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THE IMPACT OF MONETARY POLICY IN THE UK ON THE RELATIONSHIP BETWEEN THE TERM STRUCTURE OF INTEREST RATES AND FUTURE INFLATION

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The impact of monetary policy in the UK on the relationship between the term structure of interest rates and future inflation.

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Abstract

The link between London interbank interest rates and future inflation in the UK is investigated over a period which includes several changes in monetary policy regime. Recursive estimation is used to identify appropriate breakpoints in the sample and a moving-block bootstrap is used to facilitate correct inference. The general conclusion which emerges is that the informational content of the term structure is sensitive to changes in monetary policy. In particular, the relationship between interest rates and future inflation is found to break down after 1985.

Keywords

term structure, future inflation, monetary policy, recursive estimation.

JEL Classification: E52; C53.

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1 Introduction

Much recent research has focussed on the informational content of the term structure of interest rates. Specifically, the predictive power of yield spreads with respect to future inflation and real economic activity have been examined by numerous authors including Mishkin (1990b), Mishkin (1990a), Mishkin (1991), Estrella and Mishkin (1997) and more recently Siklos (2000) and Hamilton and D.H.Kim (2000). In general, it has been found that the term structure does contain useful information concerning these final goals of monetary policy, but that the relationship can change over time. In the context of European monetary policy, for example, Estrella and Mishkin (1997) make a strong argument for the use of interest-rate spreads as simple but accurate indicators of real economic activity and inflation. This paper subjects this view to a close scrutiny by examining the relationship between short-term, money-market spreads and future inflation in the UK for the period 1976 to 1999.

From the perspective of monetary policy it is crucial to assess whether the informational content of interest rates with respect to future inflation is relatively robust, even invariant, to the conduct of monetary policy. To facilitate this investigation, the interest rate data used in this paper are from the London interbank market which is a highly competitive market in sterling term deposits and is also closely integrated with the markets for other liquid assets such as Treasury bills and commercial paper. Shorter-term LIBOR interest rates[1], such as 1- and 3-month rates are influenced by the Bank of England as part of its monetary policy operations, but longer rates are determined in a highly competitive environment. It may be therefore

[1] LIBOR is an abbreviation of London Interbank Offer Rate. The interest rates in this data set are month-end middle rates.

that the informational content of the LIBOR interest rates will be sensitive to changes in the conduct of monetary policy. These would include, for example, the rapid interest rate increases following the election of the first Thatcher government in 1979-80, the sterling crisis of 1985 and the Deutschemark-shadowing experiment of 1987/88. Of particular interest is the change from a regime of exchange rate targeting to one of inflation targeting in 1992. One of the aims of the paper is to assess whether or not this change is reflected in the informational content of London interbank interest rates. The UK experience should, therefore, constitute a fertile testing ground for the impact of monetary shocks on the predictive power of the yield curve for changes in inflation.

There is another aspect of the relationship between interest rates and inflation which this paper seeks to address and that relates to the quality of the information which this approach identifies. This issue requires both the accurate identification of breaks in the relationship over time and valid assessment of the statistical reliability of the information during periods of relative stability. To facilitate this inquiry, two methodological innovations are introduced to supplement the basic methods employed in the earlier literature. The first of these relates to the detection of changes in the relationships over time. Rather than dividing the sample into sub-periods, on grounds of economic priors, a combination of recursive estimation and formal testing is used to identify the number of significant changes over the sample period. As will become apparent, this method yields some interesting results in terms of the UK data. The second innovation relates to the robustness of statistical inference in these kinds of studies, which are based on very simple dynamic models that are likely to be misspecified. Rather than merely relying on corrections to the standard errors of the estimated parameters by means of the asymptotic covariance correction of Newey and West (1987), standard errors and actual confidence intervals for the distribution of the relevant coefficient are generated by means of a moving-block bootstrap procedure (Carlstein, 1986; Künsch, 1989; Liu and Singh, 1992).

The structure of the paper is as follows. Section 2 outlines the basic method for estimating the relationship between future inflation and the term-structure of interest rates. Section 3 presents three sets of empirical results for this model, namely, the full-sample estimates, the results of the recursive estimation which is used to identify the appropriate number of sub-periods over which to estimate the model and then the sub-sample estimates. A discussion of the moving-block bootstrap procedure and the results from its implementation are contained in Section 4. Section 5 is a brief conclusion.

2 The inflation-change equation

The standard approach to analysing the information in the term structure regarding future inflation is based on the Fisher equation

$$i_t^m = \left(E\left[r_t^m\right]\right) + E\left[\pi_t^m\right]$$

which expresses the nominal interest rate, i_t^m, in terms of the expected real interest rate, $E\left[r_t^m\right]$, and the expected inflation, $E\left[\pi_t^m\right]$, over the relevant period. On the assumption of rational expectations, the forecast and actual inflation will differ by some error term, say ϵ_t^m, which is uncorrelated with any information at time t. The difference between inflation over the next m-periods and inflation over the next n-periods can then be written in the form of the so-called "inflation-change equation" proposed by Mishkin (1990b,a):

$$\pi_t^m - \pi_t^n = \alpha_{m,n} + \beta_{m,n}\left[i_t^m - i_t^n\right] + e_t^{m,n} \tag{1}$$

where

$$\begin{aligned} \alpha_{m,n} &= -\left(E\left[r_t^m\right] - E\left[r_t^n\right]\right) \\ e_t^{m,n} &= \epsilon_t^m - \epsilon_t^n \end{aligned}$$

and α and β are parameters to be estimated.

This equation purports to explain the inflation differential over a given time horizon in terms of an equivalent yield differential between two financial instruments on the assumption that the yield differential on real interest rates, over the time period concerned is constant. As a model of changes in inflation, this regression is bound to be misspecified. The argument is, however, that even though the model is misspecified, the effect of the spread on future inflation changes will be estimated consistently so long as $e_t^{m,n}$ has mean zero and the slope of the term structure of real interest rates, $\alpha_{m,n}$, is constant. The literature concludes that the slope of the term structure contains information about changes in future inflation if the hypothesis $\beta = 0$ is rejected.

Of particular importance will be the behaviour of the estimate of β at the time of entry and exit from the exchange rate mechanism of the European community. In principle, when going from exchange rate targeting to inflation targeting, one would expect the predictive power of the term structure to disappear—since if monetary policy was successful, the change in inflation ex post should be zero. So if the hypothesis $\beta = 0$ can be rejected using the sample before inflation targeting and $\beta = 0$ cannot be rejected over the sub-sample after inflation targeting, this should imply that inflation targeting has been successful.

It is clear from the above remarks that correct inference in relation to the coefficient β is the centrepiece of this method. This raises two important problems whose nature and suggested resolution are now discussed.

The first of these relates to changes in regime. In addition to the potentially important regime change in the early 1990s, there were a number of other unanticipated monetary events in the UK during the data-sample period which may have influenced the informational content of the term structure. The question of whether or not the relationship in equation (1) is constant over the entire sample period is therefore an empirical one of crucial importance. The empirical evidence on the importance of structural stability in this relationship in other countries has only recently started to emerge. For example, Estrella et al. (2000) and Schich (1999) have recently investigated the temporal evolution of the informational content of the yield curve in the Germany and the United States. The results seem to suggest a single breakpoint for US (in late 1979) but a relatively stable relationship for Germany. It may be therefore that the relationship is not influenced by minor unanticipated events but only by significant regime changes.

This question is not adequately addressed by constructing a series of sub-samples for estimating the value of β. In the first instance, the sub-samples would impose prior beliefs of the significance of particular events rather than allowing the data to determine the appropriate break points. In addition if all possible monetary shocks were to be catered for, many of the required sub-samples would be too small to allow reliable inference, especially when appeal is made to asymptotic results to justify hypothesis testing. This, in turn, would detract from investigating the major issue of the effect of the the adoption of inflation targeting. To investigate the temporal evolution of the term structure a better approach it that of recursive estimation to allow the data to indicate the significant breaks. Even this approach is problematic as the recursive coefficients will be biased after a structural break, since they also reflect the history of the previous regime, and converging to the OLS estimates for the entire sample. One remedy is to employ a backward recursive scheme to reflect the information content under the most recent regime and thus complement the information gleaned in the forward-recursive estimates. In addition, recent tests for structural change with unknown breakpoints following Andrews (1993) and Andrews and Ploberger (1994), and more traditional tests for structural change are used to formalise the information gleaned from the recursive estimation.

The second problem concerns statistical inference on the parameter β. Even within regimes, where stable relationships could exist, the standard error associated with the estimate of β is likely to be biased. Residuals from the inflation-change equation will probably be autocorrelated — because of equation misspecification and overlapping observations — and perhaps heteroskedastic. The standard solution has been to correct the standard errors, using the method of Newey and West (1987), which corrects standard errors for unknown forms of autocorrelation and heteroskedasticity. Whether this correction will be sufficient, so appeal can be made to standard distributions for inference, remains an open question. Our solution is to construct bootstrap distributions of the estimator. The construction of these distributions take into account the temporal characteristics of the residuals and enable us to assess the underlying uncertainty of the results by examining the confidence intervals within sub-samples.

3 The data and empirical results

The data are continuously compounded interest rates and inflation rates. The interest rates used are monthly LIBOR data for 1-, 3-, 6- and 12-month deposits for the period 1976:1 1999:09. Although we will report results for spreads of all combinations of maturities, we will in the following be focussing on the data involving 12-month rates. The upper panel of Figure 1 illustrates that the variance of the inflation spreads decreases later in the sample, particularly after joining the exchange rate mechanism of the European Community in on 8 October 1990. Although membership was suspended on 6 September 1992 a new framework for inflation targeting was announced in October of that year and the first Inflation Report by the Bank of England was published in February 1993. It appears therefore that volatility of inflation rates in the inflation-targeting regime has been more muted than in earlier monetary regimes. A second observation which may be made about the behaviour of inflation spreads concerns the regular pattern of spikes in the data. This indicates the likely presence of autocorrelation in the data. Given the very simple dynamic structure of the inflation-change equation, therefore, it is likely that correcting for both heteroskedasticity and autocorrelation will be an important factor in assessing the statistical significance of the information about future inflation contained in interest-rate spreads.

[Figure 1 about here.]

The corresponding interest rate spreads are plotted in the lower panel of Figure 1. As expected the most volatile of the three variables is the 12- and 1-month LIBOR spread. This is to be expected given the likely sensitivity of 1- month LIBOR interest rates to the short-term liquidity conditions in the interbank market and the influence of the Bank of England's monetary operations at the short end of the maturity spectrum. Although armed with institutional knowledge one might claim that a change occurs in the behaviour of interest-rate spreads after 1992, this information is not readily apparent from the raw data plots. There is thus a strong case for investigating the robustness of claims about the "information" of interest rate spreads for predicting inflation both on statistical grounds.

Table 1 reports the full-sample estimates of the slope coefficient β of the inflation-change equation (1). There is virtually no information in the spreads at the lower end of the "yield-curve". However, when considering spreads including 12-month rates, the explanatory power increases, and the coefficient of interest becomes significant. Furthermore, one cannot reject the hypothesis of real interest rates being constant. The results confirm those of Mishkin (1990b) and Siklos (2000) in the sense that only longer rates contain any information[2].

[Table 1 about here.]

The occurrence of numerous unexpected monetary shocks during the sample period and consequently the importance of investigating the implications of these shocks for the informational content of interest-rate spreads has already been emphasised. For this reason recursive estimation was employed to assess the stability of the information content of spreads across the monetary regimes. Figure 2 presents the forward- (upper panels) and backward-recursive estimates of the β coefficient with 24 starting observations and error bands given by the Newey-West corrected standard errors. A quick consistency check indicates the the end-points of the recursive estima-

[2] All the results were generated by programmes written in Ox (see Doornik, 1999). These are available on request.

tion conform to the results presented in Table 1 with significant coefficient estimates being recorded for these three spread variables.

[Figure 2 about here.]

The results presented in Figure 2 are particularly instructive. The forward recursions of β do not indicate any particular changes in the coefficient estimates after the estimation procedure has settled down. Although there is appears to be downward trend in the mid-1990s, the estimates are surprisingly stable, with precision increasing over the sample. There is therefore no convincing evidence of structural breaks in the information content so far. This does not really accord either with intuition or with the observed behavioural changes evident in Figure 1. This is, however, an example of the history of a previous regime dominating the coefficient estimates in the latter stages of the sample. When the recursion algorithm is run backwards, so that the sample starts with the most recent information, there are three noteworthy aspects to the estimated β's.

1. The coefficient estimates are unstable around the leaving of ERM, but they are also indicate that the β coefficient is zero after 1990, the date at which sterling entered the exchange rate mechanism of the European Community. The circumstancial evidence of Figure 2 is therefore that the information content of interest spreads is nonexistent under the monetary regime of inflation targeting.

2. The shorter maturities indicate a break in early 1985 corresponding to the sterling crisis which reached its peak in February of that year. Interestingly enough this break is not a feature of the 12,6 maturity pair indicating that the crisis affected only short-term LIBOR rates.

3. There is a change in the estimates in 1980 but as this effect is short-lived and does not appear to result in a sustained change in the value of the coefficient, this particular even is not investigated any further.

As a first step to isolating the correct breakpoints in the sample, the test for structural change with an unknown breakpoint of Andrews and Ploberger (1994) was employed. The results of this exercise confirmed that the major breaks in the relationship occurred in February 1985 (for the maturity pairs 12,1 and 12,3) and October 1990, thus confirming the results of the backward recursive estimation[3]. At this stage simple F-tests confirmed that significance of the breakpoints. To examine the behaviour of the informational content of interest-rate spreads over the sub-samples isolated by the recursive estimation, the inflation-change regression was estimated for the periods 1976:1 – 1985:1, 1985:3 – 1990:9 and 1990:11 – 1999:9. The results of this exercise are reported in Table 2.

[Table 2 about here.]

For the early sub-sample all the maturity pairs involving 12-month LIBOR are significant, indicating that the longer end of the interbank interest-rate spectrum contained important information on the future course of UK inflation. After the 1985 sterling crisis and the subsequent period of shadowing the deutschemark in the late 1980s, this relationship appears to break down for the maturity pairs involving the shorter-term rates (1- and 3-month LIBOR). Although the coefficient estimates recorded are higher than for the earlier the sub-sample, so to are the standard errors and the coefficients are

[3]The actual months recording the higherst values of the Andrews and Ploberger test varied slightly for different maturity pairs. As the maximum values of the tests were not uniformly located at a particular month we could see no reason not to use the months in which the economic event of interest occurred.

no longer significant at the 5% level. This deterioration in statistical significance is particularly noticeable for the spread involving 3-month LIBOR. Also evident is that fact that the estimate of β for the 12,6 month LIBOR pair is constant (and significant) for both these sub-samples, a result which confirms the intuition of Figure 2.

Perhaps the most significant result concerns the period after entry to the ERM in 1990. After this date there is no evidence of any relationship whatsoever between spreads of interest rates and changes in future inflation, a result consistent with a regime of credible inflation targeting. This is an interesting result given that a firm commitment to inflation targeting was only given on exit from the exchange rate mechanism in 1992. One possible interpretation is that entry to the ERM represented a fundamental commitment to fighting inflation which exit from the ERM did not significantly alter.

The story so far could therefore be described as an interesting one. There seems to be some information in the slope in the yield curve for future inflation, given the right maturities and the right monetary regime. Indeed Estrella and Mishkin (1997), on the basis of similar results for other European countries, advocate advocated the use of the term structure differential as a leading indicator for use in monetary policy decision making. The question of the reliability of the inference drawn from models of this kind remains, especially if they are to be used for in forecasting exercises as, for example, Jorion and Mishkin (1991). In particular, are the estimates and the corresponding standard errors of β reliable? Are the hypothesis tests reliable, when the models are so obviously misspecified? The next section tries to shed some light on these issues by generating the distributions of the β's by means of simulation.

4 A bootstrap approach

The usual argument made to support the validity of statistical inference in regressions of this kind relies on the consistency of OLS estimators on the basis of orthogonality conditions, valid under rational expectations, to validate hypothesis tests on the parameter of interest. It is true, however, that the residuals from the inflation-change equation are autocorrelated because of equation misspecification and overlapping observations. It is also likely that the residuals are heteroskedastic. Although the Newey-West standard errors correct for unknown forms of autocorrelation and heteroskedasticity, there is some evidence to suggest that the empirical size of the t-tests based on heteroskedasticity-corrected standard errors is too large (Horowitz, 1999). It may be, therefore, that a t- or normal distribution does not approximate the actual empirical distribution of β's particularly well. Consequently, Mishkin (1990b) derives the critical values for his hypothesis tests by Monte Carlo simulation. The approach taken here is to rely use the Newey-West correction for the basic results but to augment these by deriving confidence intervals for β by block-bootstrapping—a bootstrapping procedure which is able to take account of the temporal properties in the residuals.

The disadvantage of previous Monte Carlo approaches to this problem is that parametric models for the error process from the inflation-change equation need to be specified and estimated in order to implement the approach. This leaves scope for misspecifying the error process and introducing bias. Without a parametric model of error process, however, it is not possible to implement a residual-based bootstrap method. This conundrum may be overcome by implementing a bootstrap procedure which is capable of dealing with the correlation structure generally found in time-series data. The moving-block bootstrap with overlapping blocks (Künsch, 1989; Liu and

Singh, 1992) is used in this application

The basic idea of the moving-block bootstrap is easily outlined. Essentially blocks of data are resampled with replacement and added together to yield a time-series of approximately the same length as the original. This method preserves the structure in the data and should give accurate estimates of the standard error of the relevant parameter and also yield a distribution of estimates of this parameter so that a confidence interval may be established. The recently developed *sieve bootstrap* (Politis et al., 1997)

where each block of observations is itself regarded as a valid sub-sample for parameter estimation was considered as an alternative but it is not clear that this procedure offers any practical advantages over the moving-block bootstrap despite the fact it has asymptotic validity under weaker assumptions.

The optimal choice of block size in this bootstrap scheme has received some attention in the literature, e.g. Hall and Horowitz (1996). Rather than appeal to any asymptotically correct block size we follow Li and Maddala (1996) who infer that Künsch's 1989

suggestion to use "subjective judgement based on sample correlations" is an acceptable way to proceed. As the main autocorrelation problems in our data stem from overlapping observation problems and the maximum possible order of this overlapping interval is 12, we use two block sizes of 12 and 6.

[Figure 3 about here.]

[Table 3 about here.]

The distributions of the estimates of β recovered from the full-sample bootstrap are illustrated in Figure 3 and the corresponding full-sample results are presented in Table 3. In general, the bootstrapped distributions appear relatively well-behaved, although it is clear that blocks of size 12 provide a smoother distribution than blocks of size 6. The distributions based on the smaller block size also tend to increase the confidence interval slightly. We will therefore rely on the results using the larger blocks in the following. The summary of the results of the bootstrap are also very similar to the OLS results and the point estimates of the standard errors are similar to those obtained from the Newey-West correction. Perhaps the most striking feature of the results in Table 3 is the size of the confidence interval. Although the full sample results include a structural break and cannot therefore be regarded as a true reflection of the variation in the parameter, this feature of the results certainly requires careful examination in the context of the sub-sample results.

[Table 4 about here.]

The sub-sample results employing the moving-block bootstrap are reported in Table 4. For the earlier sub-sample the hypothesis of $\beta = 1$ cannot be rejected for the pairs of maturities involving 12-month LIBOR. From a statistical point of view, it is clear that the longer horizon LIBOR rates over this period did contain significant information about the future course of inflation. Despite this statistical significance, however, the size of the confidence interval remains large—even though the estimation takes place within one regime. This poses questions as to the practical use of the method in terms of forecasting and certainly urges caution in the use of these severely misspecified equations for purposes other than drawing very broad conclusions.

A second noteworthy feature of these results concerns the statistical significance of β during the period 1985-1990. The OLS results for the sub-samples reported in Table 2 are largely inconclusive for the maturity pairs involving 12-month LIBOR. Although the 12,6 maturity combination appears significant, the decision for the 12,3 and 12,1 maturity pairs is less clear-cut with p-values of 0.078 and 0.052 respectively for the hypothesis $\beta = 0$. The bootstrap results on the other hand are quite decisive; the hypothesis of $\beta = 0$ cannot be rejected for all of the maturity pairs.

A possible explanation for this turn of events is in the role taken by the exchange rate in policymaking. When Nigel Lawson became Chancellor of the Exchequer after the election of 1983, the stage was set for a long conflict within the government over the EMS and this itself increased the prominence of the exchange rate in policymaking. Over the period the exchange rate became accepted as at least one of the guides for policy. Indeed market perceptions of the prospect of ERM membership grew as monetary policy was allowed to be more and more guided by the deutschemark exchange rate. This led to the formal policy of "shadowing" the deutschemark between 1987 and March 1988. It appears from the results reported here that this change in monetary policy had the effect of decoupling the link between interest rates and future inflation even before ERM entry.

Given these observations, it is not surprising that the earlier results for the period of inflation targeting are confirmed and the hypothesis of $\beta = 0$ cannot be rejected. This may be taken as *prima facie* evidence that the inflation target is a credible one in the context of the UK. This result is consistent with the conclusions of Siklos (2000), who found that in New Zealand, which has a firm inflation target, the term structure contains little information about future inflation as opposed to Australia which targets inflation but with a relatively weaker commitment. What this suggests is that the nominal UK term structure is now a yardstick for future real interest rates as opposed to inflation.

5 Conclusion

This paper has examined the impact of changes in monetary policy on the informational content of the term structure with respect to future inflation in the UK. The results of combining of both forward and backward recursive estimation identified significant break points in the relationship between interest and inflation differentials, in correspondence with changes in monetary policy . The use of the moving-block bootstrap, to guard against incorrect inference in an equation where the residuals are both autocorrelated and heteroskedastic, reinforced the conclusion of the unreliability of the informational content.

Summing up, the results from this modelling exercise are fairly clear-cut. The interbank interest rates differentials involving the 12-month deposit rate contained information about future inflation prior to 1982. It also transpires, however, that the confidence interval for the crucial parameter governing this relationship is a broad one, a fact which would inhibit relying on this rather simple specification for rigorous forecasting purposes. The most interesting result to emerge is that there appears to be no information on future inflation in interest rate spreads after the sterling crisis of 1982 and the progressive importance of the exchange rate in policy formulation. There has been no change in this state of affairs after entry to the ERM in 1992. The interest rate spreads in the later period signal changes in real interest rates rather than changes in future inflation. This result is consistent with empirical work from other countries, which shows that the information content of interest spreads is higher in countries with less formal inflation targets (US and Australia) and non-existent in counties with a formal commitment to inflation targeting (New Zealand). Extracting information about the future path of inflation from the term structure is therefore to be considered a highly unreliable business.

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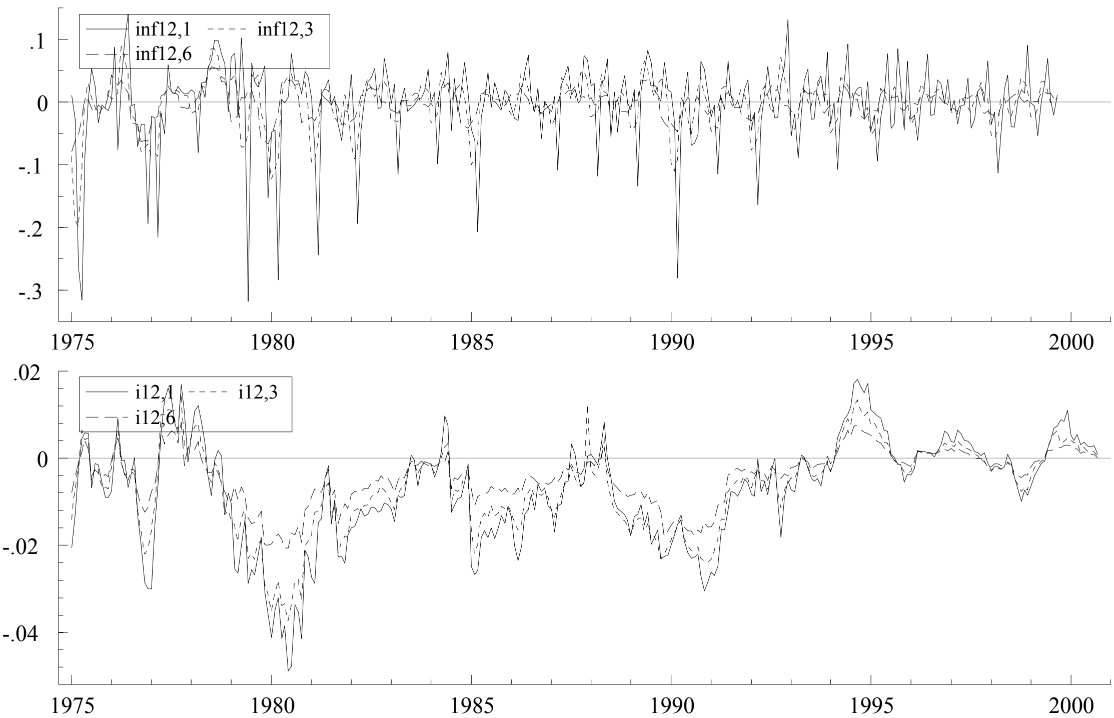

Figure 1: The spreads and changes in inflation.

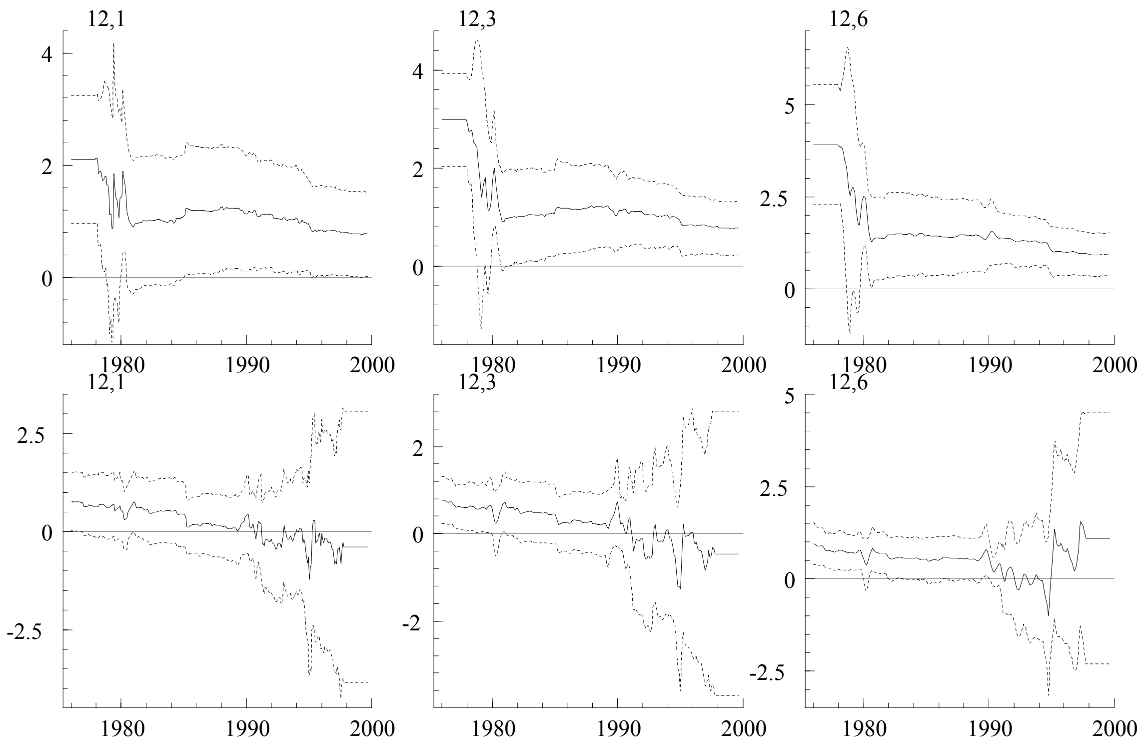

Figure 2: Recursive coefficients +/- 2 Newey-West standard errors .

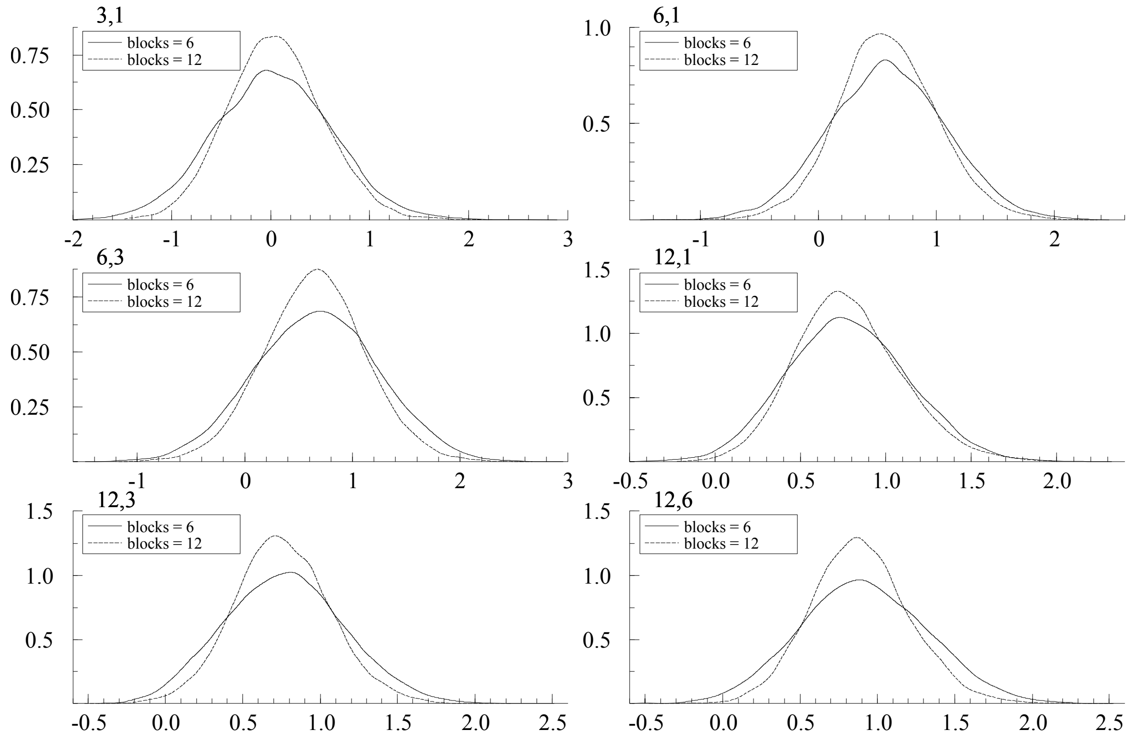

Figure 3: The bootstrap distributions of β based on the full sample.

Table 1: The effects of the slope of the 'yield-curve' on future inflation. Left hand column represents the pairs of LIBOR maturities for which the estimation was undertaken. Bootstrapped p-values lowermost.

OLS Estimates (1976:1 – 1999:9), Newey- West Standard Errors			
	β	p-value	$\bar{R}^2$
3,1	-0.030 (0.623)	0.962 0.948	-0.003
6,1	0.570 (0.551)	0.301 0.168	0.001
6,3	0.637 (0.443)	0.152 0.128	0.005
12,1	0.768 (0.376)	0.042 0.028	0.020
12,3	0.767 (0.271)	0.005 0.010	0.038
12,6	0.949 (0.283)	0.001 0.004	0.073

Table 2: The information content of interest rate spreads on future inflation for various sub-samples of interest. Left hand column represents pairs of LIBOR maturities for which the estimation is undertaken.

OLS Estimates with Newey-West Standard Errors for sub-samples									
	1976:1 – 1985:1			**1985:3 – 1990:9**			**1990:11 – 1999:9**		
	β	p-value	$\bar{R}^2$	β	p-value	$\bar{R}^2$	β	p-value	$\bar{R}^2$
3,1	-0.161 (0.864)	0.852	-0.009	1.120 (1.665)	0.504	-0.010	-0.441 (1.526)	0.773	-0.009
6,1	0.766 (0.689)	0.269	0.000	1.741 (1.796)	0.336	0.004	-0.458 (0.984)	0.642	-0.007
6,3	0.867 (0.650)	0.185	0.007	0.904 (1.012)	0.375	-0.004	-0.369 (1.044)	0.725	-0.008
12,1	1.088 (0.549)	0.050	0.036	1.766 (0.893)	0.052	0.034	-0.017 (0.507)	0.973	-0.009
12,3	1.114 (0.451)	0.015	0.070	1.219 (0.682)	0.078	0.038	0.231 (0.460)	0.616	-0.006
12,6	1.436 (0.545)	0.009	0.107	1.413 (0.547)	0.012	0.147	0.402 (0.306)	0.192	0.008

Table 3: Full sample bootstrap estimates of β with standard errors and confidence intervals. Left hand column represents pairs of LIBOR maturities for which the estimation is undertaken.

Moving-block Bootstrap Estimates (1976:1 – 1999:9)				
	Block Size = 6		**Block Size =12**	
	β	95% *C.I.*	β	95% *C.I.*
3,1	0.028 (0.595)	$-1.132 \longleftrightarrow 1.204$	0.051 (0.482)	$-0.857 \longleftrightarrow 1.024$
6,1	0.587 (0.498)	$-0.384 \longleftrightarrow 1.567$	0.587 (0.419)	$-0.227 \longleftrightarrow 1.437$
6,3	0.670 (0.571)	$-0.448 \longleftrightarrow 1.786$	0.651 (0.472)	$-0.279 \longleftrightarrow 1.584$
12,1	0.779 (0.354)	$0.115 \longleftrightarrow 1.488$	0.782 (0.313)	$0.215 \longleftrightarrow 1.450$
12,3	0.770 (0.384)	$0.049 \longleftrightarrow 1.546$	0.756 (0.310)	$0.155 \longleftrightarrow 1.394$
12,6	0.913 (0.411)	$0.129 \longleftrightarrow 1.739$	0.891 (0.321)	$0.292 \longleftrightarrow 1.542$

Table 4: Sub-sample bootstrap estimates, standard errors, and confidence intervals, using blocks of 12 observations. Left hand column represents the pairs of LIBOR maturities for which the estimation is undertaken.

Moving-block Bootstrap Estimates for sub-samples, block size =12						
	1976:1–1985:1		**1985:3–1990:9**		**1990:11–1999:9**	
	β	95% *C.I.*	β	95% *C.I.*	β	95% *C.I.*
3,1	0.010 (0.687)	$-1.143 \leftrightarrow 1.503$	0.039 (1.162)	$-2.668 \leftrightarrow 1.851$	-0.100 (1.282)	$-2.804 \leftrightarrow 2.284$
6,1	0.697 (0.498)	$-0.180 \leftrightarrow 1.795$	0.614 (1.144)	$-1.998 \leftrightarrow 2.571$	-0.522 (0.902)	$-2.622 \leftrightarrow 0.953$
6,3	0.648 (0.594)	$-0.556 \leftrightarrow 1.810$	0.779 (0.1.016)	$-1.777 \leftrightarrow 2.352$	-0.911 (1.480)	$-4.382 \leftrightarrow 1.387$
12,1	1.094 (0.514)	$0.224 \leftrightarrow 2.203$	0.715 (0.687)	$-0.798 \leftrightarrow 2.046$	-0.127 (0.487)	$-1.241 \leftrightarrow 0.679$
12,3	1.087 (0.544)	$0.124 \leftrightarrow 2.281$	0.652 (0.599)	$-0.731 \leftrightarrow 1.745$	-0.086 (0.669)	$-1.547 \leftrightarrow 1.063$
12,6	1.452 (0.651)	$0.360 \leftrightarrow 2.967$	1.365 (0.940)	$-0.559 \leftrightarrow 2.865$	0.172 (0.537)	$-0.829 \leftrightarrow 1.371$