THE EXPECTATIONS THEORY OF THE TERM STRUCTURE AND SHORT-TERM INTEREST RATES IN AUSTRALIA

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ABSTRACT

The expectations theory of the term structure of interest rates states that the yields on financial assets of different maturities are related primarily by market expectations of future yields. The expectations theory has occupied a prominent place in both theoretical and policy debates at various times. However, extensive empirical work in the United States has soundly rejected the joint (testable) hypothesis of the expectations theory and zero (or a constant) risk premium.

The aim of this paper is to test this joint hypothesis against Australian short-term interest rate data for the period since the introduction of the tender system for the sale of Treasury notes in 1979. The sample period chosen is interesting for a number of reasons. First, the market has had a greater influence on the determination of interest rates since the introduction of the tender system. Second, there was a major structural change in the Australian financial system with the floating of the Australian Dollar in 1983. This provides scope to test for the impact of policy or institutional changes on the expectations theory.

The paper finds that the joint hypothesis of the expectations theory and zero (or a constant risk premium), cannot be rejected in this period. Furthermore, this joint hypothesis cannot be rejected in the period before or after the introduction of the float. There is however, evidence of parameter instability across the pre- and post-float periods.

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1. Introduction

The relationship between yields on financial assets of different maturities is a subject that has interested economists and policy makers for decades. The most commonly discussed explanation of this relationship is the expectations theory of the term structure. The "pure" expectations hypothesis (PEH) states that, in equilibrium, the expected returns from different investment strategies with the same horizon should be equal. For example, the expected return from investing in an n-period bond should equal the expected return from investing in a one-period bond over n successive periods. If this theory holds then long-term rates can be (approximately) expressed as a weighted average of current and expected short-term rates. More importantly, it suggests that if policy makers wish to alter long-term rates through their influence on short-term rates they must succeed in altering the market's expectations of future interest rates.

The expectations theory has recently been subject to extensive empirical scrutiny in the United States. On the basis of this empirical work, the expectations theory of the term structure has been rejected in various studies. Despite this consistent rejection, Shiller, Campbell and Schoenholtz (1983) note that the theory continually reappears in policy debates. They liken this "superficially appealing" theory to the indefatigable Tom of Tom and Jerry cartoons; "The villain, Tom the cat, may be buried under a ton of boulders, blasted through a brick wall (leaving a cat-shaped hole), or flattened by a steamroller. Yet seconds later he is up again plotting his evil deeds".[1]

In addition to these extensive empirical rejections of the expectations theory, the logical consistency of many of the economic propositions derived from it have been questioned. In an important paper Cox, Ingersoll and Ross (1981) [hereafter (CIR)] re-examine several propositions about the relation between long and short rates typically associated with the expectations theory.[2] CIR show that, when interest rates are random, these different propositions are

1. Shiller, Campbell and Shoenholtz (1983, p.175).

2. Campbell (1986), notes that these various propositions can be expressed as different definitions of the term premium.

inconsistent with each other and all but one are incompatible with any continuous time rational expectations equilibrium.[3] The single proposition which obtains in continuous-time rational expectations equilibrium is the proposition that the instantaneous expected rates of return on all bonds are equal to the prevailing spot interest rate. CIR call this the Local Expectations Hypothesis. CIR also show that the various propositions are inconsistent with each other in discrete time but are compatible with arbitrage pricing equilibrium. These arguments suggest that traditional tests of the expectations hypothesis may be incorrectly specified.

Campbell (1986), however, has defended the empirical applications of the expectations theory on two grounds. First, he argues that CIR consider a more restrictive form of the theory than is considered in the empirical literature. In particular, CIR's discussion is directed to the "pure" expectations theory which states that risk premia are zero whereas most empirical applications consider the less restrictive expectations hypothesis (EH) which allows for constant risk premia. Campbell shows that the propositions derived from this less restrictive theory are not necessarily incompatible with each other or with arbitrage pricing equilibrium. Furthermore, Campbell shows that any inconsistencies are of second order and may often be ignored in empirical studies.

Previous Australian studies on this topic were conducted at a time when yields on government securities were largely set by the authorities.[4] Since that time, a move to a more market-oriented system of interest rate determination has occurred. In particular, tender systems for the sale of Treasury notes and Government bonds were introduced in 1979 and 1982 respectively.

The purpose of this paper is to test the expectations theory using data on Australian short-term financial assets for the period since 1979. The sample period encompasses two significantly different policy regimes; namely, managed

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3. Because of the non-linearities in the term structure equation the alternative definitions of the term premium cannot simultaneously hold given Jensen's Inequality.

4. See Bloch (1974) and Jüttner, Madden and Tuckwell (1975). Bloch, for instance, found that a version of the expectations hypothesis held in the short term. However, the relationship between short-term yields and much longer-term yields was not consistent with the expectations hypothesis. However, both Bloch and Jüttner et.al. caution against drawing strong inferences from the results because of the dominance of the Reserve Bank in the market at that time.

and floating exchange rate regimes. Because of this, inferences can be drawn as to whether the relationships between interest rates are altered by policy regimes. A recent paper by Mankiw and Miron (1985) on U.S. data found that the empirical tests of the expectations hypothesis are sensitive to the policy regime of the time.

The findings of this paper are at variance with the U.S. results.[5] It is found that the data are consistent with the pure expectations hypothesis for the whole period. Furthermore, in most cases, the pure expectations hypothesis cannot be rejected in either the period of managed or floating exchange rates. In no case can the less restrictive expectations hypothesis (i.e., constant risk premium) be rejected. However, there is some evidence of parameter instability in the reported equations. Although cross country and cross time comparisons of empirical results are difficult, one explanation of these results may lie in the difference in monetary regimes and the pattern of financial flows in Australia and the United States. Another explanation may lie in the different risks perceived by agents in the U.S. financial markets. These issues will be discussed in Section 3.

The paper is structured as follows. Section 2 outlines the expectations hypothesis and surveys the results of recent U.S. studies. Section 3 reports the results of the various tests of the expectations theory employed in this paper while Section 4 contains some concluding comments.

2. The Model and Empirical Evidence

The expectations theory has been described extensively in numerous papers, (see Nelson (1972), Shiller (1979), Shiller, Campbell and Schoenholtz (1983)); therefore it will be described only briefly here.

As mentioned earlier, one version of the expectations theory states that the expected return from investing in an n-period bond should equal the expected return from investing in a one-period bond over n successive periods. Thus,

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5. In the U.S. the expectations theory is rejected at each end of the maturity spectrum. The U.S. results which are most directly comparable to those reported here are those dealing with the relationship between various short-term interest rates. There are numerous rejections of the expectations theory at the short end of the market in the U.S. See, for example, Friedman (1979), Shiller, Campbell and Schoenholtz (1983), Mankiw and Summers (1984), Jones and Roley (1983) and Mankiw and Miron (1985).

after linearisation, the long-term interest rate can be expressed as a weighted average of current and expected short-term rates. In a two-period context, the interest rate on a two-period, default-free, pure-discount bond can be approximated by the linear relation,[6]

(1) $$R_t = \theta + \lambda r_t + (1-\lambda)\ E(r_{t+1}|\phi_t)$$

where R_t = per period yield on a two-period bill in period t

r_t = yield on a one-period bill in period t

λ = constant = 0.5

θ = constant risk premium. Under the PEH $\theta = 0$

ϕ_t = information set available to agents at period t

$E(.|\phi_t)$ = operator for mathematical expectation conditional on ϕ_t

The model in equation (1) states, for example, that the yield on a two-period bill equals half the sum of the current one-period bill yield and the expected one-period bill yield in period t+1 plus a constant risk premium.

Given rational expectations, the expected one-period yield, $E(r_{t+1}|\phi_t)$, can be written as,

(2) $$r_{t+1} = E(r_{t+1}|\phi_t) + e_{t+1}$$

where e_{t+1} is a white noise process othogonal to each element of the information set ϕ_t.

Substituting (2) into (1) yields,

(3) $$R_t = \theta + \lambda r_t + (1-\lambda)r_{t+1} + v_{t+1}$$

where $v_{t+1} = -(1-\lambda)e_{t+1}$.

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6. This can be easily generalised to the n-period case and also to allow for coupon payments, see Shiller (1979) and Shiller, Campbell and Schoenholtz (1983). This more general representation, however, must be linearised about the coupon rate (Singleton (1980)) before it can be used for estimation. This linearisation is needed to avoid the criticisms of CIR.

Equation (3) provides the basis for much of the empirical work on this topic. The empirical literature has basically taken two directions. The first, uses variance bounds tests to examine the relative volatility of short and long rates. The second derives a number of testable implications from (3) and subjects them to regression analysis.[7] The present paper will follow this approach.

The rationale underlying the variance bounds tests is intuitively appealing. The term structure relation expresses the long-term interest rate as a weighted average of current and expected short-term rates, thus implying that the variance of the long rate is bounded by the variance of the short rate. This simple observation has underpinned a number of variance bounds tests. In these tests the variance of the short rate, or the variance of the perfect foresight long rate, impose upper bounds on the variance of the long rate or the expected holding period yield.[8]

Both Shiller (1979) and Singleton (1980) have rejected the expectations theory using variance bounds tests. Shiller (1979) derived an inequality restriction which expressed the upper bound on the variance of the expected holding period yield as a function of the variance of short-term interest rates. Using data on 25 year bonds, consols and three to six-month short rates over a number of sample periods, Shiller found that the upper bound was violated in four of the six periods considered.

Singleton (1980), on the other hand, considered the bounds on the variance of the long-term interest rate imposed by the variance of the perfect foresight long rate and the variance of the short rate. Singleton computed consistent estimates of these variances using spectral analysis. He considered three different long rates, (5, 10 and 20-year treasury bonds) while the short

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7. Since, under the null hypothesis, e_{t+1} is orthogonal to each element of the information set ϕ_t, a sufficient condition for consistent estimation of equations derived from (3) will be that r_{t+1} appears on the LHS of the estimating equation and that only elements of ϕ_t appear on the RHS.

8. The perfect foresight long rate, R_t^*, is the value of the long rate would take if agents had perfect foresight about future short rates; thus,

$$R_t^* = \frac{1}{n} \sum_{j=0}^{n-1} r_{t+j}.$$

rate used was the six-month Treasury bill rate. He found that the upper bounds on the variance of the long rate were violated for each maturity. Hence, the expectations theory was rejected for these data.

Subsequently, Flavin (1983) and (1984) has demonstrated that these tests are strongly biased towards rejecting the null hypothesis in small samples.[9] After deriving a bias-corrected measure of volatility Flavin found that the upper bound on the variance of 10-year and 20-year bonds is not violated in postwar U.S. data.

Because of the problem with variance bounds tests, this paper will focus on the second set of tests of equation (1). Equation (1) coupled with the rational expectations assumption in (2), provides a number of distinct testable implications of the expectations hypothesis.

For instance, (1) may be rewritten as,

$$(4) \qquad (1-\lambda)^{-1}R_t - E(r_{t+1}|\phi_t) = \alpha + \beta r_t,$$

where $\alpha = \frac{\theta}{1-\lambda}$ and $\beta = \frac{\lambda}{1-\lambda}$

which, assuming (2), can be estimated as,

$$(5) \qquad (1-\lambda)^{-1}R_t - r_{t+1} = \alpha + \beta r_t - e_{t+1}$$

Equation (5) states that the expected one-period holding-period yield on a two-period bill equals the current one-period bill rate plus a risk premium. This is a two period representation of CIR's Local Expectations Hypothesis. Under the null hypothesis, β should be unity and the residual e_{t+1} should be uncorrelated with all elements of ϕ_t. Jones and Roley (1983), using weekly data on three-month and six-month U.S. Treasury bills, found that the hypothesis $\beta = 1$ cannot be rejected. However, they identified elements of ϕ_t which were correlated with the residual, i.e., that helped explain the

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9. Unbiased estimates of $\sigma^2(R_t^*)$, $\sigma^2(R_t)$ and $\sigma^2(r_t)$ can be obtained if the population means of R_t^*, R_t and r_t are known or if the sample variances are corrected for degrees of freedom when the population means are unknown. However, this latter correction will result in unbiased estimates only if the observations in the sample are not serially correlated. Flavin has argued that there is a high degree of serial correlation in r_t and thus by definition in R_t^*. Consequently, the sample variances of r_t and R_t^* are strongly downward biased estimates of the corresponding population variances.

dependant variable $(1-\lambda)^{-1}R_t - r_{t+1}$. In particular, they found that the level of the six-month yield and foreign holdings of U.S. Treasury bills were significant explanators of the expected holding-period yield. Shiller (1979) also identified a positive correlation between the level of the long rate and expected holding-period return. In both instances, the expectations theory was rejected.

Friedman (1979) found that the forward rate implicit in the yield curve[10] was not an unbiased predictor of future spot interest rates. Friedman identified a positive risk premium in rates on longer securities. Furthermore, this premium was found to vary with the level of interest rates. Clearly, these result are in contradiction to both the pure expectations hypothesis (zero risk premium) and the expectations hypothesis (constant risk premium). In a similar study Shiller, Campbell and Schoenholtz (1983) regressed the change in the three month Treasury bill rate on the expected change measured by the current forward-spot differential. They found that the expected change is a poor and biased predictor of the actual change. This is consistent with Friedman's results.

A further implication of (3) is that when the current long rate is greater than the current short rate, short interest rates are expected to rise above the current long rate. Rearranging (3) provides a test of this hypothesis,

$$(6) \qquad r_{t+1} - R_t = \alpha + \beta(R_t - r_t) + e_{t+1}$$

where $\alpha = \frac{-\theta}{1-\lambda}$ and $\beta = \frac{\lambda}{1-\lambda}$

Under the null hypothesis of the expectations theory $\beta = 1$. Mankiw and Summers (1984) test (6) using data on 20 year securities and six-month securities as long rates and three-month treasury bills as short rates. In both cases the coefficient on the spread between long rates and short rates is not equal to one; indeed it is negative.

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10. In a two period setting R_t is related to r_t and the implicit forward rate F_{t+1} by,

$$(1+R_t)^2 = (1+r_t)(1+F_{t+1}).$$

The implicit forward rate, F_{t+1}, is thus given by,

$$F_{t+1} = \frac{(1+R_t)^2}{(1+r_t)} - 1$$

In a recent paper Mankiw and Miron (1985) examine the expectations hypothesis over a number of data sets representing different periods of monetary control. They examine the ability of the spread between long rates and short rates to predict future changes in short rates. This can easily be derived from (3) and is given by,

$$(7) \qquad r_{t+1} - r_t = \alpha + \beta(R_t - r_t) + e_{t+1}$$

where $\alpha = \frac{-\theta}{1-\lambda}$ and $\beta = \frac{1}{1-\lambda}$

Hence, under the null hypothesis β equals 2. Mankiw and Miron divide their sample into four different periods of monetary control. These are, the period before the introduction of the Fed (1890-1914), the period of the gold standard (1915-1933), the period of interest rate pegging (1934-1951), and the period after the Treasury-Fed Accord (1951-1979). The null hypothesis is rejected in each period with β close to two only in the 1890-1914 period.

Mankiw and Miron attribute this result to the behaviour of the Fed. They derive an expression for β which contains the variance of expected changes in short rates and variations in the risk premium. They show that when the variance of expected changes in short rates approaches infinity the estimate of β approaches two. Whereas when the variance of expected changes in short rates approaches zero it will be dominated by variations in the risk premium and the estimate of β will approach zero. Consequently, they argue that when the Fed has attempted to stabilise interest rates, the variance of expected changes in short rates declines and β approaches zero.

The following section discusses the results of applying several of these tests to Australian short-term interest rate data.

3. Data, Estimation and Results

(a) Some Preliminaries

This paper considers the expectations hypothesis for yields on short-term securities since the introduction in 1979 of the tender system for sales of Commonwealth Treasury notes. The market for 180-day Treasury notes has, however, at times, been very thin; additionally, tenders have been irregular.

Therefore, the data used in the study are for yields on 90-day and 180-day bank-accepted bills.[11]

To obtain a maximum sample size, and therefore improve the precision of the parameter estimates, overlapping weekly observations are used. Consequently, the error terms in the estimated equations will be serially correlated (following a high order process).[12] The effects of this serial correlation on the estimation of the equations and test statistics is corrected by a procedure suggested by Hansen and Hodrick (1980).

A range of tests, based on equation (3) are conducted. In particular, the three equations discussed above (equations (5), (6) and (7)) are estimated for the period January 1980 - March 1986. Also, tests are conducted to see if the implicit forward rate (as defined in footnote 10 above) is an unbiased predictor of future spot interest rates. Several reported equations are augmented by including the level of the long rate as an additional explanatory variable. This is done because U.S. results suggest that the long rate can improve the forecasts of future short rates and also act as a proxy for time varying risk premia.

As mentioned earlier, this sample encompasses two markedly different periods in the Australian financial system. The period January 1980 to December 1983 was a period of managed exchange rates while the subsequent period encompassed a regime of floating exchange rates. Theoretically, interest rates should have been more volatile in the period of managed exchange rates.[13] Furthermore, there were distinct seasonal patterns in financial flows before the float. After the introduction of the float, however, these seasonal

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11. Typically, tests of the expectations hypothesis have used data on yields on government securities to overcome problems of default risk. The 90-day and 180-day bank-accepted bills are backed by Trading Banks. During the sample period under consideration it is unlikely that there has been any perceived solvency risk surrounding Trading Banks and hence the default risk on these bills is likely to be zero. Moreover, at worst, these data would bias the results against the null hypothesis; i.e., towards finding some risk premia.

12. The data are described in the Appendix. There are thirteen weeks between non-overlapping 90-day bills. Hence, the error terms in the equations may exhibit up to twelfth-order serial dependence under the null hypothesis. No tests for higher order serial correlation have been performed.

13. In the sense that unanticipated movements should have been greater. Trevor and Donald (forthcoming) present evidence that this was the case.

financing patterns, while still evident, have diminished.[14] For these reasons the behaviour of interest rates pre- and post-float may have been significantly different. The sample was therefore split into two periods to test for parameter instability. The first encompassed the period of managed exchange rates while the second encompassed the floating exchange rate period. Wald tests of parameter stability (which make use of the asymptotic covariance matrices consistently estimated by the Hansen and Hodrick (1980) procedure) are then calculated.[15]

Furthermore, χ^2 tests of the joint restriction of zero risk premium and the expectations hypothesis (i.e., the PEH) are reported. The test statistic has the form,[16]

$$(8) \qquad \chi^2(m) = T(\hat{\beta}_T - \beta_0)'\hat{\Omega}_T\,(\hat{\beta}_T - \beta_0)$$

where $\hat{\beta}_T$ = estimated parameter vector

$\hat{\Omega}_T$ = consistent estimate of the asymptotic covariance matrix of $\sqrt{T}\hat{\beta}$

T = number of observations

m = number of restrictions

This statistic has a χ^2 distribution with m degrees of freedom.

The full sample and sub-sample results are reported in Tables (1) through (7). These results stand in contrast to most results in the empirical literature to date. In particular, the restrictive pure expectations hypothesis is difficult to reject.

(b) Holding Period Yield

Consider Table 1. The equation tested here is a two-period representation of CIR's so-called Local Expectations Hypothesis. It states that the expected quarterly holding-period yield on a six-month bill equals the present 90-day spot rate plus a constant risk premium. As Table 1 shows $\hat{\beta}$ is significantly different from zero and within two standard deviations of unity in the full sample and in both periods.

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14. See Phillips (1985).

15. For a description of these tests see Tease (1986).

16. See Hansen and Hodrick (1980, p.836).

The PEH implies that $\alpha = 0$ and $\beta = 1$. This joint test cannot be rejected in the full sample or either of the sub-samples. Therefore, the restrictive pure expectations hypothesis cannot be rejected.

The Wald statistic, W(k), indicates that the hypothesis that the parameters are equal over the two sub-periods can be rejected at the 5 per cent but not the 1 per cent level. Thus, although the expectations theory cannot be rejected in either period, it appears that the estimated parameters are sensitive to the sample period chosen.

Table 1

$$(1-\lambda)^{-1}R_t - r_{t+1} = \alpha + \beta r_t - e_{t+1}$$

		Parameter Estimates			PEH	Stability
Obs	Period	$\hat{\alpha}$	$\hat{\beta}$	R^2	$\chi^2(2)$	W(2)
312	Full	-0.58 (0.43)	1.17** (0.12)	0.75	1.83	
193	1	-0.79 (0.47)	1.26** (0.13)	0.79	5.58	8.72*
119	2	0.04 (0.47)	0.92** (0.14)	0.72	5.48	

Notes

Full Sample 4 January 1980 - 21 March 1986
Period 1 4 January 1980 - 9 December 1983
Period 2 16 December 1983 - 21 March 1986

Standard errors are in brackets

(*) Significantly different from zero at the 5 per cent level.
(**) Significantly different from zero at the 1 per cent level.

Table 2 presents estimates of the same equation (equation (5)) with the long rate as an additional explanatory variable. If the expectations theory holds, the expected quarterly holding period yield should be uncorrelated with elements of ϕ_t. Hence, β_2 should equal zero. Using U.S. data Jones and Roley (1983) rejected the hypothesis that $\beta_2 = 0$ and consequently rejected the expectations hypothesis. In the present study, the hypothesis that $\beta_2 = 0$ cannot be rejected in either the full sample or either sub-sample. Furthermore, the pure expectations hypothesis (i.e. the joint test $\alpha = \beta_2 = 0$ and $\beta_1 = 1$) cannot be rejected in any period.
However, the parameters are found to be significantly different in the sub-periods.

Table 2

$$\lambda^{-1}R_t - r_{t+1} = \alpha + \beta_1 r_t + \beta_2 R_t - e_{t+1}$$

		Parameter Estimates				PEH	Stability
Obs	Period	$\hat{\alpha}$	$\hat{\beta}_1$	$\hat{\beta}_2$	R^2	$\chi^2(3)$	W(3)
312	Full	-0.59 (0.51)	1.15** (0.41)	0.02 (0.47)	0.75	1.85	
193	1	-0.63 (0.52)	1.55** (0.21)	-0.32 (0.23)	0.79	6.69	10.96*
119	2	0.28 (0.56)	1.31** (0.19)	-0.45 (0.31)	0.73	6.68	

See footnotes Table 1.

(c) Expected Spread

The results reported in Table 3 for the expected spread between long and short rates (equation (6)) are similar. This equation has a simple interpretation. When the current long rate is greater than the current short rate then short-term interest rates are expected to rise above the current long rate. If the pure expectations theory holds then $\alpha = 0$ and $\beta = 1$. In each period β is within two standard deviations of one. Furthermore, the joint restriction $\alpha = 0$, $\beta = 1$ cannot be rejected. Thus long rates, short rates and expected short rates behave in a way that is consistent with the expectations hypothesis. Once again, however, there is evidence of parameter instability.

Table 3

$$r_{t+1} - R_t = \alpha + \beta(R_t - r_t) + e_{t+1}$$

		Parameter Estimates			PEH	Stability
Obs	Period	$\hat{\alpha}$	$\hat{\beta}$	R^2	$\chi^2(2)$	W(2)
312	Full	-0.01 (0.09)	1.27** (0.40)	0.19	0.46	
193	1	-0.17 (0.09)	1.67** (0.44)	0.31	4.51	8.05*
113	2	0.21* (0.10)	1.08** (0.55)	0.18	4.32	

See footnotes Table 1.

(d) Expected Change in Short Rates

Table 4 considers the power of the spread between current long rates and short rates in predicting future changes in short rates (equation (7)). The pure expectations hypothesis is validated once again. The estimate of β is within two standard deviations of two and the restriction $\alpha = 0$, $\beta = 2$ cannot be rejected. Evidence of parameter instability is again found. The results in this table show that the yield curve is an unbiased predictor of future changes in spot rates. These results are in contrast to those reported in Mankiw and Miron (1985). In their study, the slope of the yield curve was a biased and poor predictor of changes in the spot rate in most samples.

Table 4

$$r_{t+1} - r_t = \alpha + \beta(R_t - r_t) + e_{t+1}$$

		Parameter Estimates			PEH	Stability
Obs	Period	$\hat{\alpha}$	$\hat{\beta}$	R^2	$\chi^2(2)$	W(2)
312	Full	-0.01 (0.09)	2.27** (0.40)	0.43	0.46	
193	1	-0.17 (0.09)	2.67** (0.44)	0.54	4.51	8.72*
119	2	0.21* (0.10)	2.08** (0.55)	0.45	4.32	

See footnotes Table 1.

(e) Implicit Forward Rate

Finally, Table 5 considers whether the implicit forward rate is an unbiased predictor of future spot rates. Like earlier findings, the results in Table 5 are consistent with the expectations hypothesis. In the full sample and the first period the joint restriction $\alpha = 0$, $\beta = 1$ cannot be rejected. However, in the later period this is not the case. It must be remembered, however, that the restriction being tested is actually more restrictive than most tests reported in the literature. Generally, only the value and significance of $\hat{\beta}$ is considered (i.e., the presence of a constant risk premium is allowed). Applying this less restrictive criterion here the expectations hypothesis cannot be rejected in the post float period, since $\hat{\beta}$ is within two standard deviations of one. Once again, there is evidence of instability.

If short term interest rates are serially correlated, then the results reported in Table 5 are weak tests of the expectations hypothesis. Because of this a test of the performance of the forward spot rate differential in predicting changes in the spot rates is reported in Table 6.

Table 5

$$r_{t+1} = \alpha + \beta F_{t+1} + v_{t+1}$$

Obs	Period	Parameter Estimates $\hat{\alpha}$	$\hat{\beta}$	R^2	PEH $\chi^2(2)$	Stability W(2)
312	Full	0.56 (0.51)	0.84** (0.14)	0.55	1.31	
193	1	0.57 (0.56)	0.81** (0.15)	0.55	3.41)	8.16*
119	2	-0.28 (0.43)	1.15** (0.13)	0.71	6.53*)	

See footnotes Table 1.

The results in Table 6 are broadly similar to those in Table 5. The estimate of β is within two standard deviations of one in each period. However, when the more restrictive joint test $\alpha = 0$, $\beta = 1$ is considered, the pure expectations hypothesis is rejected in the pre-float period but not in the full sample or post-float period. Parameter instability is evident once again.

Table 6

$$r_{t+1} - r_t = \alpha + \beta(F_{t+1} - r_t) + v_{t+1}$$

Obs	Period	Parameter Estimates $\hat{\alpha}$	$\hat{\beta}$	R^2	PEH $\chi^2(2)$	Stability W(2)
312	Full	-0.01 (0.09)	1.11** (0.20)	0.44	0.32	
193	1	-0.17* (0.09)	1.30** (0.21)	0.54	6.64*)	7.85*
119	2	0.20* (0.10)	1.02** (0.26)	0.47	4.04)	

See footnotes Table 1.

From Table 5 it can be inferred that the implicit forward rate is an unbiased predictor of future short rates. However, it may not be an optimal predictor in that additional variables which are elements of ϕ_t could improve the performance of the reported equation. Friedman (1979) found that the level of the long rate is a significant predictor of future short rates. Friedman, however, attributed this to a correlation between the level of interest rates and variations in the risk premium. To test for the optimality of the forward rate the current long rate is added to the equation reported in Table 5. The results are in Table 7.

Table 7

$$r_{t+1} = \alpha + \beta_1 F_{t+1} + \beta_2 R_t + \nu_{t+1}$$

		Parameter Estimates				PEH	Stability
Obs	Period	$\hat{\alpha}$	$\hat{\beta}_1$	$\hat{\beta}_2$	R^2	$\chi^2(2)$	W(2)
312	Full	0.58 (0.51)	1.10** (0.38)	-0.27 (0.38)	0.56	1.73	
193	1	0.64 (0.52)	1.45** (0.41)	-0.68 (0.41)	0.59	6.64	10.92*
119	2	-0.30 (0.56)	1.27** (0.41)	-0.12 (0.38)	0.72	6.69	

See footnotes Table 1.

In each period it is found that the level of the long rate cannot improve the forecasting performance of the implicit forward rate. Furthermore, the pure expectations hypothesis (i.e., joint test $\alpha = \beta_2 = 0$, $\beta_1 = 1$) cannot be rejected in any period. The finding of parameter instability is consistent with earlier results. From the results reported in Tables 5 through 7 it is clear that the implicit forward rate is an unbiased and optimal predictor of future short rates.

(f) An Explanation

These results are at variance with most recent studies in the United States. The expectations hypothesis, in its various forms, cannot be rejected in the Australian market for short-term financial assets since 1979. Even the restrictive pure expectation hypothesis is rejected (at the 5 per cent level of significance) in only two of the 21 tests.[17] Moreover, the expectations

17. These two rejections occurr in equations relating to the predictive power of the implicit forward rate.

hypothesis cannot be rejected in either the pre-float or post-float period which are, theoretically, periods of markedly different interest rate behaviour.

However, there is evidence of parameter instability. In each equation the hypothesis that the parameters are equal in both sub-periods is rejected at the 5 per cent level. It must be remembered that this test is actually a joint test that both the parameters are equal in each period and that the residual variances are also equal. To test for a change in the residual variances Goldfeld-Quandt tests of heteroskedasticity were conducted. Although not reported here, these tests indicate that the variances did not change over the sub-periods. The only exception being the equation relating to the implicit forward rate (Table 5). In this instance the hypothesis that the residual variances are equal was rejected at the five per cent level, but not the one per cent level.

There are two possible explanations for these results. The first lies in the difference in monetary regimes and the pattern of financial flows in Australia and the United States. Most U.S. studies use data sampled before 1979. In this period the Fed had a policy of targeting the Federal Funds rate. Hence, interest rates (or more importantly expected interest rates) should have been relatively stable. Mankiw and Miron (1985) argue that such a policy is likely to bias the results toward incorrectly rejecting the expectations hypothesis. On the other hand, monetary policy in Australia throughout much of the data period considered, would appear to have placed greater emphasis on monetary aggregates. In particular, from 1976 to 1984 a projection for growth of M3 was announced in each annual Commonwealth Budget. Given that this projection and knowledge of the seasonal financing patterns (at least in the pre-float period) were part of the markets information set, rational agents would expect (possibly large) changes in interest rates during the course of the year. This expected variation in interest rates is, according to Mankiw and Miron, a necessary condition to achieve estimates of β consistent with the expectations hypothesis. The results reported here suggest that the market did, in fact, use the information on policy and seasonality in a rational way when setting the prices of short-term financial assets.[18]

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18. This is in marked contrast to the behaviour of the foreign exchange market. Tease (1986) shows that the speculative efficiency hypothesis can be easily rejected in the post-float period. Lowe and Trevor (forthcoming) show that exchange rate forecasts prepared by market participants are dominated by simple rules of thumb.

The second explanation lies in the different risk perceptions of agents in the U.S. and Australian financial markets. If the rejection of the expectations hypothesis in the U.S. can be explained by a general aversion to intertemporal uncertainty in the bond market, then it is surprising that the Australian data fail to reject both the pure expectations hypothesis and the expectations hypothesis. This is because financial markets in the U.S. are more developed than those in Australia and, therefore, can provide agents with more opportunities to hedge risk. Thus, one would expect that risk premia should play a more important role in the pricing of financial assets in Australia than they do in the U.S. Since this does not appear to be the case, the rejection of the expectations hypothesis in the U.S. may be attributable to risk factors specific to the U.S. market rather than to a general aversion to uncertainty. For example, agents may be averse to the exposure of U.S. banks to Third World debtor nations - a factor unlikely to be of much importance to the Australian market, at least during the period of this study.[19]

There have been several important changes to the financial system and monetary policy which may affect the results reported here over the course of time. First, the float may significantly reduce the role of seasonality. Although a seasonal pattern is still evident in the market, it may diminish as agents adapt to the new financial environment.[20] Second, in early 1985 the Treasurer announced that the policy of restricting growth in M3 to a conditional projection would be suspended. The conditional projections of M3 growth provided the market with an indication of the stance of policy over the ensuing year. Their suspension, therefore, is likely to have altered the formation of the markets' expectations of future policy movements. This will not alter the results reported here unless policy changes are forecastable. The implicit forward rate, for instance, would remain an unbiased (but less precise) predictor of future spot interest rates. If, however, policy changes are forecastable (in that there is a feedback between major economic aggregates and policy) then the results may change. In particular, if this feedback is part of the information set but is not used by the market then the expectations hypothesis is likely to be rejected. There are not enough observations since the suspension of the conditional projections to test this conjecture rigorously.

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19. This is an example only, it clearly does not explain the U.S. results based on much earlier data.

20. One reason for the remaining seasonality is that market expectations may not have adapted to the new environment.

4. Conclusion

The aim of this paper has been to ascertain the relationship between short-term interest rates for the period after the introduction of the Treasury note tender system in 1979. In general, it has been found that the expectations hypothesis could not be rejected. Furthermore, the hypothesis could not be rejected in the period before or the period after the float.

These results are in marked contrast to the numerous rejections of the expectations hypothesis in the U.S. Possible reasons for this difference are the different policy regimes in place when the studies were conducted and different risk factors in the U.S. and Australian markets. At this stage, these must remain tentative explanations.

If the difference between the results reported here and those in the U.S. can be attributed to risk factors specific to the U.S. market then empirical research, in the U.S., should attempt to incorporate these specific risks rather than attempting to develop models of general intertemporal risk aversion.

The results reported in this paper have focussed on the yields on short-term financial assets. Future research in this area should extend this and consider the relationship between short-term yields and yields on much longer dated securities. To the extent that investment decisions are related to longer term yields, knowledge of the relationship between short and much longer term yields is important for policy purposes.

DATA CONSTRUCTION APPENDIX

Data for the yields on 90-day and 180-day bank-accepted bills were obtained from Securities Market Department of the Reserve Bank of Australia. These data are the mid-points of a range of rates quoted at 12 noon on the Friday of each week in the sample.

In the paper, these yields are expressed as quarterly rates in percentage points. All leads and lags on variables in the various equations are in terms of thirteen weeks - e.g., the estimates of the equation,

$$(1-\lambda)^{-1}R_t - r_{t+1} = \alpha + \beta r_t - e_{t+1}$$

presented in Table 1 refer to the following variables:

r_t - yield on 90-day bank-accepted bill at end of week t

r_{t+1} - the variable r_t shifted forward by thirteen weeks (observations)

R_t - yield on 180-day bank-accepted bill at end of week t.

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