parameter approach to study the time-varying cointegrating relationship between mortgage and capital markets for the period these markets are not cointegrated.

Roth (1988) used a partial adjustment model to examine the changing relationship between mortgage and capital markets. Devancy et al. (1992) further applied Sims' Granger Causality test of the mortgage and treasury yields for the period from June 1983 to August 1987 and concluded that causality is unidirectional from treasury to mortgage markets. The partial adjustment model recommended by Roth (1988) cannot determine the causality directly, while in the Sims test the choice of the lag structure is purely arbitrary.

In this paper, we employ the Minimum Final Prediction Error (MFPE) criterion for the Granger Causality test to identify the nature of the relationship between mortgage and capital markets. The approach is applied for the periods 1963–1982, 1983–2002 and for the whole period 1963–2002. The MFPE criterion differs from the Sims test in that its choice of the lag structure is well-defined and its detrending methods are based on recent statistical techniques. Our results from the causality tests suggest that causality runs unidirectional from capital to mortgage markets for 1963–1982, 1983–2002 and for

the whole period. The shortcoming of the partial adjustment model is that it does not take into consideration the effect of the number of independent variables in the equation involved.

2 Methodology

To examine the comovements between the endogenous variable y_t and the exogenous variable x_t , we first study their relationship by a simple regression

$$y_t = \alpha + \beta x_t + e_t \quad (1)$$

where e_t is the error term. The validity and reliability of the regression relationship require the examination of the trend characteristics of the variables and cointegration test as the presence of unit root processes in the variables in the spurious regression problem (Granger and Newbold, 1974). Cointegration tests consist of two steps. The first step is to examine the stationary properties of the variables being studied. If a series, say y_t , has a stationary, invertible and stochastic ARMA representation after differencing d times, it is said to be integrated of order d and denoted by $y_t = I(d)$. To test the null hypothesis $H_0: y_t = I(l)$ versus the alternative hypothesis $H_1: y_t = I(0)$, we apply the Dickey–Fuller (DF) (1979, 1981) and the Augmented Dickey–Fuller (ADF) unit root tests based on the following regression.

$$\nabla y_t = \beta_0 + \alpha_0 t + \alpha_1 y_{t-1} + \sum_i (\beta_i y_{t-i}) + \varepsilon_t \quad \text{for } i = 1, 2, \dots, p \quad (2)$$

where $\nabla y_t = y_t - y_{t-1}$ and p is chosen to achieve empirical white noise residuals. To test the null hypothesis of the presence of a unit root in y_t , it is equivalent to testing the hypothesis that $\alpha_1 = 0$ in Equation (2). If α_1 is estimated to be significantly less than zero, the null hypothesis of a unit root for the series y_t is rejected. The test statistic used is the usual t -ratio computed for α_1 , but the distribution of this test statistic is not the t -distribution under the assumption of the null hypothesis. Note that if $p = 0$, this test is known as a DF test; if $p > 0$, it is known as the ADF test. The DF test assumes that the residuals are independently and identically distributed. However, this assumption does not hold for most economic time series. The ADF test takes into account this problem of serial correlation by including lagged dependent variables to achieve white noise residuals.

In addition, it is necessary to test for the presence of deterministic trends to ensure that it does not affect the validity of cointegration tests. We should test for the hypothesis that $(\beta_0, \alpha_0, \alpha_1) = (\beta_0, 0, 0)$ (random walk with drift). The test statistic is the likelihood ratio statistic Φ_3 found in Dickey and Fuller (1981). The distribution of Φ_3 is not the usual F -distribution under the null. The decision rule is to reject the null if Φ_3 is larger than the critical value. If the hypothesis that $(\beta_0, \alpha_0, \alpha_1) = (\beta_0, 0, 0)$ is not rejected, then we may test for the hypothesis that $(\beta_0, \alpha_0, \alpha_1) = (0, 0, 0)$ (random walk without drift). The test statistic, given by Φ_2 is the usual F -statistic but it is not distributed under the standard F -distribution. The null hypothesis will be rejected if the test statistic exceeds the relevant critical value. If both the hypothesis of $\alpha_1 = 0$ and the hypothesis of $(\beta_0, \alpha_0, \alpha_1) = (\beta_0, 0, 0)$ or $(\beta_0, \alpha_0, \alpha_1) = (\beta_0, 0, 0)$ are accepted, then we may conclude that y_t is an integrated process of order 1.¹ However, if the hypotheses are rejected, then y_t is stationary or $I(0)$. Also, one could apply other stationary tests, for example, the PP test

developed by Phillips and Perron (1988) to detect the presence of a unit root. The PP test is nonparametric with respect to nuisance parameters and thereby is suitable for a very wide class of weakly dependent and possibly heterogeneously distributed data.

If both y_t and x_t are of the same order, say $I(d)$, with $d > 0$, we then estimate the cointegrating parameter in (1) by OLS regression. If the residuals are stationary, the series, y_t and x_t , are said to be cointegrated. Otherwise, y_t and x_t are not cointegrated. In the former case, Stock (1987) has shown that the regression in (1) would produce superconsistent and highly efficient estimates of the cointegrating parameter, this method has been proven to produce estimates which would approach the true parameter values more rapidly than the least squares estimator in normal situations.

Cointegration exists for variables means despite variables are individually non-stationary, a linear combination of two or more time series can be stationary and there is a long-run equilibrium relationship between these variables. If the error term in (1) is stationary while the regressors are individually trending, there may be some transitory correlation between the individual regressors and the error term. However, in the long run, the correlation must be zero because of the fact that trending variables must eventually diverge from stationary ones. Thus, the regression on the levels of the variables is meaningful and not spurious.

The three most common tests for stationarity of the estimated residuals are, namely, the Cointegrating Regression Durbin-Watson (CRDW) test, the Dickey-Fuller (CRDF) test and the Augmented Dickey-Fuller (CRADF) test. The CRDW test is based upon the Durbin-Watson statistic generated by the cointegrating regression. Under the null of non-cointegration, the CRDW should be close to zero. The null is rejected if the statistic exceeds the tabulated critical value. Engle and Granger (1987) suggested that the CRDW test might be used to obtain quick approximate result. However, the power of this test is sensitive to the dynamic structure of the equilibrium error. Thus, it is not a reliable test for cointegration. Moreover, the power of the CRDW test is greater than the DF type tests for the case where the alternative hypothesis is a simple stationary first-order autoregressive process. The more reliable tests are the CRDF and CRADF tests which are based upon the following OLS regression

$$\nabla e_t = \gamma e_{t-1} + \sum_i \gamma_i z_{t-i} + \varepsilon_t \quad \text{for } i = 1, 2, \dots, p \quad (3)$$

where e_t are residuals from cointegrating regression (1). Engle and Yoo (1987) suggested that the value of p could be selected based upon the Akaike Information Criterion (AIC). If $p = 0$, this test is referred to as the CRDF test and if $p > 0$, it is referred to as the CRADF test. The null hypothesis of non-cointegration is rejected if the t -ratio on γ is less than the relevant critical value (which is negative, since the alternative is negativity of γ). Engle and Granger have tabulated the critical values for CRDW, CRDF and CRADF tests for bivariate case for a sample of 100 observations while Engle and Yoo (1987) provided critical values for samples varying from 50 to 200 observations.

Once the cointegration relationship between the endogenous and exogenous variables has been decided, we can adopt the bivariate VAR model to test for Granger causality. The VAR model has the advantage of not having an underlying theory and does not need any assumption about the values of the exogenous variables. If the cointegration does not exist between the variables, following Granger et al. (2000), we employ

$$\nabla y_t = a_0 + \sum_{i=1}^n a_{1i} y_{t-i} + \sum_{i=1}^m a_{2i} x_{t-i} + u_{1t}$$

$$\nabla x_t = b_0 + \sum_{i=1}^n b_{1i} x_{t-i} + \sum_{i=1}^m b_{2i} y_{t-i} + u_{2t} \quad (4)$$

where n and m are the optimum lags, u_{1t} and u_{2t} are the error terms. We test the null hypothesis, $H_0 : a_{21} = a_{22} = \dots = a_{2m} = 0$ which implies that any of the exogenous variables x_{t-i} does not Granger cause the endogenous variable y_t . Similarly, we test whether the endogenous variables y_{t-i} Granger cause exogenous variable x_t as well.

If the series is cointegrated, there is a long-term or equilibrium, relationship among the variables being studied. Their dynamic structure can be exploited for further investigation. An Error-Correction Model (ECM) abstracts the short- and long-term information in the modelling process (Granger and Weiss, 1983). The ECM first used by Hendry et al. (1984) and later popularised by Hendry (1986) and Engle and Granger (1987) corrects for disequilibrium in the short run. Engle and Granger (1987) showed that cointegration is implied by the existence of an error correction representation of the indices involved. An important theorem, known as the Granger representation theorem, states that if two variables y_t and x_t are cointegrated, then their relationship can be expressed as ECM. An error correction term $e_{t-1} = y_{t-1} - \delta x_{t-1}$ is added to the equation to test the Granger causality such that:

$$\begin{aligned} \nabla y_t &= a_0 + a e_{t-1} + \sum_{i=1}^n a_{1i} y_{t-i} + \sum_{i=1}^m a_{2i} x_{t-i} + u_{1t} \\ \nabla x_t &= b_0 + b e_{t-1} + \sum_{i=1}^n b_{1i} x_{t-i} + \sum_{i=1}^m b_{2i} y_{t-i} + u_{2t} \end{aligned} \quad (5)$$

The existence of the cointegration implies causality among the set of variables as manifested by $|a| + |b| > 0$ where a and b denote speeds of adjustment (Engle and Granger, 1987). If we do not reject: $H_0 : a_{21} = a_{22} = \dots = a_{2m} = 0$ and $a = 0$, it means that any of the exogenous variables x_{t-i} does not Granger cause the endogenous variable y_t . Similarly, do not reject: $H_0 : b_{21} = b_{22} = \dots = b_{2m} = 0$ and $b = 0$ suggests that the endogenous variable y_t does not Granger cause the exogenous variable x_t (Granger et al., 2000). The Akaike's Information (AIC) is used to determine the optimum lag structures in the regressions (4) and (5), where n and m are lags in the left hand and right hand side variables respectively; and u_{1t} and u_{2t} are disturbance terms obeying the assumptions of the classical linear regression model.

To test the hypothesis $H_0 : a_{21} = a_{22} = \dots = a_{2m} = 0$, we find the sum of square of residuals for both the full regression, SSE_F and the restricted regression, SSE_R , in (4) and apply the F test:

$$F = \frac{(SSE_R - SSE_F) / m}{SSE_F / (N - m - n - 2)}$$

where N is the number of observations, n and m are defined in (4) or (5). When H_0 is true, F is distributed as $F(m, N - m - n - 2)$. So, reject the hypothesis H_0 at the α level of significance if $F > F(\alpha, m, N - m - n - 2)$. Accept the reduced model if H_0 is not rejected. Similarly, we can test for the hypothesis $H_0 : b_{21} = b_{22} = \dots = b_{2m} = 0$ and then make decision on the causality. We apply the usual simple t statistic to test $H_0 : a = 0$ and

$H_0: b = 0$. To examine the Granger Causality between treasury and mortgage yields, in this paper we apply the MFPE criterion introduced by Hsiao (1979, 1981):

$$\text{FPE}_z(J, I) = \frac{(N + J + I + 1) \sum \hat{u}_t^2}{(N - J - I - 1)}, \quad z = y \text{ or } z = x \quad (6)$$

where N is the number of observations, $u_t = u_{1t}$ or $u_t = u_{2t}$ in (4), I, J are the corresponding n and m in (4), depending on whether ∇y_t or ∇x_t are being studied.

To determine the minimum FPE_y , the first step is to run the regression equation of y_t in (4) with the lagged values of y only and exclude the x values. Thus, with $I = 0$ and starting with $J = 1$ and increasing the number of lags of y one by one, the FPE_y is calculated each time by using Equation (6). Let J^* be the optimum lag of y which gives the minimum FPE_y in this step. In the second step, with $J = J^*$, the regression Equation (4) is estimated starting with $I = 1$ and increasing the number of lags one by one. The FPE_y is calculated each time with $J = J^*$ and I being equal to the specified lag of x by using Equation (6). Let I^* be the optimum lag of x that yields the minimum FPE_y in this step. Thus, J^* and I^* are optimum lags of y and x , respectively, in Equation (4). The same procedure applied on x in (4) is used to determine the MFPE of x , FPE_x .

Now, we can test for causal relations between mortgage and capital markets by testing the following null hypotheses separately:

- 1 $c_i = 0$, for $i = 1, 2, \dots, I^*$ and
- 2 $c'_i = 0$, for $i = 1, 2, \dots, I^{**}$.

Hsiao's method of determining the MFPEs suggests a straightforward way of applying the Granger Causality tests. The null hypothesis (a) is accepted if $\text{FPE}_y(J^*, 0) < \text{FPE}_y(J^*, I^*)$ and rejected otherwise. Similarly, the hypothesis (b) is accepted if $\text{FPE}_x(J^{**}, 0) < \text{FPE}_x(J^{**}, I^{**})$ and rejected otherwise. There are four theoretically possible test results. Firstly, if both the null hypothesis (a) and (b) are accepted, it implies no causal relation exists between mortgage and capital markets. Secondly, if the null hypothesis (a) is accepted and the null hypothesis (b) is rejected, then causality runs unidirectional from mortgage to capital market. Thirdly, if the null hypothesis (a) is rejected and the null hypothesis (b) is accepted, then causality runs unidirectional from capital to mortgage market. Finally, if both the null hypothesis (a) and (b) are rejected, it is said that there exists a feedback causal relation between mortgage and capital markets.

3 Further developments in cointegration

Consider a pair of series y_t and x_t in which both are integrated of order 1, or $I(1)$. If there exists a linear combination of y_t and x_t such that

$$z_t = y_t - Az_t \quad (7)$$

is stationary or $I(0)$, then the series y_t and x_t are said to be cointegrated, with a constant cointegrating parameter A . The cointegrated model with a constant cointegrating parameter has been widely used to analyse the relationship between a pair of non-stationary time series. However, it is only appropriate to analyse the situation in

which the equilibrium relationship between series remains relatively stable for a long period of time. With changes in the environment of society and technology, the cointegrating parameter may slowly evolve with time. In this case, Equation (7) may be written as

$$z_t = y_t - A_t x_t \quad (8)$$

where A_t is a smooth deterministic function of time. A_t is called the Time-Varying cointegrating Parameter (TVP) and there could exist an A_t such that y_t and x_t are cointegrated if we release the constant restriction on A_t and let A_t to be a smooth function of time. A simple case is to set $A_t = y_t/x_t$. In this paper, we only consider the situation in which A_t is smooth and changes gradually over time.

Granger (1986) defined TVP $I(1)$ and $I(0)$ processes as follows: Suppose, we have

$$y_t = \rho(t)y_{t-1} + \varepsilon_t$$

where $|\rho(t)| < 1$ for all t . Let $f_t(w)$ denotes the time-varying spectrum for the above process. Thus, y_t is TVP $I(0)$ process if $f_t(w)$ is bounded and $f_t(w) > 0$ for all t and w . On the other hand, y_t is TVP $I(1)$ process if ∇y_t is TVP $I(0)$ process.

Let y_t denotes a TVP $I(d)$ vector process without deterministic components such that $\nabla^d y_t$ has finite variance. The Wold representation is given by

$$(1 - B)^d y_t = C_t(B)\varepsilon_t$$

where ε_t is an $N \times 1$ white noise vector with zero mean such that

$$\begin{aligned} E\{\varepsilon_t \varepsilon_s'\} &= \begin{cases} 0 & \text{if } t \neq s \\ \Omega_t & \text{if } t = s \end{cases} \\ C_t(0) &= I_N \{\text{unit matrix}\} \\ C_t(B) &= \sum_j C_{jt} B^j \text{ and} \\ \sum_j C_{jt} \Omega_t C_{jt}' &< \infty \end{aligned}$$

Let $\text{rank}[C_t(1)] = N-1$ for all values of t . The cointegrating rank is therefore 1. We will have $N \times 1$ vectors α_t and γ_t such that

$$\alpha_t'(t)C_t(1) = 0 \text{ and } C_t(1)\gamma_t(t) = 0$$

then

$$z_t = \alpha_t'(t)y_t$$

and the error correction representation becomes

$$A_t^*(B)(1 - B)y_t = -\gamma_t(t)z_{t-1} + d_t(B)\varepsilon_t$$

where $A_t^*(0) = I_N$, $A_t^*(1)$ has full rank and the roots of $|A_t^*(w)| = 0$ are outside the unit circle.

Canarella et al. (CPL, 1990) proposed a TVP cointegration approach to testing purchasing power parity. They proposed a cointegrating regression of the form

$$e(t) = b_0(t) + b_1 p(t) + u(t)$$

where $e(t)$ and $p(t)$ are the log of exchange rate and relative prices, respectively and are $I(1)$. Thereafter, they use a Kalman filter technique to generate an $I(1)$ variable $b_0(t)$ for the regression such that $u(t)$ is $I(0)$. Honohan (1993) commented on their approach that the failure of short time series to satisfy purchasing power parity cannot be brushed aside by CPL's findings. He also pointed out the fact that for any two $I(1)$ series, even two uncorrelated random walks, one can easily generate another $I(1)$ series with the property stated in CPL's TVP model. He set an example:

$$b_0(t) = p(t-1) - e(t-1)$$

in which if p and e were uncorrelated random walks

$$p(t) = p(t-1) + v(t)$$

and

$$b_0(t) = e(t-1) + w(t)$$

then this definition of b_0 would also be a random walk, with

$$e(t) - p(t) - b_0(t) = w(t) - v(t)$$

to be a stationary series.

4 Estimating and testing with time-varying parameter

In this paper, we propose to utilise a different TVP approach in which the TVP parameter follows a smooth function of time. In this approach, the procedure of testing for cointegration with TVP is similar to that of testing cointegration with constant cointegrating parameter with two steps involved. The first step is to examine the time series properties of the variables being studied by conducting unit root tests upon the variables using Fuller (1976) and Dickey and Fuller (1979, 1981) unit root testing procedures. The next step in the TVP cointegration test is to test whether z_t in Equation (7) is $I(0)$. If it fails to be $I(0)$, we proceed on to estimate A_t in Equation (8) and examine whether the corresponding z_t to follow $I(0)$. If it does, we say that the series y_t and x_t are TVP cointegrated.

To estimate the cointegrating variable A_t , we employ the Kalman filtering technique (Akaike, 1973, 1974; Caines and Rissanen, 1974; Mehra, 1974). The estimation is performed via three stages. In the first stage, we obtain the initial estimate of A_t by regressing the ratio variable w_t , which is equal to the ratio of y_t and x_t , on a time variable t such that

$$w(t) = f(t) + e_t$$

Once we obtain the estimate of the function $f(t)$, we substitute it into the following model

$$v_t = y_t - f(t)x_t$$

and perform a diagnostic check on the residuals from the above model to test whether v_t is $I(0)$. The series y_t and x_t are said to be cointegrated with time-varying cointegrating parameter $f(t)$ if we can find a smooth function $f(t)$ such that v_t is $I(0)$.

Finally we employ the Kalman filtering technique to estimate the TVP cointegrating regression given by Equation (9). We allow v_t to follow a general ARMA(p, q) process. The state-space representation of the variables is given by:

$$y_t = A_t x_t + H Z_t + a_t \quad (10)$$

Where $H = (1, 0, \dots, 0)$, Z_t is the $m \times 1$ state vector with its j th element

$$Z_{j,t} = E \left[y_{t+j-1} - A_t x_{t+j-1} \mid y_s, x_s; s \leq t \right] \quad \text{for } j = 1, \dots, m$$

in which $m = \max(p, q + 1)$. The state vector Z_t can be expressed by the state transition equation

$$Z_t = F Z_{t-1} + G \varepsilon_t \quad (11)$$

where

$$F = \begin{bmatrix} 0 & 1 & 0 & \cdots & 0 & 0 \\ 0 & 0 & 1 & \cdots & 0 & 0 \\ \vdots & \vdots & \vdots & \ddots & \vdots & \vdots \\ 0 & 0 & 0 & \cdots & 1 & 0 \\ \phi_m & \phi_{m-1} & \cdots & \cdots & \cdots & \phi_1 \end{bmatrix}$$

and

$$G = [1 \quad g_1 \quad \cdots \quad g_{m-1}]'$$

in which $g_0 = 1$ and

$$g_j = \sum_{k=0}^j \phi_k g_{j-k} - \theta_j \quad \text{for } j = 1, \dots, m-1$$

After obtaining the estimated $\hat{A}_t$, we will compute the residuals z_t such that

$$z_t = y_t + \hat{A}_t x_t$$

and test whether z_t is $I(0)$. If z_t is $I(0)$, we confirm that y_t and x_t are TVP cointegrated and we will choose the A_t with the least number of parameters.

5 Illustration

We illustrate our approach by examining the relationship between the capital market and the mortgage market. The data are the monthly yields of US 10-year Treasury bonds (T) reported in Moody's 10-year Treasury bill and the monthly home mortgage yields of new home purchase (US average) (M) reported in the Survey of Current Business from January 1963 to December 2002. We first study the unit root properties for the mortgage and treasury yields and their cointegration relationship for the four sub-periods: January 1970 to December 1982, January 1983 to December 1987, January 1963 to December 1982 and January 1983 to December 2002. The results are exhibited in Tables 1 and 2,

respectively. We note that we include the results of the first two subperiods as these subperiods are being studied by Devancy et al. (1992). We include these results so that readers could make comparison.

Table 1 Unit root test results for the mortgage and capital markets

Period	Variable	DF	ADF	ϕ_2	ϕ_1
1970–1982	M	–1.745	–1.875	3.483	2.038
	T	–2.642	–2.642	3.660	3.560
1983–1987	M	–2.069	–1.586	4.253	2.173
	T	–1.397	–1.904	0.259	1.037
1963–1982	M	–1.559	–2.261	3.679	1.850
	T	–2.568	–2.249	1.157	3.300
1983–2002	M	–1.703	–1.838	5.624*	2.598
	T	–1.334	–1.874	0.764	1.006

* $p < 0.05$, ** $p < 0.01$.

Note: T is the monthly yields of US 10-year Treasury bonds; M is the monthly home mortgage yields of new home purchase.

Table 2 Unit root test results for the mortgage and capital markets

Period	Model	R^2 (%)	CRDW	CRDF	CRADF
1970–1982	$M = 2.357 + 0.860 T$	92.78	0.698**	–5.686**	–3.838**
1983–1987	$M = 5.273 + 0.549 T$	66.23	0.153	–2.337	–2.337
1963–1982	$M = 2.060 + 0.886 T$	95.21	0.655**	–6.740**9	–4.655**
1983–2002	$M = 2.859 + 0.762 T$	79.10	0.139	–3.020	–3.020

* $p < 0.05$, ** $p < 0.01$.

Note: T is the monthly yields of US 10-year Treasury bonds; M is the monthly home mortgage yields of new home purchase.

For each series, we apply the DF unit root test by performing the OLS regression in Equation (2). Based upon the Ljung-Box-Pierce Q-statistic, the residuals from the DF regressions are tested for white noise by selecting the shortest autoregressive lag lengths. To test the null hypothesis of a unit root in time-series is equivalent to test the hypothesis that $\alpha_1 = 0$. Table 1 summarises the results of unit root test of mortgage and treasury yields for the prederegulation periods 1970–1982 and 1963–1982 and for the post-deregulation periods 1983–1987 and 1983–2002. As shown in Table 1, the t-ratio, ϕ_2 and ϕ_3 statistics for both mortgage and treasury yields are not significant at the 0.01 level of significance for all periods except ϕ_2 for the period 1983–2002 which is not significant at the 0.05 level. This enables us not to reject the null hypothesis of a unit root for both M and T for any of these four periods; indicating that the mortgage and treasury series are $I(1)$ in levels for any of these four periods.²

After revealing that both mortgage and treasury yields are $I(1)$, we proceed to test for their cointegration relationship by estimating the following regression:

$$M_t = \alpha + \beta T_t + z_t$$

To do this, we conduct unit root tests upon the residuals from the above cointegration regression to test for stationarity of the residuals z_t . This is done by performing the regression in Equation (3). A diagnostic check is then performed to the residuals by testing the hypothesis of white noise through 24 lags using the Ljung-Box-Pierce Q-statistic (Box et al., 1994).

Table 2 shows the results of the cointegration test for all periods. For the period 1970–1982, the CRDW statistic of 0.698 is significantly different from zero at the 1% level, leading us to reject the null hypothesis of non-cointegration and conclude that the mortgage and treasury yields for the period 1970–1982 are cointegrated. The CRDF statistic of –5.686 and the CRADF statistic of –4.469 are also significant at the 1% level. Hence, we reject the null hypothesis of non-cointegration at the 1% level of significance based upon the CRDW, CRDF and CRADF statistics. Similarly, we reject the null hypothesis of non-cointegration at the 1% level of significance for the period 1963–1982. Therefore, we conclude that the mortgage and treasury series for the periods 1970–1982 and 1963–1982 constitute a cointegrating vector. This conclusion differs from the existing literature on credit rationing in mortgage market during the 1970s.

Next, we proceed to investigate the cointegration relationship for the periods 1983–1987 and 1983–2002. From Table 2, the CRDW statistic of the cointegrating regression is insignificant for both 1983–1987 and 1983–2002. This suggests that M and T may not be cointegrated in these two periods. For the CRADF test, we find that the lags on the differenced residuals are not significant for both two periods. Thus, we only need to compare the CRDF statistic with the DF critical values in Engle and Yoo (1987). We do not reject the null hypothesis of non-cointegration at the 5% levels of significance. Therefore, we conclude that the mortgage and treasury series for the periods 1983–1987 and 1983–2002 are not cointegrated.

This study investigates the long-run economic relationship between mortgage and capital markets by applying time series tests for cointegration. Cointegration in mortgage and capital markets implies the presence of a long-run equilibrium relationship for the period before deregulation, revealing that the M and T series still move in a similar direction over the long run for the period before deregulation although changes in mortgage market always lag behind changes in capital market in the short run. Credit rationing did not result in non-cointegrated mortgage and capital markets. It led to a slower speed of adjustment of mortgage yield to changes in treasury yield. Thus, the mortgage and capital markets were cointegrated for the period 1970–1982 and the period 1963–1982.

However, our findings also suggest that the mortgage and capital markets are not cointegrated after deregulation. One possible explanation for this is that the deregulation of financial institutions has significantly altered the environment of mortgage market. Failure to reject the null hypothesis of non-cointegration suggests that there could be some other variables which are not captured in the cointegrating regressions or the mortgage and capital markets could follow a TVP cointegration relationship.

We next apply the TVP cointegration technique describe in previous section to analyse the TVB cointegration relationship between mortgage and capital markets for the period January 1983 to December 2002. Some research suggested that deregulation of financial institutions and development of secondary mortgage market in early 1980s have led to a more fully integrated mortgage and capital markets. However, as evidence by Table 1, the CRDW statistic of the constant parameter cointegrating regression of mortgage rates (M) on treasury rates (T) for post-deregulation period is insignificant,

indicating that M and T may not be cointegrated. This result is supported by the DF and ADF statistics which enable us not to reject the null hypothesis of non-cointegration even at the 1% level of significance. Therefore, we may conclude that the mortgage and treasury series for the period 1983–2002 are not cointegrated.

However, failure to find cointegration between mortgage and capital markets for the period January 1983 to December 2002 may suggest that cointegrating regression has omitted some relevant variables or the cointegrating parameter is not constant. In this paper, we study the latter. In the following analysis, we allow the cointegrating parameter to vary with time and test for existence of cointegration relationship in this context.

The time series plot of the ratio shows that it can be divided into the following four subperiods: 1/83 to 12/86, 1/87 to 12/89, 1/90 to 8/93 and 9/93 to 12/94. We define $t_1 = 1$ for 1/83, 2 for 2/83 and so on until $t_1 = 48$ for 12/86 and 0 otherwise; $t_2 = 1$ for 1/87 and so on until $t_2 = 36$ for 12/89 and 0 otherwise; $t_3 = 1$ for 1/90 and so on until $t_3 = 44$ for 8/93 and 0 otherwise; $t_4 = 1$ for 9/93 and so on until $t_4 = 36$ for 12/94 and 0 otherwise.

Thereafter, we fit a regression model on the ratio, w_t and the results are shown in the following:

$$\begin{aligned}\hat{w}_t = & 1.2105(0.0119) - 0.02346t_1(0.0013) - 0.02807t_2(0.0020) \\ & - 0.00778t_3(0.0016) - 0.01666t_4(0.0017) + 0.000583t_1^2(0.00003) \\ & + 0.000846t_2^2(0.000061) + 0.000181t_3^2(0.000038)\end{aligned}$$

with the adjusted $R^2 = 80.30\%$ and $s = 0.0485$. Those in parentheses are the standard errors of the corresponding estimates. The parameter estimates are significantly different from zero. This model provides a good estimate for the ratio, with adjusted R^2 of 0.8030. Now, we perform TVP cointegrating regression by regressing M against T and obtain the following results:

$$\begin{aligned}\frac{\hat{M}}{T} = & 1.2076(0.0128) - 0.0245t_1(0.0014) - 0.0284t_2(0.0021) \\ & - 0.0079t_3(0.0018) - 0.0165t_4(0.0019) + 0.000613t_1^2(0.00003) \\ & + 0.000857t_2^2(0.000066) + 0.000188t_3^2(0.000048)\end{aligned}$$

with adjusted $R^2 = 99.80\%$. This provides an initial estimate for A_t . Both DF test (CRDF = -5.665) and ADF test (CRADF = -6.227) on the residuals lead us not to reject that $M_t - \hat{f}(t)T_t$ is from $I(0)$ at the 0.01 level of significance, suggesting that the series are likely to follow a TVP cointegrated model. Applying the Kalman filter technique, we get the following model for M_t and T_t :

$$M_t - \hat{A}_t T_t = (1 + 0.81735B(0.07606) + 0.50216B^2(0.07837))\varepsilon_t$$

where

$$\begin{aligned}A_t = & 1.2239(0.0202) - 0.0242t_1(0.0022) - 0.0298t_2(0.0034) \\ & - 0.0088t_3(0.0029) - 0.0173t_4(0.0029) + 0.000569t_1^2(0.00005) \\ & + 0.000877t_2^2(0.000103) + 0.000190t_3^2(0.000075)\end{aligned}$$

with AIC = 94.6327 and $Q(24) = 25.05$ with p -value = 0.295.

This confirms that the residuals from the above model are white noise.

Finally, we test whether $M_t - \hat{A}_t T_t$ is $I(0)$ where $\hat{A}_t$ is obtained from the above model. We perform the DF test and we get CRDF = -4.919 and CRADF = -5.432, both tests lead us not to reject that $M_t - \hat{A}_t T_t$ is from $I(0)$ at the 0.01 level of significance, confirming that the mortgage and capital markets follow a TVP cointegration relationship.

Cointegration in mortgage and capital markets means that although many developments can cause permanent changes in the market conditions, there is some long-run relation tying the two together. To complete the analysis, we further test for the Granger Causality between mortgage and treasury series based on Hsiao's method (Hsiao, 1979, 1981) of determining the optimum lag length of the regressor to identify the nature of the relationship.

Roth (1988) used a partial adjustment model to examine the changing relationship between mortgage and capital markets. Devancy et al. (1992) further applied Sims test Granger Causality of the mortgage and treasury yields for the period from June 1983 to August 1987, and concluded that causality is unidirectional from treasury to mortgage markets. The partial adjustment model recommended by Roth (1988) cannot determine the causality directly while in the Sims test, the choice of the lag structure is purely arbitrary.

In this paper, we employ Hsiao's the MFPE criterion for the Granger Causality to identify the nature of the relationship between mortgage and capital markets. The approach is applied for the periods 1963–1982, 1983–2002 and for the whole period 1963–2002. The MFPE criterion differs from the Sims tests in the way that its choice of the lag structure is well-defined and detrending method which are based on recent statistical techniques. The shortcoming of the partial adjustment model is that it does not take in the consideration of the effect of the number of independent variables in the equation property.

Once the presence of stochastic trends is found in the data, the Granger Causality tests must be done on the basis of the detrended values, which are stationary in nature. We apply the MFPE criterion introduced by Hsiao (1979, 1981) as stated in (6) to study the causal relations between mortgage and capital markets. In the regression analysis for causality tests, the variables are detrended to make them stationary by taking the first differences of the mortgage and treasury yields. Then the FPEs are calculated for regression equations in (4). Experiments have been made with 6 lags for both variables in all equation within which the minimum FPEs have been found. The results are reported in Table 3A for the testing period 1963–1982, in Table 3B for the testing period 1983–2002, in Table 3C for the whole period.

Table 3A Final prediction errors for the period Jan 1963 to Dec 1982

$FPE_y(J,0)$		$FPE_y(J,1)$		$FPE_x(J,0)$		$FPE_x(J,1)$	
(1,0)	0.03689*	(1,1)	0.03306	(1,0)	0.19552	(2,1)	0.19290
(2,0)	0.03706	(1,2)	0.02895	(2,0)	0.19125*	(2,2)	0.19288*
(3,0)	0.03724	(1,3)	0.02457	(3,0)	0.19290	(2,3)	0.19451
(4,0)	0.03713	(1,4)	0.02334	(4,0)	0.19351	(2,4)	0.19601
(5,0)	0.03707	(1,5)	0.02349	(5,0)	0.19315	(2,5)	0.19646

(6,0)	0.03729	(1,6)	0.02323*	(6,0)	0.19481	(2,6)	0.19814
Table 3B Final prediction errors for the period Jan 1983 to Dec 2002							
$FPE_y(J,0)$		$FPE_y(J,1)$		$FPE_x(J,0)$		$FPE_x(J,0)$	
(1,0)	0.02887	(5,1)	0.02305	(1,0)	0.07764	(2,1)	0.07491
(2,0)	0.02631	(5,2)	0.02291	(2,0)	0.07457*	(2,2)	0.07483*
(3,0)	0.02643	(5,3)	0.02306	(3,0)	0.07473	(2,3)	0.07580
(4,0)	0.02645	(5,4)	0.02335	(4,0)	0.07576	(2,4)	0.07685
(5,0)	0.02591*	(5,5)	0.02359	(5,0)	0.07681	(2,5)	0.07708
(6,0)	0.02620	(5,6)	0.02263*	(6,0)	0.07770	(2,6)	0.07679

Table 3C Final prediction errors for the period Jan 1963 to Dec 2002

$FPE_y(J,0)$		$FPE_y(J,1)$		$FPE_x(J,0)$		$FPE_x(J,0)$	
(1,0)	0.03473	(5,1)	0.02927	(1,0)	0.15370	(2,1)	0.15160*
(2,0)	0.03372	(5,2)	0.02609	(2,0)	0.15106*	(2,2)	0.15208
(3,0)	0.03390	(5,3)	0.02436	(3,0)	0.15176	(2,3)	0.15289
(4,0)	0.03391	(5,4)	0.02405	(4,0)	0.15237	(2,4)	0.15357
(5,0)	0.03308*	(5,5)	0.02394	(5,0)	0.15159	(2,5)	0.15430
(6,0)	0.03322	(5,6)	0.02389*	(6,0)	0.15233	(2,6)	0.15511

For the prederegulation period 1963–1982, y_t in Equation (4) has only one lag of y_t to yield the minimum FPE_y , which declines with the introduction of 6 lags of x_t ; x_t in Equation (4) has two lags of x for the minimum FPE_x , which increases with the inclusion of two lags of y . Thus, causality runs unidirectional from capital to mortgage markets for the period 1963–1982. For the post-deregulation period 1983–2002, y_t in Equation (4) has five lags of y to yield the minimum FPE_y , which is reduced by the inclusion of six lags of x ; x_t in Equation (4) has two lags of x for the minimum FPE_x , which increases with the addition of two lags of y . The results clearly show that causality also runs unidirectional from capital to mortgage markets for the period 1983–2002.

To confirm the results, we apply the Minimum FPE approach to the whole period from 1963 to 2002 and the results are in Table 3C which also show that causality also runs unidirectional from capital to mortgage markets. The results from Tables 3A–3C support the hypothesis that mortgage market participants with information on treasury yields, T , can improve forecasts changes in the mortgage rate, M , in comparison with forecasts that use information on M alone but the reverse is not true. The results which show the causality from long to short rate are consistent with the expectations theory and efficient markets claimed by MacDonald and Speight (1988) and support the ‘liquidity premium’ suggested by Devancy et al. (1992).

6 Concluding comments

In this paper, we report that causality runs unidirectional from capital to mortgage markets for both prederegulation and post-deregulation periods. This suggests that

mortgage market participants with information on treasury yields can improve forecasts changes in the mortgage rate, in comparison with forecasts that use information on the mortgage market alone, through the reverse is not true. However, the cointegration tests of mortgage and treasury rates allow the null hypothesis of non-cointegration to be rejected for the prederegulation period but not the post-deregulation period. These results suggest that the mortgage and capital markets were cointegrated before the 1980s, but not cointegrated during the 1980s. The findings lead us to reject the assertion that credit rationing led to non-cointegration, while deregulation resulted in cointegrated mortgage and capital markets.

We bear in mind that deregulation of financial institution in the early 1980s significantly changed the environment of the mortgage market. Failure to find cointegration between mortgage and capital markets for the post-deregulation period may indicate that the cointegrating regression omitted some relevant variables. Thus, future research could investigate the cointegration relationship in mortgage and capital markets by using a broader set of variables.

In addition, this paper studies the estimation testing of cointegration with time-varying parameters, which enables us to investigate the change of the cointegrating parameter over time. Engle and Yoo (1987) studied the behaviour of forecasts made from a cointegrated system with a constant cointegrating parameter. In our examples, since the cointegrating parameter follows a smooth function of time, the value does not change a lot in a one-step ahead forecast.

We can assume that the cointegrating parameter will follow the same smooth function in the near future. Since cointegration implies causality in at least one direction, we believe that we would obtain a better forecast of y_t using a cointegrated model with a time-varying parameter, compared to models based only on past observations of y_t alone. We would apply the Engle and Granger two-step estimation procedure by allowing the cointegrating residuals to enter the error correction model in place of the error correction term, and estimate the short run dynamic adjustment coefficients. Note that the error correction model involves a slowly evolving coefficient vector. The research could consider the estimation of error correction model with a time-varying parameter. However, the cointegrated model with a time-varying parameter requires the cointegrating parameter to be a smooth function of time. Further developments in this field should consider the possibility of relaxation of this restriction. One could refer to Schultz (1973) and Luo and Wahba (1997) and apply spline analysis to relax the restriction. An alternative would apply a Bayesian approach (Matsumura et al., 1990; Wong and Bian, 2000) or other advanced time series approaches (Tiku et al., 2000; Wong and Miller, 1990; Wong and Bian, 2005) to improve the estimation. One could also apply the stochastic dominance approach (Broll et al., 2006; Fong et al., 2005; Wong and Li, 1999; Wong and Chan, forthcoming) to study the relationship between the US mortgage and capital markets.

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Notes

¹In this situation, we need to further test the null hypothesis $H_0 : y_t = I(2)$ versus the alternative hypothesis $H_1 : y_t = I(1)$. If $H_0 : y_t = I(2)$ is rejected and thus the series is confirmed to be $I(1)$ and will become stationary after differencing. Note that most series are integrated of order at most one.

²we have further tested the series are under the null hypothesis of being $I(2)$ versus the alternative hypothesis of being $I(1)$ and confirmed that both series are $I(1)$. The results are available on request.

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