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Treasury Note and Bank Bill Rates, the Risk Premium and Australian Monetary Policy

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Abstract

This paper examines a link in the Australian monetary transmission mechanism based on the risk structure of certain interest rates. The bank-accepted bill and Treasury note rates cointegrate, and formal tests indicate that the risk premium was stationary after, but nonstationary before, the end of 1990. Well-defined and stable error-correction mechanisms also exist since December 1990, whereas prior to that they were unstable. These changes probably indicate a reduction in uncertainty and instability associated with the conduct of monetary policy. The evidence also indicates that, since December 1990, the Reserve Bank has been able to influence the bill rate by targeting the note rate.

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I Introduction

It is generally recognised that deregulation of the Australian financial system during the 1980s undermined the role of monetary aggregates as useful intermediate targets (Milbourne, 1990). Since the floating of the Australian dollar in December 1983, the Reserve Bank of Australia (RBA) has used the official cash rate as an operating instrument for the conduct of monetary policy (Dotsey, 1991). In a deregulated environment, monetary shocks are transmitted to the real economy through changes in interest rates and asset prices. If economic activity responds to movements in long-term interest rates but monetary authorities use a short-term interest rate as an operating instrument for monetary policy, the term and risk structures of interest rates become important in the transmission process.

Recent research using time-series analysis has established some of the major links in the transmission process for Australia. Elliott and Bewley (1994) find a stable long-run relationship between the official overnight funds rate and the unofficial overnight funds rate. This implies, since the RBA can control the official rate, that the authorities can effectively target unofficial cash rates. However, this is a link between interest rates only at the shortest end of the maturity spectrum. For monetary policy to affect economic activity, it is necessary that the targeted cash rates have predictable influences on longer-term rates for both government and private-sector securities. Karfakis and Moschos (1995) examine the links between the official cash rate, the 13-week Treasury note rate and 2-, 5- and 10-year Treasury bond rates by examining the respective yield curves and their implications for the expectation theory of the term structure of interest rates.[2] They find that the spread between each of the long-term

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[1] The authors would like to thank two anonymous referees and the Editor of the Economic Record for comments which , we believe, have substantially improved the quality of this paper.

[2] The expectations theory ot the term structure has been tested for, *inter alia*, the USA, Belgium, Canada, Germany and the UK by Campbell and Shiller (1987) and MacDonald and Speight (1992).

rates and the 13-week Treasury note rate has predictive power for changes in the Treasury note rate, a finding which is consistent with the expectation theory. Further, their evidence suggests that the spread between the 13-week Treasury note rate and the official cash rate predicts movements in the cash rate, which is also consistent with the expectation theory. One implication of these findings is that, because the expectation theory linking short and long rates holds, the RBA can influence long-term bond rates by changing the cash rate. Further, the RBA can influence the Treasury note rate by intervening in the official cash market.

However, the above-mentioned research, with its emphasis on links between the official cash rate and rates on government paper of varying maturities, neglects the links between the rates on government paper and those on private paper, the rates at which the private sector can borrow. A recent paper by Lim and Martin (1994) completes some of the links in the chain by exploring the dynamic inter-relationships between short-term interest rates on both private and government paper using daily data over the sample period 1 July 1988 to 28 June 1991. They conclude that the official cash rate and the 11 am and 24-hour call rates lead, while 90-day and 180-day bank-accepted bill rates lag, in interest rate cycles. The 13-week and 26-week Treasury note rate and the 30-day bank bill appear to be approximately coincident. They also conclude that the coherence between the rates was much higher in the second of their two subperiods (January 1990 to mid-1991) than in the earlier subperiod. This they attribute to the fact that the RBA *announced* changes to the cash rate in the later, but not in the earlier, period.

This paper focuses on one link between private and government paper rates, in particular, on the risk spread between the 13-week Treasury note and 90-day bank-accepted bill rates over the period 1984-93. We explicitly test (1) whether there exist long-run comovements between the Treasury note and bank bill rates by means of cointegration techniques, (2) whether changes in the Treasury note rate predict movements in the bank bill rate using the error-correction (EC) methodology and, most importantly, (3) whether (and if so when) there have been substantial changes in the risk premium and EC dynamics over the sample period.

The raw data on the two interest rates and the relationship between them are illustrated in Figure 1 below. Although the two rates clearly move together, the relationship appears to have been much closer since mid- or late-1990. There also

seems to have been a dramatic reduction in the risk premium at about this time. Casual empiricism would suggest that the change took place a little later than Lim and Martin suggest. It may still reflect, however, the sort of announcement affect that they postulate. Alternatively, it may be attributed to the role of Bernie Fraser who was appointed as Governor of the Reserve Bank in September 1989. His early clear statements about the role of monetary policy[3], his emphasis on the official cash rate as the instrument and his publicised expectation that changes in cash rates should be reflected promptly in changes in other interest rates may have reduced the uncertainty surrounding official policy. By the time of Fraser's statements about the role of monetary policy, the RBA had stopped trying to target the exchange rate[4] and was clearly, by the end of the 1980s, targeting inflation. This promise of a more stable inflationary environment may also have helped to reduce uncertainty in financial markets. There is an obvious need to test formally for the existence and timing of any structural break in the relationship between the note and bill rates associated with a shift or shifts in the monetary policy regime. (We should like to call it the 'Bernie Fraser effect'[5] but clearly the three explanations listed above are neither easy to disentangle nor mutually exclusive). We also need to establish whether or not the closer relationship has continued beyond the period of monetary tightening examined by Lim and Martin.

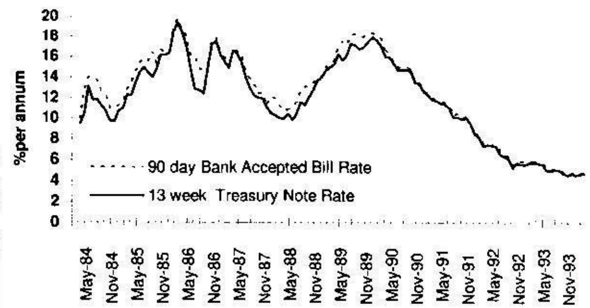

Figure 1 Treasury Note and Bank Accepted Bill Rates
(January 1984 to December 1993)

[3] Fraser (1990a) and (1990b).
[4] Evidence that the markets anticipated such intervention is contained in Karfakis and Phipps (1994).
[5] We are grateful to an anonymous referee who suggested we test for this effect and to the Editor who coined the name.

II Theoretical and Methodological Issues

The relationship between interest rates on bonds with the same maturity - the risk structure of interest rates - is explained by three factors: default risk and liquidity and income tax considerations (Mishkin, 1993). We wish to examine whether there is a stationary risk premium on bank-accepted bills over Treasury notes, and hence whether there is a stable long-run relationship between the two rates which is exploitable by the RBA. A stationary risk premium requires more than just cointegration between the two rates. Consider the relationship

$$BBR(t) = \alpha + \beta TNR(t) + u(t) \qquad (1)$$

where $BBR(t)$ and $TNR(t)$ are the logs of the bank bill and Treasury note rates[6] respectively and $u(t)$ is an error term. Cointegration between $BBR(t)$ and $TNR(t)$ is not a sufficient condition to ensure that the risk premium $RP(t) = BBR(t) - TNR(t)$ is a stationary process. It also requires that $\beta = 1$. To see this, rewrite the model (1) as follows:

$$RP(t) = BBR(t) - TNR(t) = \alpha + (\beta\text{-}1)TNR(t) + u(t) \qquad (2)$$

Since $u(t)$ is an I(0) process (we assume cointegration), if $\beta \neq 1$ equation (2) implies that the risk premium contains the same degree of persistence as $TNR(t)$. The only value of the cointegrating parameter β that implies $RP(t) \sim I(0)$ is $\beta = 1$.[7] Hence, to test for the stationarity of the risk premium we need also to test the restriction that $\beta = 1$. While stationarity of the risk premium could be tested directly, it is informative to proceed by way of cointegration because the value of any constant term in the

[6] We chose to work in logs because inspection of the data indicates that a proportional risk premium, rather than an absolute risk spread, better describes the relationship between the two rates.

[7] One reason, apart from systematic measurement error, for allowing $\beta \neq 1$ in equation (1) is that tax treatment of different interest incomes may vary. If τ and τ^* are the rates at which interest income from commercial bills and Treasury notes are respectively taxed, then $\beta = (1-\tau)/(1-\tau^*)$. Clearly if $\tau \neq \tau^*$, then $\beta \neq 1$.

cointegrating regression is of interest in its own right and testing for structural breaks in the cointegrating parameters is more readily achieved.

III Empirical Results

(i) Integration Analysis

The empirical analysis is carried out using monthly data on the 13-week Treasury note rate and the 90-day bank-accepted bill rate over the wholly-deregulated period January 1984 - December 1993. We have chosen to use monthly data because daily data is available for only a limited subperiod, and a longer sample period is probably more useful than more frequent data for exploring possible long-run relationships by cointegration.[8] Unit-root tests for both the levels and the first differences of the (log) series are reported in Table A.1 in the Appendix. The augmented Dickey-Fuller (ADF) test cannot reject the null hypothesis of a unit root in the levels of either of the interest rate series. In contrast, the hypothesis of a unit root in the first differences is rejected at the 5% significance level for both series. As far as the premium is concerned, the ADF test rejects the hypothesis of a unit root. Hence, time series modelling of the relationship between the two interest rate series requires first-difference transformation of the variables to induce stationarity

(ii) Cointegration Analysis.

Having established that the two interest rates are I(1) processes, we proceed to test for stationarity of the risk premium by estimating the cointegrating equation (1) and testing for cointegration and for the restriction that $\beta = 1$. In the presence of endogenous regressors and residual serial correlation, OLS estimates of the cointegrating equation are biased and cannot be used to test restrictions on the parameters. Phillips and Hansen (1990) have developed an approach which corrects for this bias. They have also produced a modified Wald (MW) statistic to test linear restrictions on the parameters of the cointegrating equation. We employ these procedures in the rest of this section.[9]

[8] Hakkio and Rush (1991).

[9] A brief technical description of the Phillips-Hansen method is contained in Karfakis and Moschos (1995). We have been encouraged to use this procedure for two reasons. First, we started off using the

The Phillips-Hansen estimates of the cointegrating equation (1) along with cointegration tests and MW statistics to test parameter restrictions are set out in Table 1. We focus initially on the results for the whole of our sample period (January 1984 - December 1993) which are set out in panel A. The Phillips-Perron test on $\hat{u}$ rejects the existence of a unit root in the residual of the cointegrating equation, establishing that the Treasury note and bank bill rates cointegrate for the full sample period. Further, the MW test cannot reject the null hypothesis that the cointegrating parameter, β, is unity. This suggests that the risk premium was stationary. However, casual inspection of the data casts doubt on the stability of these estimates. The existence of a structural break in the relationship is supported by two pieces of supplementary evidence. The cointegrating equation (1) was reestimated for the whole sample period using OLS with almost identical parameter estimates. This equation failed CUSUM and CUSUMSQ tests, with the critical bounds at the 5% significance level for CUSUM being broken in late 1990. We checked this result by testing for a structural break at the end of 1990[10]. The equation failed both a Chow 'structural stability' test with the sample period broken at December 1990 ($\chi^2(36)$ = 50.93 with a p-value of 0.05) and a Chow 'predictive failure' test for the period January 1991 - December 1993 ($\chi^2(2)$ = 65.50 with a p-value of 0.000). The long-run relationship was also estimated simultaneously with the short-run dynamics using the approach suggested by Mizon and Hendry (1980) again with similar parameter estimates for the long-run relationship. This equation, which is presented in panel A of Table 2, also failed CUSUM and CUSUMSQ tests with the 5% critical bounds for CUSUM being broken in March 1991. It also failed a Chow test for parameter stability with the sample broken at December 1990 ($\chi^2(36)$ = 47.34 with a p-value of 0.000). In short, the cointegrating equation of Table 1 (panel A) appears to be neither stable nor a good predictor when the sample period is broken at December 1990.

Johansen and Juselius (1990) technique in Microfit 3.0 but this produced such obviously silly estimates for the second of our two subperiods, possibly because of extreme multicollinearity among the regressors, that we abandoned it. Second, the results produced by the Phillips-Hansen method were confirmed by the joint estimation of the long-run relationship with short-run dynamics using the EC specification of Mizon and Hendry (1980). See Section III (iii)

[10] December 1990 was chosen for the break as a compromise between the breaks indicated by CUSUM and CUSUMSQ tests for the OLS estimates of the cointegrating equation and the Mizon and Hendry estimates respectively.

Table 1 Phillips-Hansen Estimates of the Cointegrating Equation

$$BBR(t) = \alpha + \beta TNR(t) + u(t)$$

A. Whole Period (Jan. 1984 to Dec. 1993)

Estimates:

$$BBR(t) = \underset{(0.05)}{0.03} + \underset{(0.02)}{1.01}\,TNR(t) \qquad ;\ R^2 = 0.99$$

Note: Figures in parentheses are asymptotic standard errors
Five lags of autocovariances used to correct for serial correlation. Number of lags selected by means of Sims (1980) LR test

MW Tests:

(1) $H_0: \alpha = 0$ $\chi^2(1) = 0.34$
(2) $H_0: \beta = 1$ $\chi^2(1) = 0.30$

Phillips-Perron Unit Root Tests on $\hat{u}$:

$Z(t_{\tilde{\tau}}) = -6.82$: Status - cointegration

Note: The $Z(t_{\tilde{\tau}})$ is the Dickey-Fuller τ_τ test (ie with mean and trend in the univariate regression)

B. Subperiod (Jan. 1984 to Dec. 1990)

Estimates:

$$BBR(t) = \underset{(0.09)}{0.43} + \underset{(0.04)}{0.86}\,TNR(t) \qquad ;\ R^2 = 0.92$$

Note: Two lags of autocovariances selected by reference to Sims LR test.

MW Tests:

(1) $H_0: \alpha = 0$ $\chi^2(1) = 20.67$
(2) $H_0: \beta = 1$ $\chi^2(1) = 15.31$

Phillips-Perron Unit Root Tests on $\hat{u}$:

$Z(t_{\tilde{\tau}}) = -6.61$: Status - cointegration

C. Subperiod (Jan. 1991 to Dec. 1993)

Estimates:

$$BBR(t) = \underset{(0.02)}{0.06} + \underset{(0.02)}{0.98}\,TNR(t) \qquad ;\ R^2 = 0.99$$

Note: Six lags of autocovariances selected by reference to Sims LR test.

MW Tests:

(1) $H_0: \alpha = 0$ $\chi^2(1) = 7.83$
(2) $H_0: \beta = 1$ $\chi^2(1) = 2.34$

Phillips-Perron Unit Root Tests on $\hat{u}$:

$Z(t_{\tilde{\tau}}) = -5.67$: Status - cointegration

To examine the instability of the cointegrating equation further, we reestimated it for the two subperiods January 1984 - December 1990 and January 1991 - December 1993. The results, which are presented in panels B and C of Table 1, indicate a dramatic change in the relationship and consequently in the risk premium. For the earlier subperiod (panel B), the Phillips-Perron test on the residuals indicates that *TNR* and *BBR* are cointegrated but the MW test clearly rejects the hypothesis that the cointegrating parameter β is unity. Hence, the risk premium prior to the end of 1990 was non-stationary. However, for the later subperiod the Phillips-Perron test rejects the hypothesis of a unit root in the residuals so that, as in the earlier period, *TNR* and *BBR* cointegrate but additionally the MW test cannot reject the hypothesis that β is unity. Hence, the risk premium since the end of 1990 has been stationary.[11] This is indicative of a more stable and predictable link in the interest rate chain that constitutes the Australian monetary policy transmission mechanism. The estimates of α also indicate that the risk premium was significantly lower in the second period than the first.

(iii) Error Correction Analysis

Having established that the bill and note rates cointegrate, it is appropriate to examine the associated EC mechanisms which describe the short-run dynamics. We do this by estimating EC models for the whole sample period and for each of the subperiods. Because of collinearity among the regressors, the standard Engle and Granger EC model failed to produce sensible estimates for the second subperiod.[12] We decided instead to use an alternative EC model devised by Mizon and Hendry (1980) which allows joint estimation of the long-run equilibrium relationship and the short-run dynamics. While this was not their preferred EC specification (the limit distribution of β involves second-order bias effects), it was ranked the second best in simulations conducted by Phillips and Loretan (1991). Using this specification, we estimate equations of the form:

[11] Since this observation regarding the second subperiod is made on the basis of data for three years only (36 observations), it should be treated tentatively.
[12] This collinearity may also explain the failure of the Johansen and Juselius (1990) method of testing for, and estimating, cointegrating vectors.

$$BBR(t) = \alpha + \beta TNR(t) + \sum_{k=1}^{t-1} \lambda_k \Delta BBR(t-k) + \sum_{k=0}^{t-1} \theta_k \Delta TNR(t-k) + \varepsilon(t) \qquad (3)$$

The results of estimating equation (3) are presented in Table 2 below.

Table 2 — Estimates and Diagnostics for the EC Model (3)

A. Whole Period (Jan. 1984 to Dec. 1993)

$$\underset{(0.03)}{BBR(t) = 0.03} + \underset{(0.01)}{1.01TNR(t)} - \underset{(0.07)}{0.19\Delta TNR(t)} + \text{6 significant lagged values of } \Delta TNR \text{ and } \Delta BBR$$

$R^2 = 0.99$ SEE = 0.04

$\chi^2_{SC}(12) = 49.98$ [0.00] $\chi^2_{FF}(1) = 12.68$ [0.00] $\chi^2_{N}(2) = 46.70$ [0.00] $\chi^2_{H}(1) = 0.20$ [0.66]
$\chi^2_{PF}(36) = 47.34$ [0.09] $\chi^2_{PI}(15) = 58.23$ [0.00]

B. Subperiod (Jan. 1984 to Dec. 1990)

$$\underset{(0.06)}{BBR(t) = 0.43} + \underset{(0.02)}{0.86TNR(t)} - \underset{(0.08)}{0.21\Delta TNR(t)} + \text{6 lagged values of } \Delta TNR \text{ and } \Delta BBR$$

$R^2 = 0.95$ SEE = 0.04

$\chi^2_{SC}(12) = 15.33$ [0.22] $\chi^2_{FF}(1) = 3.10$ [0.08] $\chi^2_{N}(2) = 3.89$ [0.14] $\chi^2_{H}(1) = 4.00$ [0.045]

C. Subperiod (Jan. 1991 to Dec. 1993)

$$\underset{(0.01)}{BBR(t) = 0.06} + \underset{(0.006)}{0.98TNR(t)} - \underset{(0.04)}{0.16\Delta TNR(t)} + \text{0 significant lagged values of } \Delta TNR \text{ and } \Delta BBR$$

$R^2 = 0.999$ SEE = 0.01

$\chi^2_{SC}(12) = 15.91$ [0.20] $\chi^2_{FF}(1) = 0.26$ [0.61] $\chi^2_{N}(2) = 2.99$ [0.22] $\chi^2_{H}(1) = 0.00$ [0.97]

Notes:

1. Numbers in rounded brackets are standard errors. In panel B (January 1984 to December 1990), they are White's heteroscedasticity-consistent standard errors.
2. Numbers in square brackets are P-values.
3. The subscripts on χ^2 indicate tests for serial correlation (SC), inappropriate functional form (FF), non-normality (N), heteroscedacity (H), 'predictive failure' (PF) and 'parameter instability' (PI) respectively.
4. For the last two (Chow) tests the sample period was broken at December 1990

The first thing to notice is that the estimated long-run coefficients, α and β, for the whole sample period and both subperiods are significant, correctly signed and confirm the Phillips-Hansen estimates of the previous section. The results for the whole period (panel A), as well as displaying the instability referred to in the previous section, also suffer from serial correlation and non-normality of the residuals and from possible functional misspecification, problems which were not removed by extending the number of lags[13]. In marked contrast, the estimated EC equation for the second subperiod (panel C) passes CUSUM and CUSUMSQ tests and is free from problems associated with serial correlation, inappropriate functional form, non-normality of residuals and heteroscedasticity. It also fits the data extraordinarily well. The estimate of β (0.98) is not significantly different from unity, confirming the MW test of the previous section. The estimated coefficient of -0.15 on ΔTNR indicates that changes in the Treasury note rate cause changes in the bank bill rate through temporary changes in the risk premium. Thus, an increase in the Treasury note rate reduces the risk premium temporarily until the long-run relationship is restored.

IV Concluding Remarks

Cointegration and error-correction estimates of the relationship between the 90-day bank-accepted bill rate and the 13-week Treasury note rate indicate some dramatic changes in the risk premium over the sample period (January 1984 - December 1993). Joint estimates of the long-run relationship and short-run dynamics in the EC model suggested by Mizon and Hendry (1980) revealed structural instability towards the end of 1990. Consequently, we decided to split our sample into two subperiods (January 1984 - December 1990 and January 1991 - December 1993). Cointegrating equations were estimated for the two subperiods separately. Formal tests of the restriction necessary for a stationary risk premium indicate that it has been stationary only since 1990. Prior to that the risk premium appears to have contained a permanent component. The risk premium was also significantly lower in the later period. These results imply a more stable, less uncertain financial environment since the end of 1990. Major changes which might explain this reduction in uncertainty are almost all associated with the conduct of official monetary policy. They include:

[13] We tried up to ten lags.

cessation of RBA attempts to target the exchange rate; RBA statements that monetary policy should primarily target inflation with some subsidiary attempt to iron out business cycle fluctuations; a shift by the RBA to *announcing* changes in the 'official rate'; and RBA assertions that changes in the official rate should be reflected promptly in changes in other interest rates. Some, but not all, of these changes in the monetary policy regime were associated with Bernie Fraser's appointment to the Governorship of the RBA.

Since 1990, a stationary risk premium, combined with the fact that changes in the note rate have Granger-caused changes in the bill rate, imply that the RBA could have influenced the bill rate by targeting the note rate. Since we already have evidence that the RBA can influence the Treasury note rate by operating in the official cash market, our results imply that the Bank can also affect the bank bill rate by altering the official rate. Together with the results of Elliott and Bewley (1994) and Karfakis and Moschos (1995), these results indicate the existence of a post-1990 monetary transmission mechanism running from the official cash rate through the Treasury note rate to rates on bank bills and long-term bonds.

Appendix

Table A.1: Unit Root Tests

Variable	Statistic	Without trend		With trend	
TNR	DF	-0.48	(-2.89)	-1.98	(-3.45)
TNR	ADF(1)	-1.09	(-2.89)	-2.23	(-3.45)
TNR	ADF(2)	-0.91	(-2.89)	-1.71	(-3.45)
TNR	ADF(3)	-0.78	(-2.89)	-2.02	(-3.45)
BBR	DF	-0.27	(-2.89)	-2.21	(-3.45)
BBR	ADF(1)	-0.52	(-2.89)	-1.71	(-3.45)
BBR	ADF(2)	-0.66	(-2.89)	-1.66	(-3.45)
BBR	ADF(3)	-0.54	(-2.89)	-1.66	(-3.45)
RP	DF	-5.39	(-2.89)	-6.95	(-3.45)
RP	ADF(1)	-3.89	(-2.89)	-4.79	(-3.45)
RP	ADF(2)	-4.09	(-2.89)	-5.52	(-3.45)
RP	ADF(3)	-4.03	(-2.89)	-5.36	(-3.45)

Notes:
Sample for DF (84M2 to 93M12) covers 119 observations.
95% critical values are in brackets.

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