THE INFORMATION CONTENT OF THE TERM STRUCTURE OF INTEREST RATES: THEORY AND EVIDENCE

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INTRODUCTION

Financial market deregulation and innovation have considerably complicated the interpretation of money aggregates in the 1980s, making them less reliable indicators for monetary policy in some countries. As a result, the problem of detecting whether monetary policy is too tight or too loose has become more difficult. A more pragmatic approach to monetary policy formulation has thus emerged in recent years, relying on more judgmental views and taking into account a wider array of potentially relevant financial and real indicators.

Among these indicators there is a growing interest in the term structure of interest rates as a measure of expected inflation[1]. It is argued that a steepening of the yield curve reflects the market's expectation of an acceleration of inflation. If these expectations tend, in the event, to be borne out, such a steepening may warrant a pre-emptive tightening of policy in order to avoid the necessity for much stronger and more disruptive tightening which would be required to counter entrenched inflation at a later stage.

The purpose of this paper is, therefore, to examine the ability of the term structure to forecast future inflation. The rationale for a relationship between these two variables is the following. When asset holders anticipate (for example) an acceleration in future inflation, they tend to substitute out of long-term bonds into short-maturity or real assets *so* as to avoid capital losses. This shift in the composition of demand raises the nominal return on long relative to short assets, i.e. the term structure becomes steeper. Loosely speaking, the more substitutable are assets of different maturities and the closer to rationality are expectations of future inflation, the more accurate will be the term structure's forecast of future inflation.

A model proposed recently by Mishkin is used as a basis for the tests. This model is derived from two maintained hypotheses, namely that the Fisher theory of nominal interest rates fully incorporating inflation premia is correct for assets of all maturities and that expectations of future inflation are formulated rationally. The first of these assumptions is equivalent to assuming the validity of the expectations theory of the term structure of interest rates.

Given the expectations theory's central importance in the term structure literature, a selective review of the arguments and evidence for and against this theory is given in Section I of the paper. The Mishkin model is outlined in Section II. The results of the empirical tests carried out for the United States, Germany, France, the United Kingdom, Italy and Canada are reported in Section III.

I. THE EXPECTATIONS THEORY OF THE TERM STRUCTURE

The expectations theory of the term structure holds that the long-term interest rate is a weighted average of present and expected future short-term interest rates. If future short rates are expected to remain constant, then the long rate will equal the short rate (plus a constant risk premium). However, if future short rates are expected to increase, then the current long rate will exceed the sum of the current short rate and the constant risk premium *so* as to yield the same expected return. Thus, the shape of the yield curve reflects the market's expectation of future short rates of interest. The expectations theory assumes that securities of varying maturities are perfect (ex ante or expected) substitutes for one another.

The expectations theory can be re-stated to imply that expected holding-period returns on bonds of all maturities are identical, or differ only by constant risk premia[2]. Without any loss of generality, let "one period" be defined by the time to maturity of the short bond. The yield and holding return on this short bond held to maturity are, by definition, the same. The "one-period" holding return on a long bond (i.e. a bond of maturity greater than one period) is the return from purchasing such a security, holding it for one period and then selling it at the prevailing price. The term premium is the difference between the expected holding return on the long bond and that on the short bond. Realised, or ex post, excess holding returns are, therefore, identically equal to the sum of the (time varying) term premium (θ_t) and expectational errors (V_{t+1}). The term premium, θ_t, reflects the extra return investors demand for the risk incurred in holding the long rather than the short bond. The expectations theory of the term structure states that θ_t is constant over time, i.e. $\theta_t = \theta$. Assuming expectations to be rational V_{t+1} reflects only 'news' about the long rate. Thus, the joint hypothesis of rational expectations and the expectations theory of the term structure states that observed excess returns cannot be forecast with information available at time t. To refute this joint hypothesis, then, all that is required is to find some variables which are known at time t and which can explain excess holding returns. Lagged values of excess holding returns seem an obvious candidate.

Tests of the expectations theory using excess holding period returns (i.e. the difference between the holding return on the long and short bond[3]) with monthly data were carried out for two periods: 1971M1 to 1979M9 and 1979M11 to

1989M4[4]**.** The results (not reported here in detail but available from the authors) show that either one or more lagged values of excess holding period yields are highly significant for all countries and for both time periods. The joint hypothesis of the expectations theory and rational expectations is thus rejected for all countries in the sample for all long and short rates employed and for both time periods.

Tests of the expectations theory reported in the literature are, as here, mostly unfavourable. The theory is rejected ***conditional*** on some hypothesis about how expectations of future short rates are generated. Recent contributions have assumed rational expectations, in line with the above presentation. However, these tests are inconclusive with respect to the expectations theory due to the fact that they are tests of a ***joint*** hypothesis: the expectations theory of the term structure and rational expectations of future nominal short-term interest rates. The rejection of this joint hypothesis is interpreted by some as implying that risk premia are time-varying (i.e. the expectations theory of the term structure is incomplete) and by others that the long rate over- or under-reacts relative to a rational expectation of future short rates (i.e. rational expectations are rejected). The expectations theory is very demanding in that it requires the long rate to vary exactly one-for-one with variations in expected future short rates and not simply that the correlation between these be positive. Perhaps, therefore, it is not surprising that it tends to be rejected (conditional on rational expectations) in ***so*** many tests (see, for some relatively recent examples, Shiller, **1979,** Jones and Roley, **1983,** Shiller, Campbell and Schoenholtz, **1983,** Mankiw and Summers, **1984,** Mankiw, **1986,** and Campbell and Shiller, **1987).**

Mankiw and Miron **(1986)** have propounded the view that acceptance or rejection of the expectations theory is sensitive to how the monetary authority implements monetary policy. The U.S. Federal Reserve relinquished the objective of interest rate smoothing between October **1979** and October **1982,** adopting, at the end of this time period, a policy of partial smoothing. Mankiw and Miron argue that this change in operating procedure will render interest rate behaviour more favourable to the expectations theory after October **1979.** One test taking into account this argument (Hardouvelis, **1988),** although rejecting the strict expectations theory, nevertheless finds that it has considerable forecasting power after **1979** but only modest forecasting ability before **1979,** lending some corroboration to the Mankiw-Miron hypothesis. Another test using survey data on interest rate expectations (Froot, **1988)** reports similar results. Indeed, for one category of long-term bonds the latter fails to reject the expectations theory. Others (Hamilton, **1988,** Fama and Bliss, **1987,** and Fama, **1988)** also find increasing support for the view that movements in long rates are highly correlated with movements in rationally expected future short rates (under the assumption of rational expectations) while still being unable to fully accept the expectations theory.

Despite the more favourable flavour of recent tests, unequivocal empirical support for the strict expectations theory is rarely reported. To what, then, is the poor performance of the expectations theory attributable? Two main factors have been advanced as explanations: time-varying risk premia and some degree of market segmentation due to the preferred habitats of investors[5].

Explaining the failure of the expectations theory in terms of ***time-varying risk premia*** is somewhat vacuous in the absence of a further theory as to why risk premia themselves vary with time. Indeed, the most striking rejections of the expectations theory tend to be for data drawn from the short end of the maturity spectrum such as 3- and 6-month Treasury bills[6]. Explaining the failure of the theory by time-varying risk premia in such cases requires identifying huge variations in risk within very short time periods, for which the objective conditions do not seem to exist. Much effort has nevertheless been devoted to the search for measures of risk that prove to be significantly related to the long-short spread but there has been little success'.

If it were possible to obtain such a measure of risk, and providing expectations are rational, the slope of the yield curve could be purged of this element, leaving the remainder to be interpreted in accordance with the tenets of the expectations theory. In other words, an upward-sloping ***risk-adjusted*** yield curve would then reflect the markets' expectation of a future increase in short rates and ***vice versa.***

An alternative explanation for the apparent inadequacies of the expectations theory is that ***the term structure is compartmentalised*** to a greater or lesser extent. If so, changes in the supply and demand for securities of a particular maturity will affect the shape of the yield curve. The view that borrowers and lenders may have preferred habitats at particular maturities which suit their specific needs, and from which they can only be coaxed by large interest differentials, has long been debated.

Empirical investigations of this theory have used structural supply-demand models and have employed the restrictions suggested by the theory of portfolio behaviour (see Brainard and Tobin, 1968, and Smith, 1975), to constrain the term structure equation implied by the data-admissible structural portfolio model. Friedman and Roley, (Friedman, 1977, 1979, Roley, 1981, and Friedman and Roley, 1979) have examined the determinants of the term structure using this approach. Their implied expressions for the term structure of interest rates fit the data about as well as does the typical alternative approach of a single-equation reduced form. Nevertheless, their simulations for the effects of debt management policies indicate that such effects are quite small. Other single equation evidence also suggests that debt management effects are small (e.g. Modigliani and Sutch, 1966). Other authors, who have attempted more recently to test for market segmentation by including proxy variables in single-equation reduced form term structure equations, have also found only limited support for this view[8].

Since the expectations theory is couched in terms of unobserved variables (expected future short rates), it can, as noted above, only be tested by first assuming some expectations generating scheme for short rates. The typical assumption is rational expectations. Thus, another possible reason why the expectations theory tends to be rejected by the data is because this expectations assumptions is incorrect. Shiller (1979), Shiller, Campbell and Schoenholtz (1983), Campbell and Shiller (1987), and Mankiw and Summers (1984) have tested the hypothesis that myopic expectations provoke an ***overreaction in*** long ***rates*** to changes in contemporaneous short rates and that this is responsible for the rejection of the expectations theory. However, empirical tests of this excess sensitivity hypothesis decisively reject it. Long rates are found to underreact to ***current*** short rate changes. As a result, the long rate reacts disproportionately to expected ***future*** short spot rates. In sum, the excess sensitivity hypothesis cannot account for the observed departures from the predictions of the combined rational expectations and expectations theory of the term structure.

In sum, although the strict expectations theory is almost invariably rejected in empirical tests, neither time-varying risk premia nor segmentation effects appear to be of such a large magnitude as to significantly compromise the potential of the term structure to act as a leading indicator of future inflation changes. Of course, empirical rejection of the expectations theory does not necessarily mean that there is no information in the term structure that is useful as an indicator for monetary policy. Variation in expected future short rates could still account for the bulk of the systematic variation in the current long rate. The problem is in extracting this information, given the noise created either by time-varying risk premia or by excess volatility of the long rate, or both (see below). In consequence the usefulness of the term structure as a forecaster, or leading indicator, of future inflation does not require the expectations theory to be strictly valid, although if it were valid then the term structure would evidently be a better forecaster.

II. THE TERM STRUCTURE AS A LEADING INDICATOR OF INFLATION

A. The Mishkin model

Mishkin (1988) has recently proposed a model for testing the forecasting ability of the term structure for future inflation. His model (for a detailed derivation see Annex 1) is based on the Fisher equation for different maturities plus rational expectations. His forecasting equation is:

$$\pi_{m,t} - \pi_{n,t} = \alpha_{m,n} + \beta_{m,n}\,[i_{m,t} - i_{n,t}] + \eta_t, \qquad [1]$$

where

$$\alpha_{m,n} = \bar{rr}_n - \bar{rr}_m,$$

$$\eta_t = \epsilon_{m,t} - \epsilon_{n,t} - u_{m,t} + u_{n,t},$$

$$u_{m,t} = rr_{m,t} - \bar{rr}_m,$$

$$u_{n,t} = rr_{n,t} - \bar{rr}_n.$$

$\pi_{m,t}$ and $\pi_{n,t}$ are the realised inflation rates over m and n periods corresponding to the maturities of the long and short assets, the yields of which are denoted by $i_{m,t}$ and $i_{n,t}$ respectively. $rr_{m,t}$ and $rr_{n,t}$ are the corresponding ***ex ante*** real yields. The intercept in equation [1] is the difference in the average ***ex ante*** real rates over the corresponding maturities. $\epsilon_{m,t}$ and $\epsilon_{n,t}$ are inflation forecast errors over the m– and n– period horizons. If the real term structure is constant, (i.e., $u_{m,t}, u_{n,t} = 0$), the somewhat complex error term boils down to $\epsilon_{m,t} - \epsilon_{n,t}$. Given rational expectations the latter is independent of the contemporaneous nominal term structure (i.e., $i_{m,t} - i_{n,t}$), a necessary condition for consistent estimation. However, in the present context of overlapping observations this condition does not guarantee efficient parameter estimates (this issue will be taken up below).

For heuristic purposes it is worthwhile noting that equation [1] can be rewritten as:

$$rr^P_{m,t} - rr^P_{n,t} = -\alpha_{m,n} + (1 - \beta_{m,n})\,[i_{m,t} - i_{n,t}] - \eta_t, \qquad [2]$$

where rr^P_m and rr^P_n are the ***ex post*** real interest rates on m and n period bonds at time t.

If $\beta_{m,n}$ in equation [1] is significantly different from zero then the slope of the term structure, defined as $i_m - i_n$ in period t, has ***some*** power to forecast inflation over the time interval $(t+m) - (t+n)$, $n < m$. The rejection of $\beta_{m,n}$ equal to zero would be tantamount to a rejection of the null hypothesis that the term structure of ***ex post*** real interest rates moves ***pari-passu*** with the term structure of nominal interest rates (see equation [2]). On the other hand, a rejection of the hypothesis that $\beta_{m,n}$ equals one is a rejection of the hypothesis that ***all*** changes in the slope of the nominal term structure arise from variations in (rationally expected) inflation and that the slope of the real term structure remains invariant with time. This case can also therefore be interpreted as saying that changes in the slope of the nominal term structure are informative with respect to changes in the slope of the term structure of real interest rates. Allowing for the possibility that $\beta_{m,n}$ is not equal to one, but continuing to assume rational expectations and constant real rates, equation [1] yields:

$$\alpha_{m,n} = [(\beta_{m,n}\, i_{n,t} - E_t\pi_{n,t}) - (\beta_{m,n}\, i_{m,t} - E_t\pi_{m,t})]$$

which is equal to the difference between the (constant) ***ex ante*** real interest rates only if $\beta_{m,n} = 1$. Care should therefore be exercised in interpreting the $\alpha_{m,n}$ coefficient in the results reported below.

B. Measurement

It is important to note that the interest and inflation rates entering equation [1] are expressed in an analogous manner on a continuously compounded basis. To be more precise, let $i^A_{j,t}$ be the annualised interest rate prevailing in period (month) t (i.e., 8.65 per cent on 3-month treasury bills in April 1989 for the U.S. economy, for example). The continuously compounded rate over the j periods (months) to maturity of the rate prevailing in period t converted to an annualised basis is:

$$i_{j,t} = [(i^A_{j,t}/100 + 1)^{j/12} - 1]*100, \qquad j = m,n.$$

Hence it represents the nominal return on investing 1 dollar in a fixed rate bond that matures at t +j.

The inflation rate is compounded, not over the actual rate prevailing in period t (as with the interest rate), but over the actual inflation rates prevailing in all the future periods to maturity of the corresponding asset whose yield is being compared with the inflation rate, i.e.:

$$\pi_{j,t} = [(\pi^A_t/100 + 1)^{1/12}(\pi^A_{t+1}/100 + 1)^{1/12} \ldots\ldots (\pi^A_{t+j-1}/100 + 1)^{1/12} - 1]*100 \qquad j = m,n$$

where π_t is the annualised inflation rate in period t. Note that, given these definitions, the constant term as estimated in the equations reported below is the average difference over the sample period between the cumulative or compound annualised real ***ex post*** returns on an m- and an n-period bond (m > n) conditional on $\beta_{m,n}$ being equal to one. To convert this o an annualised per period difference $\tilde{\alpha}_{m,n}$, a more easily interpretable measure, requires the following transformation:

$$\tilde{\alpha}_{m,n} = [\alpha_{m,n}]^{1/(m-n)}$$

The sign of $\tilde{\alpha}_{m,n}$ depends on whether the term structure of real interest rates is upward or downward sloping.

C. Some econometric considerations

The literature on estimation of models containing rational expectations has proposed three basic different approaches:

i) A ***full-information*** method (see, for example, Hansen and Sargent, 1980 and 1982) consisting of the estimation of the complete simultaneous system. The model solution is used to generate the expectational variables, giving rise to cross equation parameter constraints that assure consistency. The efficiency gains of this approach must be weighted against the difficulties in estimating a (possibly) large system, and against the risk that mis-specification in any single equation will lead to inconsistent estimates of all the parameters.

ii) The ***unobservable variable approach*** does not impose a unique (convergent) solution to the rational expectation model, but treats expectational variables as unobservable state variables and applies Kalman filtering techniques. This approach is useful when testing for bubble-like phenomena, *so* that one does not want to rule out non-convergent solutions in the first place (see Burmeister and Wall, 1982). This is not the case here.

iii) The most popular (and easy) method is the ***limited information*** approach proposed by McCallum (1976). This consists of substituting actual future for expected values in the equations (see Annex 1), and using instrumental variables to take care of the non-independence of those regressors not entering the conditioning information set and the "composite" error term, which now includes structural disturbances as well as forecast errors.

OLS estimation is problematic because a phenomenon of "overlapping observations" arises in models with forward-looking expectations whenever the sampling interval is finer than the forecast horizon. Forecast errors are not known until the forecast horizon is reached. Referring back to equation [1], we see that the forecast error term $\epsilon_{m,t}$ is not realised until period $t+m$. Rational expectations can only rule out any correlation between the forecast error (realised at $t+m$) and the variables entering the conditioning information set (at t) which includes forecast errors realised at t or earlier. They cannot rule out serial correlation in forecast errors realised from $t+1$ to $t+m$, since they do not enter the conditioning information set. Hence, the composite error term is likely to follow a MA process of order $(m-1)$[9]. The consequence is the familiar efficiency loss of OLS estimates and inconsistent standard errors due to serial correlations. This shows up in virtually all the OLS estimated equations in the very high values of the χ^2_A tests for autocorrelation (see below).

When the structure of ***real*** interest rates does not change over time, OLS estimates of the equation [1] will produce consistent estimates of $\beta_{m,n}$. This is *so* because the composite error term, η, will then involve only the forecast errors for inflation, the ϵs, that will be orthogonal to all information available at time t, including the nominal interest rate structure. Conversely, if the real interest rate structure varies over time, $\beta_{m,n}$ will no longer be an optimal predictor, as the information contained in the u variables will not be exploited. Moreover, if the nominal and real yield slopes are correlated, the composite error term η will be correlated with $i_m - i_n$, and OLS estimates of $\beta_{m,n}$ (and $\alpha_{m,n}$) will be inconsistent. Hence some instrumentation, as in the McCallum procedure, is called for.

The McCallum technique must, however, be extended to handle the case of serially correlated structural disturbances as may occur in the case in hand. Standard ways of correcting for autocorrelation (for example, the generalised two stage least squares proposed in Theil, **1961**), will make things "worse" as they would re-introduce inconsistency through the "filtering" of the instrumental variables and of the "composite" error term (see Cumby, Huizinga and Obstfeld, **1983).** The two-step-two-stage least squares **(2S2SLS)** estimator employed here is designed to achieve efficient estimates by removing autocorrelation in residuals ***without*** loosing consistency in a rational expectation context. It is consistent, asymptotically normal and asymptotically efficient in the class of "generalised method of moments" estimators developed by Hansen **(1982).** All that is required is the existence of some instruments that are "predetermined" with respect to the error term, and that serial correlation in the model dies out in a ***finite*** time. In order to assure that both these conditions are met, the autoregressive error component had to be removed from the error term, the observations quasi-differenced (filtered) and appropriately lagged instruments have to be used (with the "nearest" lag bigger than the order of the moving average component of the error term).

Equation **[1]** can be specified for a variety of forecasting horizons. These forecasting horizons overlap for the interest rates chosen for the present exercise. However, since contemporaneous forecasting errors for different term structure forecasting equations are likely to be highly correlated, estimating these separate equations as a system using SURE estimation is likely to enhance the efficiency of parameter estimation.

The first step therefore consists in obtaining a consistent estimate of the system by applying instrumental variables. The instrumental variables employed were industrial output and lagged interest rates. From this, one recovers a consistent estimate of the variance-covariance matrix of residuals, and then applies GLS to the system (second step). This estimator accounts for serial correlation in the error terms as well as cross-equation correlation in disturbances, while preserving consistency through instrumental variables. Furthermore, it takes account of heteroskedastic residuals (a problem in only a few of the equations) when obtaining an estimate of the variance-covariance matrix in the first step. It employs the weights for the sample moments suggested by Newey and West **(1986)** for the case in which this matrix is not positive definite.

III. RESULTS

A. Ordinary least squares

In order to gauge the efficiency gains from employing the 2S2SLS technique, equation [1] was first estimated by OLS for all forecasting horizons and for all the countries in the sample. The detailed results are reported in Table A1 in Annex 2. Since OLS is not the appropriate estimation technique, the results merit only a brief comment[10].

These results suggest that the term structure for maturities at the short end of the maturity spectrum (from 3 to 12 months maturities) contain information about inflation over the corresponding future horizons (see the results for the United States, France and Italy). However, except for Germany, and perhaps Canada, the slope of the term structure fails to forecast inflation when rates at the longer end of the maturity spectrum are employed as long rates[11]. This result is not at all implausible. The more distant the future horizon the greater the scope for unpredictable shocks to influence the actual outcome for future inflation. For the slope of the yield curve formed by the 5-year government bond rate, as the long rate, to have been a good predictor of changes in the rate of inflation in the 1970s, for example, would have required bond holders to have been able to predict the first and second oil crises and also general government policy reaction to these events. For the 1980s, bondholders would have had to predict the disinflationary Federal Reserve policy of the early 1980s. They would, furthermore, have had to been able to make predictions of these events approximately 5 years in advance of their occurrence.

There are several disappointing features of the OLS results, however. Only one "non-perverse" result was estimated for the United Kingdom. Several of the β estimates are significantly in excess of unity. Also, there is a good spattering of negative βs although, with the exception of the results for the United Kingdom and $\beta_{60,3}$ for Germany, these are not statistically significant.

B. Two-step two-stage least squares results

The order of the moving average processes for the error terms was not always that suggested by theory (i.e., $m - n$) particularly when this was very large. Rather, the order was determined pragmatically using the equation diagnostics.

The 2S2SLS results[12] are reported in Table 1. In one qualitative respect, the overall inference from the 2S2SLS results is not altered *vis-à-vis* that inferred from the OLS results. The term structure has, generally, better forecasting ability

Table 1. Term structure forecasting equations for inflation
(2S2SLS estimates)

Months (m,n)	$\alpha_{m,n}$	$\beta_{m,n}$	81	$\bar{R}^2$
(6,3)	0.86	0.49	1.11	0.979
	(2.76)	(3.88)	(21.32)	
(60,3)	0.04	0.32	0.99	0.990
	(0.00)	(1.07)	(135.74)	
(60,6)	0.52	0.35	0.98	0.989
	(0.02)	(1.09)	(126.17)	
(60,3)	37.59	-0.39	0.87	0.948
	(12.26)	(4.47)	(28.89)	
(132,3)	13.40	0.27	0.81	0.958
	(2.51)	(5.86)	(20.99)	
(132,6)	-20.68	0.53	0.77	0.958
	(4.73)	(10.33)	(18.57)	

$\hat{\epsilon}_{1t}$ ~

Months (m,n)	$\alpha_{m,n}$	$\beta_{m,n}$	81	$\bar{R}^2$
(3,1)	47.10	0.57	1.00	0.992
	(0.00)	(2.31)	(44.28)	
(6,1)	−2.46	0.38	0.99	0.998
	(0.49)	(1.73)	(110.02)	
(6,3)	−0.32	0.53	0.97	0.994
	(0.50)	(1.76)	(65.88)	

$\hat{\epsilon}_{1t}$

Months (m,n)	$\alpha_{m,n}$	$\hat{\beta}_{m,n}$	61	R^2
(60,3)	−27.32	−0.01	0.99	0.999
	(0.01)	(0.32)	(17.13)	
(120,3)	229.60	0.01	1.04	0.999
	(6.66)	(0.28)	(48.10)	
(120,6)	125.88	0.03	1.04	0.999
	(4.78)	(0.88)	(61.49)	

Cross equation correlation matrix:

1.00		
−0.27	1.00	
−0.39	0.98	1.00

$\hat{\epsilon}_{1t} \sim$ ARMA (1,5), $\hat{\epsilon}_{2t} \sim$ ARMA (1,5), $\hat{\epsilon}_{3t}$ − ARMA (1,5)

(6,3)	−0.01	0.92	0.99	0.886
	(0.04)	(3.16)	(256.10)	
(12,3)	0.26	2.39	1.03	0.960
	(1.00)	(8.16)	(39.37)	
(12,6)	0.02	0.50	1.01	0.973
	(0.16)	(4.71)	(55.01)	

(24,3)	14.19	0.25	1.04	0.993
	(6.49)	(2.58)	(64.21)	
(48,3)	37.26	0.14	1.02	0.997
	(7.06)	(2.21)	(134.21)	
(48,24)	25.79	0.08	1.01	0.997
	(2.43)	(1.11)	(166.60)	

$\hat{\epsilon}_{1t}$ $\sim$

for future inflation over shorter than longer horizons, with the one outstanding exception again being Germany. However, in several other respects the results are a considerable improvement over the **OLS** results. β estimates in excess of unity vanish (except for $\beta_{12,3}$ for Italy). Negative βs also virtually disappear. Indeed, the only outstanding unsatisfying feature of the **2S2SLS** results is the statistically significant negative value for $\beta_{60,3}$ for Germany.

For the United States all the $\beta_{m,n}$ parameter estimates lie between zero and one but, in common with the **OLS** estimates, only the $\beta_{6,3}$ coefficient is significantly different from zero. The slope of the nominal term structure at the short end of the maturity spectrum contains information about future inflation over the time interval from 3 to 6 months ahead. This contrasts with a result reported in Mishkin **(1988)** that "the term structure for maturities of six months or less contrains almost no information about the path of future inflation" (p. 16, op. *cit.*). However, since $\beta_{6,3}$ is also significantly different from one the nominal term structure also contains information about the term structure of real interest rates. When the 5-year rate is used as the long rate the results suggest that the term structure of nominal rates only contains information about the real term structure.

For Germany, the $\beta_{m,n}$ parameter estimates are not greatly altered from the corresponding **OLS** estimates with both $\beta_{132,3}$ and $\beta_{132,60}$ continuing to lie between zero and one again reflecting information in the term structure about both the real term structure and future inflation. As with the **OLS** estimates, the German **2S2SLS** results are again the only ones which suggest forecasting power for the term structure for inflation over very long time horizons[13].

The estimated $\beta_{m,n}$ parameters are all lower and their corresponding estimated standard errors are all much higher for **2S2SLS** compared to **OLS** for France. Inferences from these results are also qualitatively different in that the **2S2SLS** $\beta_{m,n}$**'s** all lie between zero and one with both $\beta_{6,1}$ and $\beta_{6,3}$ significantly different from zero only at the **10** per cent level. For the United Kingdom no forecasting power for the term structure is detectable with all s being effectively zero.

Except for the implausible value for $\beta_{12,3}$ the results for Italy suggest considerable forecasting ability for inflation. $\beta_{6,3}$ is not significantly different from one while $\beta_{12,6}$ lies half-way between zero and one. The **2S2SLS** results for Canada **also** indicate some forecasting power for future inflation which, as per other countries, fades with lengthening maturity and forecast horizons.

Overall, when a β estimate is significantly positive, it also tends to be significantly less than one, the theoretical value under the joint null hypothesis of the expectations theory of the term structure and rational expectations. The one exception is $\beta_{12,3}$ for Italy which is significantly greater than one. The **2S2SLS** estimation technique accounts for possible sources of simultaneity and thus least-squares bias can be dismissed as the cause. Thus, all the results tell us is that the joint null hypothesis is rejected without allowing us to infer which of the constituent hypotheses is responsible[14].

To the extent that the slope of the nominal yield curve fails to predict future inflation it, by definition, reflects variation in the slope of the ex post real yield curve (refer back to equation [2]). Changes in the real term structure affect the real economy. Therefore, the following general pattern could be tentatively inferred from the cross-country results reported: changes in the slope of the term structure, formed by using a long interest rate from the long end of the maturity spectrum, reflect exclusively changes in the real term structure (except for Germany) and may, therefore, forecast future output changes: changes in the slope of the term structure based on "long" and short rates taken from the short end of the maturity spectrum are informative about both future inflation and the real term structure and, via the latter, future real output.

IV. CONCLUSIONS

The breakdown of the long-run relationship relating nominal money to real output and prices has left monetary policy formulation in some OECD countries without a reliable anchor. Monetary aggregates are now being supplemented by an array of real and financial indicators in a more eclectic approach to policy formulation. One of the financial variables that attracts considerable attention in this respect is the term structure of interest rates, which has been advanced as an indicator of the markets' expectation of future inflation. If these expectations are responding to actual current incipient inflationary pressures that have not yet become known through published price indices, then such an indicator would be extremely valuable. It would allow the monetary authorities to take pre-emptive action to prevent inflation emerging and becoming ingrained in peoples' expectations.

Tests of the predictive power of the term structure for future inflation have been carried out here for six major OECD countries and for a variety of asset maturities. Careful attention had to be paid to econometric issues. To obtain consistent and relatively efficient parameter estimates in a context of rational expectations, overlapping observations, simultaneity and the likely contemporaneous cross-correlation in the forecast errors for different, but overlapping forecast horizons, a special estimation procedure had to be employed. A general pattern emerges from the results. The term structure does have considerable forecasting ability but this fades as yields on assets of increasingly distant maturities are employed as "long" rates. For those countries that conform to the general pattern, the term structure at the longer end of the maturity spectrum reflects variation in the real term structure and may thus contain useful information about the evolution of future output.

Thus, the current practice of using the spread between a very long rate (typically a 10-year government bond rate) and the 3-month treasury bill rate, as the short rate, to provide an indicator, albeit tentative, of the markets' expectation of future inflation may be suspect. The results obtained here would suggest that yields taken from the shorter end of the maturity spectrum (in the region of, say, three months to two years) are more reliable indicators of the markets' expectation. The use of such yields would also have the advantage of conforming more closely to the likely policy horizons of central banks.

NOTES

1. Among other indicators that have been cited are commodity prices, the exchange rate, credit aggregates, cyclical indicators of real activity, as well as, of course, money aggregates.
2. See Cox, Ingersoll and Ross (1981). Using linearised holding period yields the equivalence is exact as demonstrated by Shiller, Campbell and Schoenholtz (1983).
3. The holding return on the long bond (H_t) is defined as:

$$H_t = \frac{1}{P_t} + \frac{P_{t+1} - Pt}{P_t}$$

$$= R_t - \frac{(R_{t+1} - R_t)}{R_{t+1}}$$

   where R_t is the yield on a long-term coupon bond which can be approximated as a consol. Thus, $R_t = 1/P_t$ when P_t is the price of the consol. For a one-period asset, the holding return equals the yield.
4. The choice of sample period in conducting such tests is important. Some researchers (Blanchard, 1984, Mankiw and Miron, 1986, and Belongia and Koedijk, 1988) have argued, in the United States context, that tests of the expectations theory of the term structure are likely to be sensitive to the changes in the monetary control procedures implemented by the Federal Reserve in October 1979. These changes allowed short-term interest rates to fluctuate much more than under the previous procedure. Given the importance of the United States in the world financial system, this increased variability is likely to have been transmitted to other OECD countries, also implying structural change for them after October 1979. This is discussed further below.
5. Segmentation effects could also arise from various types of official regulation and tax policies.
6. There are numerous examples. For a recent selection of the literature see, for example, Jones and Roley (1983), Shiller, Campbell and Schoenholtz (1983), Mankiw (1986), and Hardouvelis (1988).
7. The risk measure employed by Mankiw (1986) is the absolute value of the percentage first forward difference in the long bond yield. Another measure frequently used is a moving standard deviation of short rates (see, for example, Modigliani and Shiller, 1973, and, more recently, Jones and Roley, 1983). Other measures of interest rate variability have been used by Mishkin (1982) and Bodie, Kane and MacDonald (1984). These measures tend to be ex *post* measures of volatility and to that extent are less than ideal.
8. Jones and Roley (1983) test for four such variables; treasury bill supplies, the unemployment rate, a risk variable and foreign holdings (specifically foreign central bank

holdings of U.S. treasury securities), in a test of the joint rational expectations-expectations theory of the term structure hypothesis. The authors find that the last-mentioned variable has a significant effect at the 5 per cent level in a model that rejects the joint null hypothesis. Their rationale for this effect is as follows. Foreign central banks have a "preferred habitat" in three-month U.S. treasury bills. When investors observe high foreign holdings of treasury bills, they expect further purchase of this security in the next period and thus a lower short-term interest rate and an increase in the six-month holding-period yield. This implies that the risk of a capital loss is reduced, which lowers the required term premium. Shiller, Campbell and Schoenholtz (1983) also test to see if the relative volume of trade in securities at either end of the maturity spectrum succeeds in explaining the term premium. They find that the volume variable does help to explain excess holding returns and indeed it displaces the risk variable (measured as a moving standard deviation of the short interest rate) which was significant in an excess return equation when entered without the volume variable.

9. Of course, if the real interest rate is not constant, $u_{n,t}$ could follow a higher-order MA process than $(m - 1)$. However, allowing for higher order processes does not qualitatively alter the results.
10. The actual forecasting horizon chosen is dictated by data availability considerations. A full description of the data is given in Annex 3 of the paper.
11. For a fairer comparison to have been made, at least two rates at the short end, and at least one at the long end would have had to have been available. Unfortunately this was not always the case.
12. Note that using the 10-year government bond rate is not possible in the post-October 1979 regime since the calculation of the cumulative inflation rate over the following 10 years for April 1979 uses up all the remaining observations to April 1989. More generally, the data transformations required to obtain cumulative inflation over the relevant future maturity horizon, combined with the 2S2SLS estimation procedure, involves such a large loss of observations that estimation over the post-October 1979 period is not feasible except for France and Italy where only relatively short maturity assets are employed.
13. The yield on the secondary market on public sector bonds are quoted for (3 – 7) and (7 – 15) year maturities. In both cases the mid-point of these ranges was chosen as the actual term to maturity. Data availability constraints dictated that tests could only be carried out for Germany by combining yields on public and private sector securities.
14. The magnitudes of some of the constant terms in the equations, although apparently very large, are of a reasonable size when transformed as indicated in Section III.B in the text. Also note that the efficiency gains from SURE estimation can be appreciated by noting the very large cross-equation correlations between the errors in different equations.

Annex 1

THE MISHKIN MODEL IN DETAIL

Mishkin (1988) builds on Fama's (1975) study on variations in the level of nominal interest rates as forecaster of future inflation rate movements. According to the Fisher equation expected inflation over m periods is equal to the m-period nominal interest rate minus the m-period ex ante real interest rate, i.e.:

$$E_t\pi_{m,t} = i_{m,t} - rr_{m,t} \qquad [A1]$$

where $E_t\pi_{m,t}$ is the expectation of inflation over the m period horizon at time t, $i_{m,t}$ is the m-period nominal interest rate of time t and $rr_{m,t}$ the m-period ex ante real interest rate at time t. It is assumed that inflationary expectations are generated rationally, i.e.:

$$\pi_{m,t} = E_t\pi_{m,t} + \epsilon_{m,t} \qquad [A2]$$

where $\epsilon_{m,t}$ is the forecast error over the full m periods. Substituting A2 into A I and rearranging yields:

$$\pi_{m,t} = \alpha_m + \beta_m\, i_{m,t} + \eta_{m,t} \qquad [A3]$$

where:

$$\alpha_m = -\overline{rr}_{m,t}$$

$$\beta_m = 1$$

$$\eta_{m,t} = \epsilon_{m,t} - u_{m,t}$$

$$u_{m,t} = rr_{m,t} - \overline{rr}_m$$

To examine the information in the term structure of interest rates, as opposed to the level of the interest rate, about future changes in the inflation rate, Mishkin substracts equation A3 for the n-period inflation rate from equation A3 for the m-period inflation rate which gives the following:

$$\pi_{m,t} - \pi_{n,t} = \alpha_{m,n} + \beta_{m,n}\, [i_{m,t} - i_{n,t}] + \eta_t \qquad [A4]$$

where:

$$\alpha_{m,n} = \overline{rr}_n - \overline{rr}_m$$

$$\beta_{m,n} = 1$$

$$\eta_t = \epsilon_{m,t} - \epsilon_{n,t} - u_{m,t}\ {}_{-}\ u_{n,t}$$

$u_{n,t} = rr_{n,t} - \overline{rr}_{n,t}$, and

$\epsilon_{n,t}$ = the forecast error over the n-period horizon.

As the estimated value of $\beta_{m,n}$ varies from zero to one the information content of the yield curve for future inflation increases while its information content for the term structure of real interest rates decreases.

Annex 2

TABULAR OLS RESULTS

Months (m,n)		$\alpha_{m,n}$	$\beta_{m,n}$	$\hat{R}^2$	$\hat{F}(l)$	$\hat{\chi}^2_A(12)$	$\hat{\chi}^2_B(1)$
(6,3)	(a)	−0.65	1.52	0.75	282	76	0.4
		(4.11)	(16.80)				
	(b)	−0.55	0.84	0.42	78	97	1.9
		(2.46)	(8.84)				
(60,3)	(a)	43.1	0.17	0.02	2.8	91	14.6
		(10.051	(1.67)				
	(b)	30.9	−0.09	0.01	1.4	48	9.7
		(5.06)	(1.20)				
(60,6)	(a)	43.7	0.11	0.00	1.1	91	9.6
		(9.79)	(1.04)				
	(b)	30.1	−0.11	0.03	2.5	48	9.6
		(5.64)	(1.59)				
(60,3)	(a)	29.4	−0.12	0.49	65.2	65.4	8.3
		(43.0)	(8.08)				
	(b)	0.75	0.25	0.07	4.8	49.8	10.6
		(0.13)	(2.20)				
(132,3)	(a)	19.15	0.20	0.88	374.4	37.2	4.1
		(12.64)	(19.35)				
	(b)	—	—	—	—	—	—
(132,6)	(a)	−6.38	0.32	0.90	462.3	40.3	7.5
		(4.32)	(21.50)				
	(b)		—	—	—	—	—
FRANCE							
(3,1)	(a)	−0.51	1.06	0.53	146	123.8	23.6
		(3.26)	(12.09)				
	(b)	−1.27	1.41	0.80	438	91.7	6.0
		(10.21)	(20.94)				
(6,1)	(a)	−1.56	1.11	0.51	134	125.6	22.8
		(3.59)	(11.58)				
	(b)	−3.63	1.49	0.80	431	94.8	3.6
		(10.83)	(20.76)				
(6,3)	(a)	−1.01	1.13	0.49	123	126.1	19.9
		(3.66)	(11.08)				
	(b)	−2.32	1.52	0.78	388	95.9	2.7
		(10.61)	(19.70)				

(60,3)	(a)	107.9 (13.2)	-0.24 (2.00)	0.03	4.0	92	3.3
	(b)	6.30 (0.97)	0.33 (4.26)	0.25	18.1	46	27.5
(120,3)	(a)	284.9 (11.62)	-0.42 (3.79)	0.13	14.3	85	0.2
	(b)	—	—	—	—	—	—
(120,6)	(a)	171.0 (10.26)	-0.47 (4.37)	0.17	19.1	85	0.3
	(b)	—	—	—	—	—	—
				ITALY			
(6,3)	(a)	-0.92 (3.06)	1.04 (13.46)	0.56	181.1	117.3	0.2
	(b)	-1.a2 (5.03)	1.26 (13.22)	0.62	174	92.3	0.2
(12,3)	(a)	-3.10 (2.51)	1.09 (10.35)	0.44	107	131.9	3.6
	(b)	-5.42 (4.35)	1.21 (11.63)	0.57	135	87.3	0.7
(12,6)	(a)	0.22 (0.24)	0.79 (6.66)	0.24	44	135.9	19.9
	(b)	-3.68 (4.28)	1.19 (11.27)	0.56	127	87.0	0.7
				CANADA			
(24,3)	(a)	13.96 (11.35)	0.27 (3.12)	0.09	9.7	88.6	5.2
	(b)	-2.69 (1.09)	0.68 (6.25)	0.30	39.1	84.2	1.4
(48,3)	(a)	29.6 (13.7)	0.36 (5.73)	0.26	32.8	89.0	17.2
	(b)	13.44 (2.23)	0.20 (2.02)	0.05	4.07	60.5	3.0
(48,241	(a)	14.15 (6.46)	0.50 (4.62)	0.18	21.4	85.6	13.2
	(b)	12.60 (5.60)	-0.04 (0.56)	-0.01	0.3	56.1	3.2

$\chi^2_B(1)$: χ^2_A

Annex 3

DESCRIPTION OF DATA

All interest rate data were obtained from the Directorate for Financial, Fiscal and Enterprise Affairs of the OECD while the source for data on inflation (based on the GNP/GDP deflator) and output is the Economics and Statistics Department of the OECD.

Some of the internally available OECD interest rate data are not ideal for the purposes for which they are being employed. The severity of these data limitations vary from country to country. From the United States, at one extreme, where four reasonably good time series are available to Japan, at the other extreme, for which sufficiently good interest rate data were not available to do even a single test. A general problem which characterises the data for most countries is the following: the reported monthly interest rate are, for the most part, monthly averages of daily data. For conformity with price and output data, end-month figures may be preferable. However, as noted by Mishkin (1988), the appropriate dating for the CPI in a particular month is not clear (at least for the United States) since price quotations on the component items of the index are collected at different times during the month.

United States: The following four interest rate series are available for the United States, three of which are used in the tests performed: yields on 3- and 6-month treasury bills and on 5- and 10-year-to-maturity U.S. government bonds. The treasury bill yields are monthly averages computed from closing-bid quotations reported daily. Yields are calculated on a bank discount rate basis and, therefore, only approximate a true yield. The government bond yields are also monthly averages of daily data for 5- and 10-year notes and bonds on the secondary market.

Germany: Given data limitations tests were only possible by ignoring the private-public sector division. Thus the results reported for Germany may be affected by different perceptions of default risk on public and private sector bonds. Three-month Fibor (Frankfurt interbank offered rates) are available for the full-time period of interest. However, 6-month Fibor data are only available from September of 1985 and the time series for 3-month treasury bills ends in November 1981. Thus the only short maturity yield available is that for the 3-month Fibor which is a private sector economy yield. The yields on public sector bonds on the secondary market with (3 – 7) and (7 – 15) years maturity are the only long bond yields available. The yields for the type of government securities mentioned above are weighted by the outstanding amount of bonds included in the calculation. Yields are generally monthly averages of daily data. However, up to January 1986 the yields on government bonds outstanding were based on the yield on four bank week return dates in each month. A private sector industrial bond yield is available but the maturity to which it refers is unknown.

France: The yields used here relate to private sector assets. 1-, 3-, and 6-month PIBOR (Paris interbank offer rates) are used. These are again monthly average rates. The yield on private sector bonds on the secondary market is also considered. This has a term to maturity of at least seven years. It is not used in the empirical tests herein, however.

United Kingdom: Three interest rate series are employed in tests for the United Kingdom. The 91-day treasury bill rate is, until August 1977, the weighted average rate of discount on allotment for 91-day bills at the weekly tender on the last Friday of the month. Since then, the rates are monthly averages of weekly data. The government security yields are gross redemption yields for selected maturities, derived from fitting observed yield maturity curves to a mathematical model of the government bond markets. Taxes are ignored. The figures are averages of Wednesdays until February 1980; from March 1980 until December 1981, figures are the average of all observations (3 per week): from January 1982 figures are averages of all working days.

Canada: The 3-month treasury bill rate is a weighted average of the yields on successful bids for three-month treasury bills sold by tender the last Thursday of the month. As from April 1981, the data are monthly averages of the Thursday rates. The government bond yields on the secondary market employed are unweighted averages of yields for issues other than guaranteed issues. The quotes used are mid-points between bid and asked prices at the close of business on the last Wednesday of the month. As from April 1981, the data are monthly averages of Wednesday rates.

Italy: The **3-, 6-** and 12-month treasury bill yields employed in the tests are end-of-month rates.

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