Expectations, long-term interest rates and monetary policy in the United Kingdom

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During the last few years the monetary authorities have expressed a willingness to allow interest rates to fluctuate more widely than in the past. As was indicated on numerous occasions, this shift to a more flexible interest rate policy represented a change in the Bank of England's market tactics rather than a change in its basic objectives.[1] Apart from the needs of government finance the Bank's main aim is to achieve the degree of monetary restraint judged to be appropriate to the economic situation and the overall direction of policy. Its intention in moving to greater flexibility in its policy on interest rates has been to allow market forces to be more fully and more quickly reflected in prices ([**2**], page 456).

The purpose of this study is to attempt to identify the 'market forces' that have determined the short-run (*i.e.* the month-to-month) movements in long-term interest rates in the United Kingdom. In the past, most econometric investigations of long-term bond rates have been divided into two parts. In the first part, emphasis is placed on analysing the term structure relations between the bond rate and short-term rate, while in the second part the objective has been to explain the movements in the short rate. An important assumption underlying this approach is that the current level of long-term interest rates is critically dependent upon expected future rates and that the latter may be explained by a weighted average of observed past rates. Section 1 of the paper provides a critical review of the evidence that has been presented in support of this view. Considering the extensive use of the term structure analysis in both the United Kingdom and the United States, such a review seems well worthwhile.

Econometricians are, of course, not the only ones who attribute an important role to expectations of future rates as a determinant of the demand for, and hence the current prices and yields on, long-term government securities. As R. S. Sayers puts it, "Public opinion about the future course of interest rates is a very powerful factor in the long-term market, and a mere reduction of short-term rates may not do much to modify that opinion; and unless opinion is modified little impression will be made upon long-term rates even if the authorities are prepared to operate in the market on a very large scale." ([**28**], page 142). There are, however, substantial differences between the views of econometricians on the determinants of expectations and those of other observers. For example, in several issues of this *Bulletin* it has been argued that investors' expectations are such that a fall in the price of government securities may result in a decline rather than an increase in the quantity

1 The first mention of such a change in the Bank's tactics appeared in the March 1969 issue of the *Bulletin* [**3**]. A subsequent statement of the Bank's views on monetary and debt management policy was presented in the Jane Hodge Memorial Lecture delivered by the Governor at the University of Wales, Institute of Science and Technology, on 7th December 1970 [**4**]. Further changes in the Bank's operations were announced in the consultative document *Competition and credit control*, 14th May 1971, but this study was completed before that date.

demanded in the immediate future. Such an argument implies that investors sometimes have a tendency to extrapolate recent changes in interest rates into the relatively near future. It has also been argued, however, that interest rate expectations are frequently highly volatile and unpredictable, and for this reason "the volume of purchases [of government securities] by the public tends to vary erratically" ([5], page 365).[1]

These and other questions are taken up in Section 2. The main purpose of that section is to identify the factors, other than expectations generated from past experience, that have influenced the month-to-month movements in the U.K. long rate. Among the variables considered are the level of economic activity, the quantity of money, the expected rate of inflation and foreign interest rates. The resulting model is then used to test for changes arising from the modification in the Bank of England's market tactics in the period up to 1970. A concluding section summarises the results and discusses some of their more general implications.

1 The distributed lag theory of interest rate expectations[2]

In recent years, the starting point for many empirical studies of long-term interest rates has been the expectations theory of the yield curve. According to this theory the expected returns from all bonds, regardless of term to maturity, will be identical over any given interval of time where the return is defined as the sum of cash payments plus any increase (or minus any decrease) in the market value of the bond.[3] In the absence of transactions costs and different preferences for interest income and capital gains this implies that

$$r_L(t) + g^e(t) = r_S(t) \qquad (1)$$

where $r_L(t)$: the ruling long-term rate

$r_S(t)$: the ruling short rate, which is not expected to change over the holding period

$g^e(t)$: the expected capital gain (or loss) on long-term bonds over the holding period[4]

In addition if $g^e(t)$ is taken as inversely proportional to the expected change in the long rate $[\Delta r_L{}^e(t)]$, equation (1) may be written as

$$r_L(t) = \alpha + r_S(t) + \beta\Delta r_L{}^e(t) \qquad (2)$$

where α and β are constants. Thus, to a first approximation the current long-term rate may be expressed as a linear function of the current short rate and the expected change in the long rate. The only remaining problem is the formulation of a theory to explain $\Delta r_L{}^e(t)$.

1 A more complete discussion of the Bank's view of the formation of interest rate expectations may be found in Goodhart [12] and Rowan and O'Brien [26].

2 This section is somewhat technical and to a large extent independent of Section 2. Readers who are unfamiliar with the literature in the area may wish to proceed directly to Section 2.

3 Following Hicks [16] the theory may be modified to allow for a less than perfect degree of substitutability between different maturity bonds due to the greater risk associated with the capital values of longer-term securities. As a result, holding period yields may differ by an amount of a liquidity premium paid to compensate for this risk. However, in view of the difficulties that have been experienced in attempting to measure such premiums (see, for example, Cagan [6] and Modigliani and Sutch [19]) they have been omitted.

4 Strictly speaking, equation (1) requires either that the holding period is equal to the maturity period of short-term debt or that the expected short rate is equal to the ruling short rate. If neither of these conditions is fulfilled, then $g^e(t)$ must be viewed as the difference between the expected capital gain (or loss) on long and short-term debt.

Following de Leeuw [**7**] most investigators assume a systematic relationship between past movements in interest rates and investors' expectations of the future.[1] De Leeuw's procedure provides a means of synthesising two widely held views concerning the formation of expectations. The first, the Keynesian [**17**] normal-rate hypothesis, assumes that on the basis of past experience investors have in mind a normal level of long-term interest rates, towards which current rates are expected to move. When current rates are higher than normal investors expect interest rates to fall, and vice versa. However, as Duesenberry points out, ". . . on a priori grounds there is no reason why the argument should not be turned just the other way . . . It would not . . . be surprising if it turned out that a rise in rates led to an expectation of a further rise and vice versa." ([**8**], page 318).

Since both of these hypotheses may be expressed in terms of the difference between the current long-term rate and some weighted average of past rates we have

$$\Delta r_L^e(t) = \gamma[r_L(t) - \sum_{i=1}^{m} \delta_i r_L(t-i)] \qquad (3)$$

where γ is a constant, m denotes the number of periods that investors look back in forming their expectations about the future and the δ_i sum to one and denote the weight attached to the rate in each past period.[2] Substituting equation (3) into (2) and solving for $r_L(t)$ yields equation (4), which has served as the basis for most applications of this approach.

$$r_L(t) = \frac{\alpha}{1-\beta'} + \frac{1}{1-\beta'} r_S(t) - \frac{\beta'}{1-\beta'} \sum_{i=1}^{m} \delta_i r_L(t-i) \qquad (4)$$

where $\beta' = \gamma\beta$

In two recent articles Rowan and O'Brien [**25**, **26**] use this type of analysis in an attempt to explain the behaviour of the long-term rate of interest in the United Kingdom.[3] The conclusions they reach are quite modest; nevertheless they argue that the approach is a promising one and ought to be pursued by other researchers. After examining a number of ways of expressing the distributed lag $[\sum_{i=1}^{m} \delta_i r_L(t-i)]$ Rowan and O'Brien find that the most satisfactory results are obtained when the δ_i are described by an exponential decay function of the form:[4]

$$\delta_i = \lambda(1-\lambda)^{i-1} \quad ; \quad i=1, \ldots, m. \qquad (5)$$

Substituting this expression into equation (4) and using the Koyck [**18**] transformation to eliminate the distributed lag

1 Among others see Ball [**1**], Gibbs [**11**], Modigliani and Sutch [**19**, **20**], Norton [**22**] and Pierson [**24**].

2 Since the term in the summation $\sum_{i=1}^{m} \delta_i r_L(t-i)$ represents the sum of two distributed lags there is no reason to expect it to be of a simple geometric form.

3 The following discussion does not provide a complete review of the Rowan and O'Brien model. In particular, nothing is said of their very interesting treatment of portfolio allocation under conditions of uncertainty and the effects of the maturity composition of government debt on the term structure of interest rates.

4 Commenting on more complicated lag structures such as the fourth degree polynomial employed by Modigliani and Sutch [**19**, **20**], the authors state that "In general the estimates of the weights did not exclude their representation by a single exponential function . . ." ([**25**], page 289) similar to equation (5); and that although ". . . the reduced forms of these investigators perform impressively, the values they imply for the structural parameters are not always acceptable, which suggests that the equations they estimate are essentially forecasting devices." ([**25**], page 310).

yields, after the rearrangement of terms and the addition of a random error term,

$$r_L(t)=\frac{\alpha\lambda}{1-\beta'}+\frac{1}{1-\beta'}r_S(t)-\frac{1-\lambda}{1-\beta'}r_S(t-1)+\frac{1-\lambda-\beta'}{1-\beta'}r_L(t-1)+u(t) \tag{6}$$

or more simply

$$r_L(t)=B_0+B_1r_S(t)+B_2r_S(t-1)+B_3r_L(t-1)+u(t) \tag{6a}$$

where

$$B_0=\frac{\alpha\lambda}{1-\beta'}\ ;\ B_1=\frac{1}{1-\beta'}\ ;\ B_2=-\frac{1-\lambda}{1-\beta'}\ ;\ B_3=\frac{1-\lambda-\beta'}{1-\beta'}$$

Hence, if the distributed lag theory of interest rate expectations holds

$$B_3=1-B_1-B_2 \tag{7}$$

which implies that equation (6a) may be rewritten as

$$r_L(t)-r_L(t-1)=B_0+B_1[r_S(t)-r_L(t-1)]+B_2[r_S(t-1)-r_L(t-1)]+u(t) \tag{6b}$$

To test their model Rowan and O'Brien compare the explanatory powers (*i.e.* the standard errors of the residuals) of equations (6a) and (6b). The argument is that if the variance explained by (6a) is not significantly greater than that explained by (6b), constraint (7) is satisfied and the model is not rejected by data. Table A presents the results obtained when the equations are fitted to monthly observations for several overlapping time periods, with $r_L(t)$ defined as the 2½% Consol rate and $r_S(t)$ defined as the discount rate on 91-day Treasury bills.[1] Inspection of the table reveals that the parameters of equation (6a) conform very closely to the conditions imposed by equation (7). Thus it is not surprising that the estimates of the standard errors of the residuals for equation (6b) are never significantly greater than those for equation (6a).[2] However, as Rowan and O'Brien point out, these results represent a necessary but not a sufficient condition for the acceptance of their model.

Another essential feature of the distributed lag theory of interest rate expectations [see equation (5)] is that λ does not equal zero, or, equivalently,

$$B_1 \neq -B_2$$

A test of this hypothesis is provided by comparing the fit of equation (6b) with that of the first difference relationship:

$$r_L(t)-r_L(t-1)=B_0+B_1[r_S(t)-r_S(t-1)]+u(t) \tag{6c}$$

which assumes $B_1=-B_2$. If the Rowan and O'Brien model is not to be rejected by the data, equation (6b) must fit the data significantly better than (6c). The parameter estimates and other relevant statistics for the latter equation are reported in Table B. The estimates of $\bar{R}^2$ indicate a signifi-

1 The time periods are the ones used by Rowan and O'Brien and the specification of the variables is quite similar to theirs. Nevertheless, there are some minor differences in the estimates for the equations fitted to the data after 1958. See the Appendix for precise definitions of the variables and their sources.

2 The nature of constraint (7) suggests that the appropriate statistic to be used in evaluating the difference in the standard errors of equations (6a) and (6b) is the likelihood ratio. The estimates of the ratio, none of which are significant at the 10% level, are similar to those reported by Rowan and O'Brien [26].

cant but fairly weak association between monthly first differences in the rates on short and long-term government securities since the early 1950s. This suggests some substitution in the demand for these securities and hence is consistent with the expectations theory of the yield curve. On the other hand, there appears to be no evidence at all to support the distributed lag theory of the formation of expectations. For each of the time periods considered the estimate of the standard error of the residuals for the first difference relationship (6c) is *less* than the estimate for equation (6b). As noted above this implies that

$$\lambda=\delta_1=\delta_2= \ . \ . \ . =\delta_m=0$$

that is, past interest rates (taken by themselves) have relatively little influence on investors' expectations of the near future.

One possible limitation of the analysis is that in following Rowan and O'Brien the most appropriate measures of the variables may not have been used, particularly in the case of the short-term interest rate. An alternative measure that has been suggested by a number of writers is the rate on three-month local authority debt. In their study of the demand for money in the United Kingdom, Goodhart and Crockett use the latter variable instead of the Treasury bill rate ". . . on the grounds that in recent years the local authority market has attracted a wider range of active participants and has been less dominated by the direct influence of the authorities than has the Treasury bill market." ([**13**], page 191.) To test the hypothesis that during the 1960s the local authority rate was a better indicator of the short-term rate of interest than the Treasury bill rate, equations (6a), (6b) and (6c) were estimated for the period February 1960–December 1969 with $r_s(t)$ defined as the yield on three-month local authority deposits.

In all important respects the results (see Table C) are identical to those presented above. The estimate of $\bar{R}^2$ for equation (6c), the first difference relationship, is very low. However, once again the statistical fit of this equation (as measured by the standard error of the residuals) is as good as or better than that provided by either of the other equations. Thus, no matter which measure of the short rate is used, it is not possible to find any support for the hypothesis that there is a systematic relationship between past movements in rates and investors' expectations of the future. Moreover, the similarity in the behaviour of the yields on Treasury bills and local authority debt during the latter part of the 1960s suggests that in recent years the monetary authorities have exerted considerable influence over both of these variables. Hence, one should not expect them to perform very differently in most econometric analyses of this period. The simple correlations between monthly first differences in the Treasury bill rate and the local authority rate for several recent periods are given below:

Time period	Correlation
February 1960 – December 1964	0·455
January 1965 – December 1969	0·739
January 1965 – October 1967	0·766
November 1967 – December 1969	0·703

Another question regarding our findings concerns the differences between them and the results that have been obtained for the United States. Modigliani and Sutch [**19**, **20**] are generally considered to have presented an impressive case for the distributed lag theory of interest rate expectations. However, in their most recent paper [**21**] they report that after allowance is made for serial correlation (something they had not done previously), the distributed lag term contributes only marginally to the explanatory power of their model. The summary statistics for their regressions, which are used to explain the rate on long-term U.S. government bonds from the second quarter of 1953 to the fourth quarter of 1966, including and excluding the distributed lag term are:

	R^2	Standard error of estimate	Durbin-Watson statistic
Including the lag	0·979	0·091	1·93
Excluding the lag	0·976	0·093	2·00

It is easily verified that since the distributed lag introduces five parameters into the equation, the small increase in R^2 is not statistically significant even at the 20% confidence level.[1] Thus, on closer examination, the results for both the United States and the United Kingdom indicate that no-one has yet succeeded in developing a reliable statistical model relating past interest rates to expected future rates. The problem would appear to be that in the real world there are so many factors other than the history of interest rates that influence investors' expectations of the future that it is difficult to isolate a stable relationship between the latter variables.[2] On the other hand, it may be argued that the exponentially distributed lag is too simple and that one should try more complicated distributed lag functions. However, in view of the very weak relationship between month-to-month movements in short and long-term interest rates in the United Kingdom, such an approach does not seem promising.[3]

2 An alternative approach to the explanation of the U.K. long rate

The poor performance of the term structure analysis suggests the need for an alternative explanation of the behaviour of long-term interest rates. The one adopted here may be viewed as an application of the reduced form approach to interest rate estimation, which has been employed in a number of recent studies.[4] Among the

1 Four parameters are used to describe the shape of the distributed lag (*i.e.* the degree of the polynomial) and one is needed to specify the length of the lag. Even if the latter parameter is not considered, the contribution of the distributed lag on past rates is not significant at the 10% level. For a more thorough analysis of the U.S. data see Hamburger and Latta [**14**]. Other investigators who have reported difficulties in their attempts to express expected future rates as a function of observed past rates are Parkin [**23**] and White [**30**].

2 A similar view has recently been expressed by Duesenberry [**9**] with respect to price expectations.

3 The correlations between monthly first differences in short and long-term interest rates in the United Kingdom are considerably lower than those that have been observed in the United States. The estimates of $\bar{R}^2$ for equations relating monthly changes in the yield on long-term U.S. government bonds to monthly changes in the three-month U.S. Treasury bill rate, for the period 1920-65 and nine shorter periods, are generally around 0·25. Perhaps equally important, though, is the stability of the estimated relationship. In only one case (the 1934-38 period) do the regression coefficients differ significantly from those obtained when all the data are pooled. See Hamburger and Latta [**14**].

4 See Ball [**1**], Feldstein and Eckstein [**10**], Hamburger and Silber [**15**], Pierson [**24**] and Walters [**29**]. A somewhat different reduced form model is presented by Sargent [**27**].

variables suggested as determinants of interest rates in these studies are: the level of income, the quantity of money, the expected rate of inflation and the level of interest rates in previous periods. Considering the importance of London as an international financial centre and the short-run nature of this study, it was decided to include foreign interest rates and the forward exchange rate as additional determinants of U.K. rates. The latter variable may be viewed either as the explicit cost of covering investments in sterling or as some general indicator of confidence. In both cases a fall in the value of sterling in the forward exchange markets would be expected to reduce the demand for U.K. securities and put upward pressure on interest rates.[1] Measures of all the variables mentioned above were included in regression equations used to explain monthly first differences in the Consol rate for the period 1965-69. Thus far, the only variables for which we have been able to identify clear and sustained influences on the Consol rate are the three-month euro-dollar rate, the three-month forward discount (or premium) on the pound and the lagged value of the Consol rate.[2]

First differences are used to mitigate the serial correlation problem and to focus the power of the analysis on the problem of most concern to policy-makers, explaining the changes in rates from one period to the next. One of the disadvantages of using first differences is that it tends to emphasise the short-run determinants of rates, such as foreign interest rates, relative to the longer run and perhaps more fundamental determinants *e.g.* the level of economic activity and the expected rate of inflation.[3]

To test the stability of the estimated relationships the sample period is divided into pre- and post-devaluation sub-periods: January 1965 – October 1967 and November 1967 – December 1969. As only the second of these includes the period after the Bank of England adopted a more flexible policy on interest rates, it is also possible to test for the effects of this change.[4] Finally, the behaviour of the Consol rate in 1970 and 1971 is examined to determine how well the relationship has held outside the sample period.

The U.K. long rate and the euro-dollar rate Regression results for the two sub-periods are presented in Table D. Several aspects of the results warrant comment. Note first the relatively high coefficients of determination and the stability of the relationship between month-to-month move-

1 Some problems might be encountered in identifying the direction of the flow of causality between the forward exchange rate and U.K. rates in any current period. However, as will be seen below, this problem creates no serious difficulties in the present analysis.

2 The Consol rate is used as the measure of the U.K. long rate, first because it is the rate that has been used most often in other studies and secondly because it is published on a monthly average basis (this is not true of the War Loan rate).

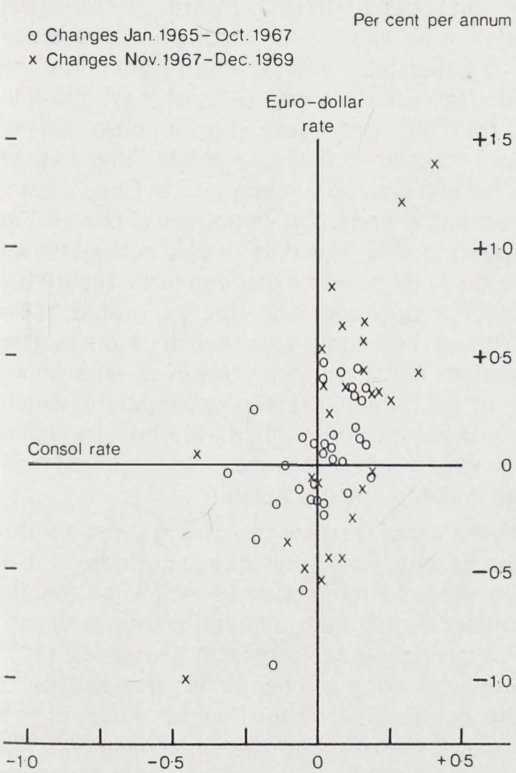
3 The effects of the latter variables on U.S. rates are discussed by Feldstein and Eckstein [**10**], Sargent [**27**] and Yohe and Karnosky [**31**]. The equations used in these studies were fitted to quarterly and annual levels of the variables and not monthly first differences. Measurement problems have presumably also contributed to the difficulties we have experienced in obtaining statistically significant coefficients for some variables. Income is a particularly difficult variable to measure for periods of less than a quarter. The variable used in this study was the industrial production index, but as is well known it has some serious shortcomings.

4 A closer matching of the time periods with the 1968 change in tactics would greatly reduce the number of degrees of freedom in the later period. As an indication of the change in policy, we note that during the second period the standard deviation of the monthly first differences in the Consol rate was approximately 60% greater than it was during the first period.

Chart A

The euro-dollar rate and Consol rate, 1965-69

Monthly first differences



ments in the euro-dollar rate and the Consol rate.[1] Compare the estimate of the euro-dollar rate coefficient in equation D-1 with the one in equations D-2 and D-3. In addition, as Chart A indicates, the relationship is not dependent upon a few extreme observations. Such an association between the euro-dollar rate and the Consol rate is not very surprising; what is surprising is that it does not seem to work through U.K. short rates. If it did, one would expect U.K. short rates to provide a better 'explanation' of U.K. long rates than the euro-dollar rate, but this does not appear to be the case. The estimates of $\bar{R}^2$ for equations D-1, D-2 and D-3, when the local authority rate (r_{LA}) and the Treasury bill rate (r_{TB}) are substituted for the euro-dollar rate, are:

Equation	r_{LA}	r_{TB}
D-1	0·1831	0·2044
D-2	0·1728	0·0484
D-3	0·2643	0·2691

Moreover, in the period after devaluation the simple correlation of the euro-dollar rate with the medium and longer-term gilt-edged rates was higher than that of the three-month local authority rate or of the Treasury bill rate, though shorter and longer-dated gilt-edged stocks are generally more highly correlated than are the euro-dollar rate and any gilt-edged rate at all. This is demonstrated in the following table of simple correlations.

	Euro-dollar	Treasury bill	Local authority	British government stocks: 5-year	10-year	2½% Consols
	(Nov. 1967-Dec. 1969)					
Euro-dollar	1·00					
Treasury bill	0·21	1·00				
Local authority	0·40	0·70	1·00			
British government stocks:						
5-year	0·47	0·63	0·64	1·00		
10-year	0·63	0·48	0·57	0·84	1·00	
2½% Consols	0·63	0·29	0·45	0·63	0·92	1·00

It is particularly interesting to see that whereas the correlations between domestic rates decline as the differences in their maturities increase, those for the euro-dollar rate move in the opposite direction.

A large portion of these paradoxical results can, no doubt, be attributed to the control that the monetary authorities have exerted over U.K. short rates. Nevertheless, some discussion of the process linking the rates on euro-dollar deposits and long-term British government securities does seem appropriate. One possibility is that there is some degree of substitution in the demand for these securities. Alternatively, it may be argued that the linkage is purely expectational *i.e.* changes in the euro-dollar rate have a major impact on investors' expectations of the future trend in U.K. long rates. This line of reasoning does not imply much actual substitution between euro-dollars and long-term U.K. securities. Indeed, the statistical information on the ownership of gilt-edged securities, which is quite com-

1 The constant term is included in the equations to permit identification of any trend in the dependent variable that is not explained by the independent variables.

prehensive, suggests hardly any movement of overseas funds into or out of these securities over the periods considered. However, such direct substitution is not required for the first explanation put forward. Changes in the euro-dollar rate could put downward (or upward) pressure on U.K. long rates simply by causing U.K. companies to move out of (or into) the domestic capital market and into (or out of) the euro-dollar market. During 1970, at least, there is some indication that this did occur. Furthermore, if the linkage were purely expectational one might also expect U.S. long-term rates to have some effect on the Consol rate. The evidence does not support this hypothesis. The partial correlation between monthly first differences in the Consol rate and the rate on U.S. government long-term securities, holding the euro-dollar rate constant (for the period 1965-69), is $-0{\cdot}03$. Without additional evidence it is difficult to pursue the argument much further. However, one thing does seem clear: to the extent that it is expectations which provide the link between the euro-dollar rate and the long-term U.K. rate, it would seem that this expectational response appears quite stable and predictable.

Monetary policy and other factors Turning next to the differences in the results for the two sub-periods, it becomes possible to examine the extent to which the results reflect the authorities' announced intention of allowing gilt-edged rates to react more fully to market pressures ([**2**], page 456). Evidence of such a change is provided by a comparison of the coefficients of the lagged value of the Consol rate in each period. The lagged value of the dependent variable is generally included in a regression to allow for an exponentially distributed lag in the response of the dependent variable to changes in the independent variables. On this interpretation, the results are consistent with the authorities' intentions. The coefficient of $\Delta r_L(t-1)$, which measures the percentage of the adjustment that is *not* completed during any time period, is statistically significant (at the 0·05 level) in equation D-1, but not in equations D-2 or D-3. This implies that the response of the Consol rate to changes in its determinants was significantly faster in the second period than it was in the first. Indeed, the evidence suggests that since devaluation the adjustment in the Consol rate has been completed within the period of observation – one month.[1]

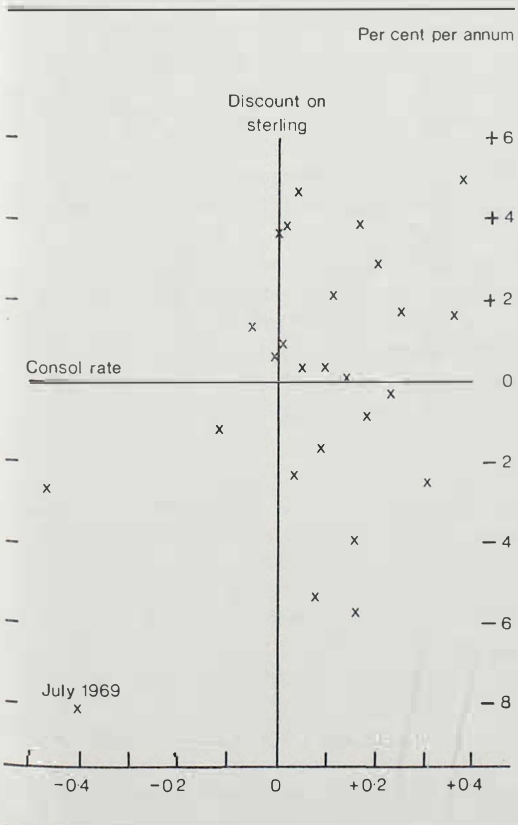
The differences in the estimates of $\bar{R}^2$, which measure the relative explanatory power of the equations, are also consistent with the hypothesis that in recent years official intervention in the gilt-edged market has been on a more limited scale than it was previously. The higher values in the second period suggest that since devaluation the relationship between the Consol rate and the euro-dollar rate has become more predictable. Compare the estimates of $\bar{R}^2$ for equations D-1, D-2 and D-2a (to be reported below). Thus, in terms of both speed and reliability, the evidence supports the view that during the past few years the Consol

1 When $\Delta r_L(t-1)$ is included in equations D-2 and D-3, the parameter estimates are negative but the absolute values of the t-statistics never exceed 0·5.

Chart B

The Consol rate and lagged changes in the three-month discount on sterling, Nov. 1967 – Dec. 1969

Monthly first differences



rate has become more responsive to market forces than it was before.[1]

Another apparent difference in the regression results for the two sub-periods concerns the relationship between the Consol rate and the value of sterling in the forward exchange markets. In the first period, the month-to-month movements in these variables appear to be almost totally unrelated. Neither the first nor the second differences in the three-month discount on the pound are statistically significant when they are added to equation D-1. In the post-devaluation period, the second difference, lagged one period, is significant well beyond the 0·05 level while the first difference falls a little short. This suggests that changes in the three-month discount on the pound first have a positive influence on the Consol rate in the current and subsequent month, but that the influence becomes significantly negative in the month thereafter. As a result, fluctuations in the forward value of sterling since devaluation have had little net effect on the Consol rate over a quarter. However, even this transitory effect is subject to question. As is indicated in Chart B, the relationship between $\Delta^2£D(t-1)$ and $\Delta r_L(t)$ is heavily dependent upon the inclusion of one observation, July 1969. If this observation is excluded from the sample, the simple correlation between these variables falls from 0·39 to 0·18 and neither $\Delta^2£D(t-1)$ nor $\Delta£D(t)$ is significant when it is included in equation D-3. In addition, the estimate of $\bar{R}^2$ for the equation containing only the change in the euro-dollar rate rises to 0·47. See equations D-2a and D-3a.[2]

$$\Delta r_L(t) = \underset{(0{\cdot}0459)}{0{\cdot}2148}\Delta r_{E\$}(t) + \underset{(0{\cdot}0265)}{0{\cdot}0534} \qquad \text{D-2a}$$

$$\bar{R}^2 = 0{\cdot}4651 \qquad \text{S.E.R.} = 0{\cdot}1230 \qquad \text{D.W.} = 1{\cdot}6338$$

$$\text{D-3a}$$

$$\Delta r_L(t) = \underset{(0{\cdot}0478)}{0{\cdot}1924}\Delta r_{E\$}(t) + \underset{(0{\cdot}0120)}{0{\cdot}0157}\Delta£D(t) + \underset{(0{\cdot}0083)}{0{\cdot}0102}\Delta^2£D(t-1) + \underset{(0{\cdot}0261)}{0{\cdot}0548}$$

$$\bar{R}^2 = 0{\cdot}4912 \qquad \text{S.E.R.} = 0{\cdot}1200 \qquad \text{D.W.} = 1{\cdot}6563$$

Note: Standard errors of the estimated coefficients are shown in brackets.

Consequently the only important difference we find in the model used to explain the behaviour of the Consol rate in the two sub-periods is the speed with which the rate has adjusted to changes in the euro-dollar rate. Other than this the same regression equation seems to fit the data reasonably well in both periods. This result suggests that the Bank of England's change of tactics has not affected the nature of the market's response to external stimuli. Fluctuations in rates were certainly wider in the second period than in the first, but there is no evidence to suggest that this impaired the functioning of the market.

The 1970-71 experience Finally, it is useful to consider the extent to which movements in the Consol rate in 1970 and 1971 can be explained by the behaviour of the euro-dollar rate. During the autumn of 1970 the performance of the

1 As will be seen, the relationship was not nearly so strong in 1970, so to that extent conclusions drawn from the difference in $\bar{R}^2$ between the two periods should be somewhat reserved.

2 If only the least significant variable in equation D-3a [*i.e.* $\Delta^2£D(t-1)$] is eliminated, the t-statistic for $\Delta£D(t)$ drops from 1·3 to 1·27.

model was not as good as might have been expected. In three months, September, November and December, the rates moved in opposite directions. However, taking the fourteen months January 1970 – February 1971 as a whole, the results are impressive. The simple correlation between monthly changes in the euro-dollar rate and the Consol rate is 0·49, which compares quite favourably with the negative correlations between the latter variable and the local authority rate (r_{LA}) and the Treasury bill rate (r_{TB}). The matrix of simple correlations among these variables is presented below:

Simple correlations among monthly changes in various interest rates January 1970 – February 1971

	r_L	$r_{E\$}$	r_{TB}	r_{LA}
r_L	1·00			
$r_{E\$}$	0·49	1·00		
r_{TB}	−0·30	−0·01	1·00	
r_{LA}	−0·45	−0·15	0·63	1·00

It is apparent that the equations estimated above do not provide a foolproof account of changes in the U.K. long rate. Nevertheless, the behaviour of the rate appears to be far less erratic than is often asserted. Furthermore, the 1970-71 results reinforce our earlier finding that in recent years:

(a) there has been a close correspondence between the U.K. Treasury bill rate and the three-month local authority rate; and

(b) monthly movements in these rates have been almost totally unrelated to movements in both the euro-dollar rate and the U.K. long rate.

The results also indicate that researchers ought to look for other variables, in addition to the history of interest rates, to help explain investors' expectations of the future.

3 Conclusion

It may be helpful at this point to summarise the major empirical results reported above and to discuss some of their implications. First, it was found that past interest rates do not provide a reliable indicator of expected future rates. This is not to say that investors ignore the past in forming their expectations about the future, but it does suggest that the relation between past and expected interest rates is not very stable.

Secondly, there appears to be only a very weak relationship between month-to-month movements in the yields on short and long-term British government securities. This contrasts quite sharply with the finding that there is a close and very stable relationship between the euro-dollar rate and the U.K. long rate. The available evidence is not sufficient to permit a choice between two alternative interpretations of the latter result: one, that the linkage is purely expectational and, two, that there is some substitution (direct or indirect) between short-term international claims and long-term government securities. To the extent that there is some substitution between these securities, the widely held view that it is long rates which affect the domestic economy and short rates which influence international capital flows would require some rethinking. It is worth repeating, however, that the dominant role attributed to the euro-dollar rate as a determinant of the U.K. long rate in this study may be very much influenced by the short-run nature of the analysis *i.e.* the use of monthly first differences.

Finally, the evidence suggests that in moving to greater flexibility in their policy on interest rates, the authorities have accomplished their objective of allowing market forces to be more fully reflected in the prices of gilt-edged securities. There is no indication, however, that this has impaired the functioning of the market in any way.

Appendix

Sources and definitions of variables

$2\frac{1}{2}$% Consol rate *Annual Abstract* of *Statistics*	Average of working days, based on the mean of the middle opening and middle closing prices each day, and ignoring tax considerations.
Treasury bill rate *Annual Abstract* of *Statistics*	Average rate of discount on allotment for 91-day Treasury bills at weekly tender.
Euro-dollar rate *Bank* of *England Quarterly Bulletin*	Average of Friday middle closing rates on three-month euro-dollar deposits.
Three-month discount on sterling *Bank* of *England Quarterly Bulletin*	Average of Friday middle closing prices. (Per cent per annum).
Local authority rate *Bank* of *England Quarterly Bulletin*	Average of the range of Friday rates on local authority temporary loans for a minimum term of three months and thereafter at seven days' notice.

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Table A

Estimates of equations 6a and 6b for several overlapping time periods

6a $r_L(t) = B_0 + B_1 r_S(t) + B_2 r_S(t-1) + B_3 r_L(t-1) + u(t)$

6b $r_L(t) - r_L(t-1) = B_0 + B_1 [r_S(t) - r_L(t-1)] + B_2 [r_S(t-1) - r_L(t-1)] + u(t)$

Data period	Equation	Estimated coefficients of: B_0	B_1	B_2	B_3	Coefficient of determination	Standard error of estimate	Durbin-Watson statistic
Jan. 1946 – Dec. 1967	**6a**	0·0310	0·1205 (0·0199)	−0·1178 (0·0199)	0·9945 (0·0104)	0·995	0·0901	1·4872
	6b	0·0144	0·1192 (0·0197)	−0·1185 (0·0198)		. .	0·0899	1·4909
Jan. 1946 – Dec. 1956	**6a**	0·0644	0·1026 (0·0363)	−0·0949 (0·0366)	0·9824 (0·0167)	0·988	0·0691	1·7489
	6b	0·0218	0·0981 (0·0358)	−0·0938 (0·0366)		. .	0·0690	1·7595
Jan. 1957 – Dec. 1967	**6a**	0·0856	0·1230 (0·0266)	−0·1289 (0·0266)	0·9926 (0·0185)	0·977	0·1074	1·3376
	6b	0·0082	0·1201 (0·0264)	−0·1264 (0·0265)		. .	0·1074	1·3442
Jan. 1946 – June 1952	**6a**	−0·0137	0·0523 (0·0623)	−0·0212 (0·0708)	1·0046 (0·0178)	0·984	0·0641	1·5198
	6b	0·0029	0·0857 (0·0589)	−0·0822 (0·0593)		. .	0·0647	1·4305
July 1952 – Dec. 1958	**6a**	0·1759	0·1331 (0·0303)	−0·1125 (0·0322)	0·9448 (0·0403)	0·973	0·0824	1·8786
	6b	0·0114	0·1327 (0·0304)	0·1250 (0·0306)		. .	0·0826	1·9240
Apr. 1959 – Mar. 1968	**6a**	0·1979	0·1161 (0·0316)	−0·1010 (0·0321)	0·9578 (0·0287)	0·970	0·1081	1·3024
	6b	0·0240	0·1114 (0·0316)	−0·1059 (0·0321)		. .	0·1087	1·3123

Note: Standard errors of the estimated coefficients are shown in brackets.
. . not available.

Table B
Estimates of equation 6c for several overlapping time periods

6c $\Delta r_L(t) = B_0 + B_1 \Delta r_S(t) + u(t)$

Data period	Estimated coefficients B_0	Estimated coefficients B_1	Coefficient of determination	Standard error of estimate	Durbin-Watson statistic
Jan. 1946 – Dec. 1967	0·0134	0·1189 (0·0195)	0·121	0·0898	1·4911
Jan. 1946 – Dec. 1956	0·0130	0·0992 (0·0357)	0·049	0·0689	1·7580
Jan. 1957 – Dec. 1967	0·0145	0·1232 (0·0257)	0·144	0·1071	1·3397
Jan. 1946 – June 1952	0·0193	0·0842 (0·0581)	0·014	0·0643	1·4339
July 1952 – Dec. 1958	0·0048	0·1294 (0·0300)	0·186	0·0824	1·9140
Apr. 1959 – Mar. 1968	0·0176	0·1091 (0·0308)	0·097	0·1083	1·3108

Note: Standard errors of the estimated coefficients are shown in brackets.

Table C

Estimates of equations 6a, 6b and 6c with $r_s(t)$ defined as the yield on three-month local authority deposits: Feb. 1960-Dec. 1969

Equation	Estimated coefficients B_0	B_1	B_2	B_3	Coefficient of determination	Standard error of estimate	Durbin-Watson statistic
6a	0·0573	0·1074 (0·0385)	−0·0950 (0·0385)	0·9846 (0·0240)	0·981	0·1359	1·4647
6b	0·0368	0·1063 (0·0380)	−0·0952 (0·0383)		. .	0·1354	1·4742
6c	0·0281	0·1022 (0·0377)			0·051	0·1354	1·4691

Note: Standard errors of the estimated coefficients are shown in brackets.
. . not available.

Table D

Regressions explaining the monthly first differences in the Consol rate

Data period	Equation	Estimated coefficients of: $\Delta r_L(t-1)$	$\Delta r_{E\$}(t)$	$\Delta£D(t)$	$\Delta^2£D(t-1)$	Constant	Coefficient of determination	Standard error of estimate	Durbin-Watson statistic
Jan. 1965 – Oct. 1967	1	0·4173 (0·1431)	0·1464 (0·0615)			0·0054 (0·0176)	0·3030	0·1008	1·8605
Nov. 1967 – Dec. 1969	2		0·2248 (0·0567)			0·0334 (0·0320)	0·3710	0·1520	2·0637
	3		0·1933 (0·0535)	0·0190 (0·0133)	0·0206 (0·0080)	0·0412 (0·0286)	0·5096	0·1342	1·8208

Note: Standard errors of the estimated coefficients are shown in brackets.

$r_L(t)$: 2½% Consol rate.

$r_{E\$}(t)$: three-month euro-dollar rate.

£D(t): three-month forward discount on the pound (per cent per annum).