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The Information Content of the
Yield Curve in Australia*

by

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This paper examines the expectation theory of the term structure of interest rates in Australia by looking at the information content of the yield curve. Cointegration results provide evidence that the slope coefficient of the yield curve is unity. Bivariate vector autoregressive analysis (VAR) indicates that the spread between the long term and the short term rates is informative about changes in the short rate. In addition, the spread between the short term rate and the official cash rate has predictive power for changes in the cash rate. These findings imply that the Reserve Bank of Australia could influence the long term rate by intervening on the official cash rate. Finally, the efficient market restrictions were tested and accepted by the data.

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1. Introduction

The financial 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 (the cost of funds for interbank clearing) as an operating instrument for the conduct of monetary policy (Dotsey, 1987, 1991).

In a deregulated environment, monetary shocks are transmitted to the real economy through changes in interest rates and asset prices.¹ If economic activity responds to movements in the long term interest rate but the monetary authorities use the short term interest rate as an operating instrument of monetary policy, the term structure of interest rates is becoming important in the monetary transmission mechanism.

According to the expectation theory of the term structure of interest rates, which is an arbitrage equilibrium condition, the long term rate is determined by the current and the expected short term rates over the same time horizon as the long term rate. In addition, under the expectation theory, the money market determines the short term interest rates as a weighted average of the current and expected levels of the interbank rate over the the same horizon as that of the short rate. Consider, for instance, a corporation which has the options of buying a treasury bill or lending its funds overnight for three months (Goodfriend, 1988, 1990). To the extent that the monetary authorities affect the current and expected future interbank rates and the theory is valid, the long term rate can be influenced through the effect of monetary policy actions on the short term rate.

Recent developments in the theory of cointegration have been used to test

the expectation theory of the term structure. In a earlier work, Campbell and Shiller (1987) examined post-war term structure data for the US using a vector autoregressive (VAR) system which incorporated an error-correction model for the short rate.² They found evidence in favour of the weak implication of the expectation theory, namely that the spread Granger-causes changes in the short rate. However they could not accept the full set of efficient market restrictions implied by the theory. MacDonald and Speight (1992) have investigated the expectation theory in a VAR context using Belgian, Canadian, German, UK and US data. They found mixed evidence concerning the weak implication of the theory and also rejected the efficient market restrictions for all countries apart from the UK, for which the tests confirmed the results in MacDonald and Speight (1988). In a recent study Taylor (1992) tested alternative models of the yield curve using weekly data on UK interest rates. His evidence indicates that the expectations and the risk premium models are both rejected. On the other hand, the data supported a market segmentation model.

The present paper examines comovements between interest rates in Australia over the deregulated period 1984-1991³ using the cointegration approach developed by Phillips and Hansen (1990) and the relevant Wald test for the unity restriction of the yield curve. Next, it investigates the information content of the yield curve by testing (a) whether the spread between the short term interest rate and the official cash rate is informative in forecasting changes in the official cash rate, and (b) whether the spread between the long term and the short term interest rates has predictive power for future changes in the short rate as predicted by the expectation theory of the term structure. It then proceeds with testing the restrictions implied by the efficient markets view of the term structure.

2. The Expectation Theory of the Term Structure of Interest Rates

The expectation theory of the term structure of interest rates holds that the n period bond rate ($R_{n,t}$) is a weighted average of present and expected future short term interest rates (r_t). In the case of discount bonds, the theory is written as (Mishkin, 1990)

$$R_{n,t} = (1/n)E_t(r_t + \dots + r_{t+n-1}) \quad (1)$$

where E_t is the expectation at time t . Subtracting r_t from both sides of (1) and rearranging, we obtain

$$S_{n,t} = (1/n)E_t[(n-1)\Delta r_{t,1} + (n-2)\Delta r_{t,2} + \dots + \Delta r_{t,n-1}] \quad (2)$$

where $S_{n,t} = R_{n,t} - r_t$ denotes the slope of the yield curve. A positive (negative) yield curve predicts that the short rates are expected to rise (fall) in the future. Thus, the expectation is that future short rates, and hence the long rate, will be above (below) the current short rate.

In the case of coupon bonds, the expectation theory is written as (Shiller, 1979)

$$R_{n,t} = (1-\alpha)\sum_{i=0}^{n-1} \alpha^i E_t r_{t+i} \quad (3)$$

where $0 < \alpha = (1/(1+R^*)) < 1$ and R^* is the mean long term rate. According to equation (3), more recent values of r play a greater role than earlier values, and hence recent ones are weighted more heavily. Subtracting r_t from both sides of (3) and rearranging, we obtain

$$S_{n,t} = R_{n,t} - r_t = \sum_{i=1}^{\infty} a^i E_t \Delta r_{t+i} \quad (4)$$

Equation (4) shows that the spread is equal to the optimal forecast of future changes in the short rates.

3. Methodology

As Campbell and Shiller (1987) argue, if Δr_t is stationary then equation (4) implies that S_t is also stationary, or alternatively that R_n and r are cointegrated with a cointegrating vector [1 -1] (Engle and Granger, 1987). Thus, we can model the data generation process of Δr and S as a bivariate VAR system, in the context of which the information content of the yield curve can be examined.

Consider the following bivariate VAR model

$$\Delta z_t = \Psi + A(L)\Delta z_{t-1} + \varepsilon_t \quad (5)$$

where $z = [\Delta r \ S_n]'$ is a 2x1 vector of endogenous variables; $\Psi = [\Psi_1 \ \Psi_2]'$ is a fixed intercept vector; $A(L) = \{A_{\Delta r}(L) \ A_{S_n}(L)\}$ is a 2x2 polynomial matrix in the lag operator, with $A_{\Delta r}(L) = \{a(L) \ c(L)\}'$ and $A_{S_n}(L) = \{b(L) \ d(L)\}'$, $\varepsilon_t = [\varepsilon_{1t} \ \varepsilon_{2t}]'$ is a 2x1 vector of white noise errors with properties: $E(\varepsilon_t) = 0$, $E(\varepsilon_t \varepsilon_{1-s}') = \Sigma$, when $t=s$ and zero otherwise, with Σ denoting the variance-covariance matrix of residuals. The polynomials in the lag operators $a(L)$, $b(L)$, $c(L)$, and $d(L)$ are all of order p .

An investigation of the expectation theory of the term structure for the VAR system (5) is that S_t will Granger-cause the future path of Δr_t , if agents have information beyond the history of Δr_t . Moreover, following Campbell and Shiller

(1987), the full set of restrictions implied by the efficient market hypothesis are derived as $c_i = -a_i$ ($i=1,2,\dots,p$), $d_1 = (1/\alpha) - b_1$ and $d_i = -b_i$ ($i=2,\dots,p$). Using this set of restrictions to eliminate c_i and d_i from the VAR system and adding the equations for Δr_t and $S_{n,t}$ we obtain $S_{n,t} - (1/\alpha)S_{n,t+1} + \Delta r_t = \Psi_1 + \Psi_2 + u_{1t} + u_{2t}$. Thus under the efficient market restrictions the expression $\xi_t = S_{n,t} - (1/\alpha)S_{n,t+1} + \Delta r_t$ is unpredictable. Hence a test of this set of restrictions can be obtained using a Wald test of the orthogonality of ξ_t on the information set consisting of lagged Δr_t and $S_{n,t}$.

Estimation of the VAR system (5) requires investigation of the time series properties of the variables involved. To test whether a time series is integrated of order one $I(1)$ against the alternative of zero order integration $I(0)$, a number of unit root tests have been developed (Fuller, (1976), Dickey and Fuller (1981), Phillips (1987) and Perron (1988)). The hypothesis that $S_{n,t}$ is stationary can also be investigated in a bivariate context by testing whether or not R_n and r are cointegrated with a cointegrating vector [1 -1]. In the presence of endogenous regressors and residual serial correlation, the OLS estimates of the parameters of the cointegrating equation are biased and cannot be used to test parameter restrictions. Phillips and Hansen (1990) have developed an approach which corrects for this bias. They have also produced a modified Wald statistic to test linear restrictions on the parameters of cointegrating regressions.⁴

Consider the following model,

$$y_t = \lambda x_t + \omega_t \quad (6)$$

$$\Delta x_t = v_t \quad (7)$$

We assume that y_t and x_t are $I(1)$ processes and cointegrated, and $u_t = (\omega_t, v_t)'$ is stationary with zero mean. To test for $\lambda=1$, we should correct for the bias of λ arising from the presence of endogeneity and serial correlation. Phillips and Hansen method gives a fully modified (bias-adjusted) least squares estimate of λ as:

$$\lambda_t^* = \left[\sum_{t=1}^T y_t^* x_t - T \delta^* \left(\sum_{t=1}^T x_t^2 \right)^{-1} \right] \quad (8)$$

where $y_t^* = y_t - \Omega_{12} \Omega_{22}^{-1} \Delta x_t$, $\delta^* = [I - \Omega_{12} \Omega_{22}^{-1}] [\Lambda_{21} \Lambda_{22}]'$

$$\Omega \equiv [\Omega_{ij}] = T^{-1} \sum_{t=1}^T u_t u_t' + T^{-1} \sum_{\tau=1}^q \sum_{t=\tau+1}^T \eta(\tau, q) \Sigma(u_t u_{t-\tau}') + u_{t-\tau} u_t'$$

$$\Lambda \equiv [\Lambda_{ij}] = T^{-1} \sum_{t=1}^T u_t u_t' + T^{-1} \sum_{\tau=1}^q \sum_{t=\tau+1}^T \eta(\tau, q) \Sigma(u_t u_{t-\tau}') + u_{t-\tau} u_t'$$

and Ω is the Newey and West (1987) estimator of the variance-covariance matrix with q lags of autocovariances and weights η . The correction for endogeneity and serial correlation is provided respectively through the use of y_t^* and δ^* in expression (8). The modified wald (MW) test statistic for testing $\lambda=1$ is given by:

$$MW = (\lambda^* - 1)^2 \left[\Omega_{11.2} \otimes \left(\sum_{t=1}^T x_t^2 \right)^{-1} \right]^{-1} X^2(1) \quad (9)$$

where $\Omega_{11.2} = \Omega_{11} - \Omega_{12} \Omega_{22}^{-1} \Omega_{21}$.

4. Empirical Results

4.1 Integration and Cointegration Analysis

The empirical analysis is carried out using quarterly data on the 3-month treasury note rate (i), the official cash rate (o), the two (R_2), five (R_5) and ten (R_{10}) year treasury bonds rates over the period 1984-1991, taken from dx v2.1 data base.

The strategy recommended by Perron (1988) is used to test for unit roots. The test statistics of the single hypothesis of one unit root, and the joint hypotheses of a unit root and a zero trend as well as a unit root and a zero mean for both the levels and the first differences are reported in Table 1.

With regard to the interest rates levels, none of the tests can reject the null hypothesis of a unit root. On the contrary, the hypothesis of a unit root in the first differences of the interest rates is rejected at the 5% significance level for all the series. As far as the spreads are concerned, the tests reject the joint hypothesis for the level of S_{30} and the single hypothesis of one unit root (without trend and mean) for the levels of S_{2r} , S_{5r} and S_{10} . Overall, the univariate tests suggest that the interest rates are $I(1)$, whereas the spreads are $I(0)$ processes.

The comovements of the interest rates entering the definitions of the various spreads are also investigated by means of cointegration tests. In this respect, it should be noted that the existence of cointegration between y_t and x_t is not a sufficient condition to ensure that the spread $S_t = y_t - x_t$ is a stationary process. It also requires that $\lambda=1$. To see this, rewrite the model (6) as follows:

$$S_t = y_t - x_t = (\lambda - 1)x_t + u_t \quad (10)$$

Since u_t is an $I(0)$ process (we assume cointegration), if $\lambda \neq 1$ equation (10) implies

that the spread contains the same degree of persistence as x_t . Then the only value of the cointegrating parameter λ that will imply $S_t - I(0)$ is $\lambda=1$.⁵

The Phillips and Hansen tests for cointegration between the short term and the long term rates, and between the official cash and the short term rates, which appear in Table 2, reject the null hypothesis of noncointegration in all cases. In addition, the modified Wald test does not reject the hypothesis that $\lambda=1$, confirming the results obtained from the integration analysis. The nonrejection of the restriction that $\lambda=1$ implies that the interest rate spreads do not contain a permanent component.

Overall, the tests indicate that the interest rates are $I(1)$, but their spreads are $I(0)$ suggesting that the data generation mechanism of $(S_{t,v}, \Delta r_t)$ is modeled as a bivariate VAR system in the context of which the information content of the yield curve can be examined.

4.2. The VAR Analysis

The strategy adopted in specifying the number of lagged Δz terms in equation (5) was based on Sim's (1980) likelihood ratio test corrected for the degrees of freedom.⁶ A maximum number of four lags was considered in order to preserve a reasonable number of degrees of freedom. The selected lag length was also checked for the presence of serial correlation associated with the truncation of the lag structure. Table 3 reports the selected lags and residual misspecification tests. The diagnostic tests reject the presence of serial correlation, functional misspecification and heteroscedasticity. The normality test suggests that the error process represents a random sample from a normal distribution.

The results of Granger causality tests given in Table 4, suggests that the past history of the spread at two, five and ten year maturity has information

content for short term interest rate changes, as predicted by the expectation theory of the term structure. In addition, the spread between the treasury note rate and the official cash rate has predictive power for changes in the cash rate.

The existence of causality from Δo to $(i-o)$ seems to suggest that monetary policy actions can affect the interest rate spread at short horizons by intervening on the cash rate. The evidence that the interest rate spreads are Granger exogenous with respect to changes in the treasury note rate may indicate that they are only affected by inflationary expectations.

The results of the efficient markets (EM) view of the term structure are reported in Table 5. The evidence indicates that the EM restrictions of the VAR model are not rejected by the data.

The nonrejection of the expectation theory of the term structure seems to indicate that the Australian monetary authorities have not been strongly committed to maintaining short term interest rates targets in the deregulated period. Had this been the case, expected future short rates would have been equal to the actual short rate and variations in the spread would only reflect variations in the term-premium (Mankiw and Miron, 1986). A further implication of these findings is that the monetary authorities could influence the short term rate and, thereby, the long term rate by intervening on the current level of the cash rate, as well as on its expected path.

5. Concluding Remarks

This paper has examined the expectation theory of the term structure of interest rates by looking at the information content of the of the yield curve. Given the evidence on the presence of a unit root in the levels of the interest rates, the time series properties of the spreads were investigated by unit root and

cointegration tests. The results suggest that the the spreads are the same order of integration as the first differences of the interest rates. An interesting aspect of the VAR analysis is the evidence that the spread between the long term rates, and the 3-month treasury note rate has predictive power for changes in the short rate, which is consistent with the expectation theory. Furthermore, the spread between the 3-month note rate and the official cash rate has predictive power for movements in the cash rate. One implication of these findings is that the Reserve Bank of Australia could influence the long term rate by changing the cash rate. Finally, the EM restrictions are tested and accepted by the data.

NOTES

1. For a theoretical discussion of the monetary transmission mechanism, see Brunner and Meltzer (1990).

2. In a recent contribution, Campbell and Shiller (1991) using US post war-data showed that the spread between the long and the short rate forecasted rising short term interest rates, but a declining yield on the long term bond a finding which was inconsistent with the expectation theory of the term structure.

3. In our analysis, we take the start of the deregulated period to be 1984, so it coincides with the floating of the Australian dollar, the adoption of an interest rate operating instrument procedure (Dotsey, 1987, 1991), and the abolishment of all remaining controls on bank deposits interest rates. Fahrner and Rohling (1992) have also used 1984 as the start of the deregulated period.

4. A simple exposition of the Phillips and Hansen methodology is given in Evans and Lewis (1992).

5. Taking variances in equation (10) yields

$$\sigma_x^2 = (\lambda - 1)^2 \sigma_x^2 + \sigma_u^2$$

This equation shows that σ^2 is dominated by σ_u^2 , unless $\lambda = 1$.

6. The selection of the lag order was also based on Akaike's (1969) information criterion as implemented by Hsiao (1979). The results were similar with those obtained by the Sin's criterion.

Table 1
Phillips and Perron Unit Root Tests

STATISTICS									
V/ble	Φ_3	Φ_2	Φ_1	τ_t	τ_p	τ	τ_{nc}	τ_{pc}	τ_{op}
i	2.31	1.55	1.13	-2.0	-2.23	-0.3	1.99	-0.47	2.21
o	3.26	2.20	1.55	-2.28	-2.63	-0.14	2.31	-0.5	2.64
R ₂	2.08	1.40	0.72	-1.81	-2.07	-0.38	1.79	-0.65	2.04
R ₅	2.41	1.67	2.00	-2.10	-2.18	-0.55	2.07	-0.95	2.14
R ₁₀	3.69	2.67	3.59	-2.85	-2.33	-0.98	2.81	-1.69	2.29
S ₁₀	18.66*	12.59*	-	-6.43*					
S ₅	2.71	1.85	3.04	-2.33	-2.47	-2.24*	-1.2	0.15	-1.22
S ₂	2.48	1.69	1.87	-2.14	-2.34	-2.16*	-1.10	0.12	-1.11
S ₁₀	2.41	1.66	1.57	-2.04	-2.31	-2.13*	-1.10	0.13	-1.10
Δi	11.25*	7.51*	-	-5.05*					
Δo	12.07*	8.14*	-	-5.17*					
ΔR_2	10.77*	7.19*	-	-4.93*					
ΔR_5	13.50*	9.02*	-	-5.57*					
ΔR_{10}	15.05*	10.08*	-	-5.98*					

Notes: The calculated statistics are those reported in Dickey and Fuller (1981). The critical values at 5% with T=25 are: $\Phi_3=7.24$, $\Phi_2=5.68$, $\Phi_1=5.18$, $\tau_t=-3.60$, $\tau_p=-3.00$, $\tau=-1.95$, $\tau_{nc}=3.20$, $\tau_{pc}=2.85$, $\tau_{op}=2.61$. The critical value of τ at 5% with T=30 is equal to -1.90 (Blangiewicz and Charemza, 1990). The lag length of the Phillips and Perron test is set equal to 5 to ensure white noise residuals. $S_j=R_j-i$, $j=2,5,10$ and $S_{10} = i-o$.

* Indicates significance at 5%.

Table 2
Cointegration Analysis of Interest Rates: $y_t = \lambda x_t + u_t$

$y_t \times x_t$	λ	R ²	Wald-Test $H_0: \lambda=1$									
			Φ_3	Φ_2	Φ_1	τ_t	τ_p	τ	τ_{nc}	τ_{pc}	τ_{op}	
i,o	0.99	0.90	0.69	18.47*	12.46*	-	-6.40*					
R ₂ ,i	0.96	0.08	3.00	2.75	1.88	3.27	-2.37	-2.50	-2.53+	-0.26	0.12	
R ₅ ,i	0.95	-	2.29	2.49	1.71	1.97	-2.16	-2.35	-2.38+	-2.83	0.09	
R ₁₀ ,i	0.94	-	2.07	2.42	1.67	1.64	-2.05	-2.31	-2.35+	-0.34	0.10	

Notes: The modified Wald test is distributed as $\chi^2(1)$. A negative R² is not reported. * (+) indicates significance at 5% (10%).

The critical value of τ at 10% with T=30 is -2.35(Blangiewicz and Charemza, 1990).

The critical value of τ_t at 5% with T=50 is -3.29(Engle and Yoo, 1988).

It is worth noting that the Engle and Yoo model includes an intercept whereas ours does not, and hence the critical values are only indicative.

See also notes to Table 1.

Table 3
Residual Misspecification Tests in the VAR Model (5)

System	Lags	Dep. V/ble	R ²	SEE	SC Q(13)	FF $\chi^2(12)$	NO $\chi^2(1)$	HE $\chi^2(1)$	ARCH $\chi^2(12)$
S ₂ , Δi	1	S ₂	0.52	0.9280	18.61	17.73	0.18	0.41	0.00
		Δi	0.16	1.7041	15.84	13.74	2.09	0.92	0.00
S ₅ , Δi	1	S ₅	0.59	1.1325	18.96	18.19	0.21	0.19	0.15
		Δi	0.13	1.7375	14.32	13.40	0.99	0.51	0.00
S ₁₀ , Δi	1	S ₁₀	0.64	1.2924	20.28	16.79	0.29	0.02	0.37
		Δi	0.11	1.7568	13.18	13.45	1.23	0.26	0.00
S ₁₀ , Δo	4	S ₁₀	0.58	0.4583	18.16	20.73	0.01	0.81	1.53
		Δo	0.62	1.2695	6.92	18.52	2.33	0.24	0.05

Notes: SEE is the standard error of the equation. Q(13) is Ljung Box Q statistic for serial correlation at 13 degrees of freedom. The Q statistic of the system S₁₀, Δo is calculated at 12 degrees of freedom. SC is the Lagrange multiplier test of residual serial correlation. The Ramsey's RESET test of functional specification (FF) uses the square of the fitted values. The Jarque Bera normality test statistic (NO) is based on a test of skewness and kurtosis of residuals. The test of heteroscedasticity (HE) is based on the regression of squared residuals on squared fitted values. ARCH refers to the test statistic for autoregressive conditional heteroscedasticity. Numbers in parentheses are p-values.

Table 4
Granger Causality Tests

System	Dep. V/ble	F-test	p-value
$S_2, \Delta i$	S_2 Δi	$H_0^1: \Delta i \rightarrow S_2$ $H_0^2: S_2 \rightarrow \Delta i$	0.1817 0.0429*
$S_5, \Delta i$	S_5 Δi	$H_0^3: \Delta i \rightarrow S_5$ $H_0^4: S_5 \rightarrow \Delta i$	0.1886 0.074+
$S_{10}, \Delta i$	S_{10} Δi	$H_0^5: \Delta i \rightarrow S_{10}$ $H_0^6: S_{10} \rightarrow \Delta i$	0.1684 0.1026+
$S_{10}, \Delta o$	S_{10} Δo	$H_0^7: \Delta o \rightarrow S_{10}$ $H_0^8: S_{10} \rightarrow \Delta o$	0.04* 0.0068***

Notes: The F tests are calculated for the joint hypothesis that the coefficients of the lags of the independent variable are zero. The symbol \rightarrow means that the independent variable does not Granger cause the dependent variable. ***, *, + denote significance at 1%, 5%, and 10%, respectively.

Table 5
Tests of the Efficient Market Restrictions

System	F-test
$S_2, \Delta i$	F(2,24)=0.5111
$S_5, \Delta i$	F(2,24)=0.0338
$S_{10}, \Delta i$	F(2,24)=0.3827
$S_{10}, \Delta o$	F(8,15)=1.6731

Notes: The tests are obtained as the F-tests of the joint hypothesis that the lagged values of S_n and Δi are insignificant in the regressions of $\hat{\epsilon}_n = S_{n,t} - (1/\alpha)S_{n,t-1} + \Delta i_t$ on lagged S_n and Δi .

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