An assessment of the relative importance of real interest rates, inflation and term premia in determining the prices of real and nominal UK bonds.

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Abstract

We use a vector autoregression to decompose the causes of unanticipated movements in bond prices into news about fundamentals (expected future real interest and inflation rates) and expected future risk premia. This decomposition is applied to UK short- and long-maturity nominal bonds, and to UK index-linked (i.e. approximately real) bonds. We also examine the causes of changes in *relative* conventional and real bond prices. The results suggest that for *both* bond types, real-rate news plays an insignificant role, and that even for 'real' bonds it is dominated by news about inflation. Both bonds are strongly influenced by news about future risk premia, but these appear to be a common factor which has little influence on *relative* prices. It seems that news about inflation dominates relative price movements, and that such movements provide a reliable source of information about inflation expectations.

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1 Introduction

Theoretical models typically identify 'news' as the proximate cause of previously unexpected returns. This, however, merely raises a question as to which news is relevent, and why. In this paper, we investigate the relative importance of news about 'fundamentals' (real interest rates and inflation) and term, or risk, premia, in generating unexpected returns on UK real and nominal bonds.

The results also provide informal tests of models of the yield curve that are based on expected future short-term interest rates [see Campbell and Shiller (1993)] and on expected future inflation (the Fisher hypothesis). In the latter case, we consider whether investors *expect* future real rates to be constant despite the observation that *ex post* real rates are variable. Recent work on the strength of the Fisher effect in the United States [Mishkin (1992)], suggests that while the assumption of a constant expected real rate over short horizons is too extreme, it may be quite reasonable over the long term.

The separate identification of significant and insignificant types of news allows us to assess the extent to which information about fundamentals can be 'backed out' of asset prices. In particular, we investigate the reliability of using real and nominal-bond yields to make inferences about inflation expectations that might subsequently be used to condition monetary policy. Our results also offer some informal evidence about the credibility of the tight monetary policy introduced in the late 1970s, and suggest that the change of regime went largely unnoticed at the time. Finally, we offer some quantification of the extent to which uncertainty in financial markets could be reduced by a credible anti-inflation policy.

The analysis in this paper uses the discounted present value model of Campbell and Ammer (1993) to provide a rigid structure within which news about fundamentals competes with news about risk and other premia to explain unexpected asset price movements.

The discounting of expected future income streams is central to most of asset pricing theory. It follows that much of the observed volatility of asset prices reflects either the volatility of these expected income streams, or some form of volatility in the discounting process.

That income streams are not sufficiently variable to explain observed price volatility has been demonstrated in numerous papers, such as Shiller (1982), LeRoy and Porter (1981) and Grossman and Shiller(1981). This has provided the impetus for a range of papers aimed at generating and testing models with variable discount factors. Many of these have investigated theoretical models linking asset prices to consumption growth, following the consumption CAPM of Lucas (1978). A complementary approach, initiated by Roll (1988) looks instead at the relative importance of different forms of contemporaneous news in influencing prices. While this approach is appealing, the choice of news events is not constrained by an optimising model of asset price formation and, therefore, leaves open the question of why a particular news event should be influential. For example, to be consistent with the present-value model, contemporaneous news must contain information about current or future income streams or discount rates if it is to cause prices to change. Since many types of news may work through either, or both channels, this approach is not informative about the underlying economic mechanisms through which the news variables influence the asset prices.

An alternative to the contemporaneous-news approach was developed in Campbell and Shiller (1988) and Campbell (1991). In principle, the relatively unstructured range of news of Roll (1988) is replaced by two news items derived from what Campbell and Ammer (1993) call a 'dynamic accounting identity', namely future income and discount rates. 'News' here takes the form of revisions to explicit numerical values for forecasts of these variables over the entire life of the asset. The empirical analysis reported in this paper is based on Campbell and Ammer's methodology [4] which uses a vector auto regression to construct these forecasts.

Previous research in this area has concentrated exclusively on the excess returns on equities and nominal bonds. This paper offers two principal innovations. First, we report the results of including real bonds in both the theoretical structure and in the empirical work. Second, the structure *per se*,which provides a simple log-linear framework for the analysis of real bonds, may be of use in a range of further research.

The empirical results have relevance for two quite separate areas of economics. First, they develop the results of papers such as Campbell and Ammer (1993) by taking advantage of the additional information contained in the prices of real bonds to explain the behaviour of nominal bonds, while offering the obvious advantage of a similar analysis of real bonds *per se*.

Second, they offer some insight into the extent to which bonds of both varieties can be used to extract information about 'market' expectations of future real interest rates and inflation. The latter are of particular importance to monetary authorities who regard 'policy credibility' as an essential factor in their fight against inflation, and who might be tempted to use the information contained in asset prices in the policy-making process. An obvious example of such an authority is the Bank of England which, alone among central banks, publishes a quarterly term structure of inflation expectations derived from the relative yields on UK nominal and real gilts. The extent to which this term structure reflects news about expected inflation depends crucially on the relative importance of fundamentals and risk premia in determining its movements. Similarly, in the United States there is a continuing debate on the question of whether real bonds should be issued; the results reported here should offer some encouragement to those who base their support for such issuance on the ability of real bonds to convey important information about developments in economic

[4]We should like to thank John Campbell and John Ammer for making their programs available to us; these provided the basis for those used in our own work.

fundamentals.

2 An introduction to real bonds

Real, or index-linked, bonds were introduced in the United Kingdom in 1981 in an attempt to reduce the cost of financing the fiscal deficit. The high and variable rates of inflation throughout the 1970s had, it was argued, added an inflation risk premium to the required returns on conventional bonds; a premium that real bonds would not have to bear. Further encouragement for real-bond issuance came from the subsequent development of rational expectations models which emphasised the advantages of real bond's in enhancing the credibility of monetary policy; unlike conventional bonds, real bonds offer no incentive for the authorities to 'cheat' by using previously unexpected inflation to reduce the real value of the public sector's bond liabilities [5].

Despite the obvious attractions of index-linked assets the demand for them grew very slowly. The bonds tended to be purchased by long-term investors, and the residual float available to the market was insufficient to generate the liquidity necessary to attract the more active investors. A potential consequence of a lack of liquidity is a high degree of price volatility generated by factors specific to the asset, the market or, in extreme cases, particular investors. This raises obvious questions about the extent to which it is possible to distinguish the 'signal' that real bond prices contain about economic fundamentals from the noise generated by asset-specific premia.

In principle, the yield on a real bond should be its real rate of return. The near-zero credit risk of government bonds suggests that, subject to market-risk and liquidity premia, this yield should offer the clearest available indication of investors' marginal rates of time preference. Similarly, in principle, the real-bond yield can be subtracted from the appropriate nominal-bond yield to generate a measure of investors' expectations of inflation. In practice, however, UK index linked bonds are not precisely 'real' and the calculation of real yields and inflation expectations is considerably more complicated than suggested above.

A perfectly indexed bond would pay a nominal coupon equal to the coupon rate announced at the time of issue multiplied by the increase in the consumer/retail price index between the issue date and the time of payment. In practice, the period over which the indexation occurs lags both of these dates by eight months (this permits traders to calculate the amount of accrued interest to be exchanged in trades that take place between coupon payments). Consequently any inflation that occurs in the eight months prior to payment reduces the real value of the coupon. Thus UK index-linked bonds are actually a combination of real and nominal bonds.

This feature of index-linked bonds creates a number of technical difficulties. For example, it is not possible to calculate their yield to maturity in the usual

[5] See Lawson (1992) for an entertaining account of the introduction of index-linked bonds.

way since neither the real nor the nominal value of their coupons can be known in advance. Similarly, and more relevant for this paper, changes in the price of an index-linked bond may reflect changes in inflation expectations, albeit with a sensitivity well below that of a purely nominal bond. It is this joint dependence of nominal and 'real' UK bond prices on inflation that renders the simple calculation of inflation expectations suggested above inappropriate.

3 Using index-linked bonds to infer inflation expectations

The yield to maturity on an index-linked bond can be calculated conditionally on an assumed profile of inflation throughout its remaining life. Quoted index-linked yields typically assume a constant 5% inflation rate, and are usually presented as a 'real' rate. This creates a temptation to subtract this real rate from the nominal yield on a nominal, or conventional, bond of equivalent maturity (or duration) in order to generate a figure for average expected inflation over the remaining life of the bonds. The potential inconsistency between the derived rate and that assumed at the outset is obvious. This conflict can, however, be resolved by an iteritive process whereby the generated expected inflation is used to recompute the real yield on the index-linked bonds, from which a new figure for inflation can be obtained, and so on. This approach, which generates so-called 'break-even inflation rates' suffers from two problems. First, it does not generate a term structure of inflation since it can be applied only to those maturities where there are equivalent pairs of real and nominal bonds. Second, it takes no account of the premia that were discussed above i.e. the inferred rates of inflation will be contaminated by factors such as risk or liquidity premia.

The first of these problems has been overcome by the Bank of England by applying an iteritive process to complete yield curves rather than to the yields on specific bonds. First, a real yield curve is calculated from real-bond prices on the basis of an initial constant 5% profile for inflation. This is then subtracted from the familiar nominal curve. The derived profile (which will almost certainly not be flat) is then used to recalculate the real curve, and so on. The second problem remains, however. There is no guarantee that these iteritive processes will produce inflation term structures that are close to the actual expectations of market participants. Under reasonable conditions, the inferred expectations will tend to exceed agents' actual expectations.

One of the objectives of this paper is to measure the extent to which variations in these structures are due to changing expectations of inflation, as opposed to changes in risk or other premia. Clearly, in the absence of a model of the premia, a dominant role for inflation expectations in driving a wedge between real and nominal bonds is crucial to their use as a source of fundamental economic information. It is not obvious, *a priori*, that this condition will be satisfied.

4 Asset price movements and 'news'

4.1 An analytical framework

The focus of the empirical analysis will be excess one-month returns on nominal and real bonds. The starting-point in deriving a suitable analytical framework, however, is an equation for asset prices. The log real price p_t^r of any asset can be approximated by the following log linear version of the present discounted value equation [see Campbell and Shiller (1988) and Campbell (1991) for details of the log linearisation procedure];

$$p_t^r = k + (1-\rho)\sum_{i=0}^{m-1}\rho^i E_t d_{t+1+i}^r - \sum_{i=0}^{m-1}\rho^i E_t h_{t+1+i}^r + E_t \rho^m p_{t+m}^r \tag{1}$$

where, for each period (t+1+i), the asset's log real dividends d_{t+1+i}^r are discounted (arithmetically) by the log of the real 'required holding period return' (which may include risk premia) from (t+i) to (t+i+1), denoted h_{t+1+i}^r. The geometric discounting factor ρ derives from the approximation procedure, as does the constant k.

For the empirical analysis in this paper it will be convenient to work in terms of *nominal* rather than *real* asset prices. The appropriate conversion of equation (1) is accomplished by replacing the general terms in real dividends d_{t+1+i}^r, and the redemption value p_{t+m}^r, by their asset-specific values (ie their known nominal dividends deflated by the general price index) and rearranging. Having achieved this, for the assets under consideration here, the *unexpected* nominal return will then be simply the unexpected movement in the asset's nominal price.

4.2 Nominal bonds

In the case of nominal, or 'conventional', bonds the real dividend and redemption payments are simply the declared nominal payments c deflated by the general price index z, ie [6]

$$d_{c,t+1+i}^r = c_{c,t+1+i}^n - z_{t+1+i} \tag{2}$$

and

$$p_{c,t+m}^r = -z_{t+m} \tag{3}$$

[6] Superscript n and r are used to denote nominal and real variables respectively. Subscript c and g denote conventional (nominal) and index-linked (real) bonds.

where all nominal payments are scaled so that the log of the nominal redemption value is zero. The coupons are assumed to be paid in equal amounts each period.

Substituting these definitions into equation (1) generates the following expression for the log nominal price of an m-period conventional bond,

$$p^n_{c,m,t} = k'_c - E_t \sum_{i=0}^{m-1} \rho^i_c h^n_{c,m-i,t+1+i} \tag{4}$$

where $k'_c = k_c + c^n_c(1 - \rho^m_c)$.

The *unexpected* nominal one-month holding period return $\tilde{h}^n_{c,m,t+1}$ on an m-period conventional bond is simply the *unexpected* movement in the bond's nominal price (since there can be no surprises with regard to the nominal coupons). Hence,

$$\tilde{h}^n_{c,m,t+1} = -(E_{t+1} - E_t) \sum_{i=0}^{m-1} \rho^i_c h^n_{c,m-i,t+1+i} \tag{5}$$

Almost all asset pricing models express the risk premium as the excess of the total return over the 'safe' rate. Since our interest is in the role of revisions to expected risk premia in determining current asset prices and returns we recast the above framework in terms of excess returns $x_{c,m,t+1}$ i.e.[7]

$$x_{c,m,t+1} = h^n_{c,m,t+1} - r^n_t \tag{6}$$

If we define the real interest rate r^r as

$$r^r_{t+1} = r^n_t - \pi_{t+1} \tag{7}$$

where π is inflation in the period t to t+1, then

$$x_{c,m,t+1} = h^n_{c,m,t+1} - r^r_{t+1} - \pi_{t+1} \tag{8}$$

The determinants of the *unexpected* excess return $\tilde{x}_{c,m,t+1}$ are obtained by substituting for $h^n_{c,m}$ throughout the right-hand side of equation (5). For the left-hand side, since the nominal rate r^n_t is known at time t, $\tilde{x}_{c,m,t+1}$ is equal to $\tilde{h}_{c,m,t+1}$. Hence,

$$\tilde{x}_{c,m,t+1} = -(E_{t+1}-E_t)\{\sum_{i=1}^{m-1} \rho^i_c r^r_{t+1+i} + \sum_{i=1}^{m-1} \rho^i_c \pi_{t+1+i} + \sum_{i=1}^{m-1} \rho^i_c x_{c,m-i,t+1+i}\} \tag{9}$$

which we can rewrite as

$$\tilde{x}_{c,m,t+1} = -\tilde{x}_{c,r,t+1} - \tilde{x}_{c,\pi,t+1} - \tilde{x}_{c,x,t+1} \tag{10}$$

[7]Since we are concerned only with nominal excess returns, no superscript is attached to x.

where $\tilde{x}_{c,r,t+1}$ represents news about real rates, and so on. Note that each of these terms is defined for conventional bonds only, since they are functions of the conventionals' linearisation constant ρ_c.

Equation (10) expresses an *ex-post* unobservable in terms of three *ex-ante* unobservables. This may appear to be an unpromising focus for an empirical investigation. However, data for each of the *ex-ante* variables can be generated iteritively using an estimated vector auto regression. While the details of this will be explained below, it is worth noting at this point that the *ex-post* unexpected excess return $\tilde{x}_{c,m,t+1}$ can be expressed in terms of the unexpected yield change in the yield to maturity of the bond i.e.

$$\tilde{x}_{c,m,t+1} = -\phi_{m-1}(y_{c,m-1,t+1} - E_t y_{c,m-1,t+1}) \tag{11}$$

where $\phi_{m-1} = (1 - \rho_c^{m-1})/(1 - \rho_c)$ is the duration of the bond, and measures the sensitivity of its price to movements in its yield. The VAR that will be used to make long-horizon forecasts for the right-hand side of (10) will also be used to generate one-period-ahead forecasts of yields. These will then be compared with the yield outturns to obtain the unexpected *ex-post* return.

4.3 Short-term interest rates and long-short yield spreads

It is common practice among portfolio managers to decompose movements in the yield curve into two components; a movement parallel to the initial position of the curve, and a change in its slope. Similarly, spread variables, measuring the slope of the curve, play a significant role in the empirical analysis of asset prices in many papers [see, for example, Campbell and Shiller (1991)]. These approaches can be combined, as in Campbell and Ammer (1993), to permit a decomposition of the two movements of the curve in the same way as that for the return on a long-bond.

We first have to construct two 'assets', each of whose returns depend on only one of the two indiviual types of yield curve movement. To do this we construct two notional portfolios. First, a portfolio of two-month bonds captures the 'parallel' movements in the curve. Second, a 'duration-weighted' spread-portfolio, which is 'short' of short bonds and 'long' of long bonds, captures the change in the slope of the curve. The duration-weighting, which requires the ratio of short-bonds to long-bonds in the spread-portfolio to equal minus ϕ_{m-1}, ensures that parallel shifts in the curve have no impact on the portfolio's returns.

4.3.1 Unexpected returns on short-term bonds

The unexpected one-month return on a two-month bond, purchased at time t, will depend only on the unexpected movement in the nominal one-month rate from t to t+1. Thus,

$$\tilde{x}_{c,2,t+1} = -(E_{t+1} - E_t)(r^r_{t+2} + \pi_{t+2}) \tag{12}$$

which we rewrite as

$$\tilde{x}_{c,2,t+1} = -\tilde{x}_{c,1,r,t+1} - \tilde{x}_{c,1,\pi,t+1} \tag{13}$$

where the additional subscript indicates that the revisions are to expectations for one future period only.

As with long-term bonds, the *ex-post* unexpected excess return is calculated in terms of a forecast error i.e.

$$\tilde{x}_{c,2,t+1} = -(y_{1,t+1} - E_t y_{1,t+1}) \tag{14}$$

4.3.2 Unexpected returns on a spread portfolio

The unexpected returns on the spread portfolio are,

$$\tilde{x}_{s,m,t+1} = -\tilde{x}_{s,r,t+1} - \tilde{x}_{s,\pi,t+1} - \tilde{x}_{s,x,t+1} \tag{15}$$

where the subscript s indicates that the variable is relevent to the spread portfolio only. The 'explanatory' right-hand side variables here are

$$\tilde{x}_{s,r,t+1} = (E_{t+1} - E_t) \sum_{i=1}^{m-1} \rho_c^i \phi_{m-1-i} \triangle r_{t+2+i} \quad (16)$$

$$\tilde{x}_{s,\pi,t+1} = (E_{t+1} - E_t) \sum_{i=1}^{m-1} \rho_c^i \phi_{m-1-i} \triangle \pi_{t+2+i}$$

$$\tilde{x}_{s,x,t+1} = (E_{t+1} - E_t) \sum_{i=1}^{m-1} x_{s,m-i,t+1+i} \quad (17)$$

The *ex-post* excess return on the spread portfolio is,

$$\tilde{x}_{s,m,t+1} = -\phi(s_{m-1,t+1} - E_t s_{m-1,t+1}) \quad (18)$$

where $s_{m-1,t+1} = y_{m-1,t+1} - y_{1,t+1}$ is the long-short yield-to-maturity spread.

4.4 Index-linked bonds

The six-monthly coupon payments on index-linked bonds are indexed over a period that starts two months before the issue of the bond, and ends eight months before payment is made. Thus the declared 'real' coupon is adjusted by the difference between the price level eight months ahead of payment z_{t-8} and the appropriate reference level $\bar{z}$. Hence, the real value of the coupon depends on its declared value c_g, the efficiency of the indexation (which, for the algebraic derivation, we assume to have a general lag of l months) $(z_{t-l} - \bar{z})$ and the contemporaneous price level z_{t+1+i} i.e.

$$d^r_{g,t+1+i} = c + (z_{t-l} - \bar{z}) - z_{t+1+i} \quad (19)$$

similarly, for the redemption value of the bond,

$$p^r_{g,t+m} = (z_{t+m-l} - \bar{z}) - z_{t+m} \quad (20)$$

In either case, the effects of inflation on the real value are quite transparent. For example, in the case of the coupon, equation (19) can be rewritten as,

$$d^r_{g,t+1+i} = (c - \bar{z}) - \triangle_l z_{t+1+i} \quad (21)$$

ie the real value of the coupon, given the known values of c and $\bar{z}$ depends on inflation during the period of the indexation lag. The price of the bond will of course reflect the expected level of this inflation.

The log nominal price of the index-linked bond can be rewritten as[8],

$$p^n_{g,t} = k'_{g,t} - E_t \sum_{i=0}^{m-1} \rho_g^i h^n_{g,m-i,t+1+i} + E_t \sum_{i=l}^{m-1} \rho_g^i \pi_{g,t+1+i-l} \qquad (22)$$

The *unexpected* nominal one month holding period return on an m-period index-linked bond is, as in the case of a conventional bond, simply the *unexpected* movement in the bond's nominal price. Hence[9],

$$\tilde{h}^n_{g,m,t+1} = -(E_{t+1} - E_t) \sum_{i=1}^{m-1} \rho_g^i h^n_{g,m-i,t+1+i} - (E_{t+1} - E_t) \sum_{i=l}^{m-1} \rho_g^i \pi_{g,t+1+i-l} \qquad (23)$$

Substituting for $\tilde{h}$ and h in terms of r^r, π and x, as in the case of conventional bonds, gives

$$\begin{aligned} \tilde{x}_{g,m,t+1} &= -(E_{t+1} - E_t)\{(\sum_{i=1}^{m-1} \rho_g^i r^r_{t+1+i} + (\sum_{i=1}^{m-1} \rho_g^i \pi_{t+1+i} - \sum_{i=l}^{m-1} \rho_g^i \pi_{t+1+i-l}) \\ &\quad + \sum_{i=1}^{m-1} \rho_g^i x_{g,m-i,t+1+i})\} \end{aligned}$$

or

$$\tilde{x}_{g,m,t+1} = -\tilde{x}_{g,r,t+1} - \tilde{x}_{g,\pi,t+1} - \tilde{x}_{g,x,t+1} \qquad (24)$$

The excess return on index-linked bonds will be one of the variables in the VAR, so the *ex-post* unexpected excess return will be calculated as the forecast error in this variable directly, rather than in terms of errors in forecasts of yields.

4.5 Relative returns on real and nominal bonds.

Relative movements in the prices of real and nominal bonds are likely to contain valuable information about market expectations of inflation. However, as discussed earlier, this information can be very difficult to extract, largely because the prices may depend on factors other than expected inflation. The analytical structure outlined above can be applied to relative returns by subtracting the excess returns on index-linked bonds from those on conventionals.

$$\tilde{x}_{rel,m,t+1} = -(\tilde{x}_{c,r,t+1} - \tilde{x}_{g,r,t+1}) - (\tilde{x}_{c,\pi,t+1} - \tilde{x}_{g,\pi,t+1}) - (\tilde{x}_{c,x,t+1} - \tilde{x}_{g,x,t+1}) \qquad (25)$$

[8]The constant k' is not importatnt for our purposes. It is, however, discussed in some detail in an Appendix.

[9]The appendix demonstrates that $(E_{t+1} - E_t)k'_{g,t+1} = 0$.

As is clear from the earlier discussion, news about inflation will affect the unexpected returns of both assets. The returns will, however, be affected in different ways; nominal bonds being the more sensitive to changes in inflation expectations. In contrast, news about future excess returns may be unique to each asset, although it is likely that some correlation between these news items will be present since, apart from their relation to inflation, the assets are virtually identical. News about real rates, however, is common to both assets, and appears in the two returns equations in almost exactly the same way. The only minor difference arises from the difference between the linearisation coefficients ρ. Since these are both very close to unity the net effect of real rate news on the relative returns should be very small.

5 Empirical analysis

5.1 An overview

The analytical structure set out above identifies candidate causes of unexpected excess returns. We now set out the empirical methodology that enables us to estimate the relative importance of these causes for UK conventional and index-linked gilts markets during the period 1983-93.

The first step is to estimate a model capable of generating two types of forecasts. First, we require one-period-ahead forecasts of yields (or excess returns, in the case of index-linked bonds). Second, we need forecasts of real rates, inflation, and excess returns over longer horizons. We use a single vector-autoregression to meet both of these requirements. In the case of the long-term excess return the subsequent analysis is, broadly, as follows. At any sample period t we calculate both the one-period-ahead forecast of the long-term yield, and the three sets of long-horizon forecasts. The latter are then aggregated over the entire horizon using the geometric discount factor ρ. This generates four numbers dated t. In the next period, we can find the error in the previous one-period- ahead forecast from the raw data dated $t + 1$; this gives us the left-hand side of equation (10) for the period $t + 1$. We also calculate new long-horizon forecasts to correspond to those made in the previous period, and aggregate over the horizon as before. The differences between the three long-horizon forecasts made at t and $t + 1$ are the forecast revisions that appear on the right-hand side of (10). Thus for period $t + 1$ we have each of the required four numbers. This procedure is then repeated for every period until we have four time-series corresponding to the four elements of (10). It is then a straightforward exercise to calculate the variances and covariances of these series.

The final step is to decompose the variance of the *ex-post* unexpected excess return into its components in the form of variances and covariances. Each component is then measured as a share of the total variance. For the long-term bond we have, from (10),

$$
\begin{aligned}
var(\tilde{x}) &= var(\tilde{x}_r) + var(\tilde{x}_\pi) + var(\tilde{x}_x) \\
&\quad +2cov(\tilde{x}_r, \tilde{x}_\pi) + 2cov(\tilde{x}_r, \tilde{x}_x) + 2cov(\tilde{x}_\pi, \tilde{x}_x)
\end{aligned}
$$

Similar expressions can be derived for each of the other assets introduced above.

In practice, since we are concerned only with errors in, or revisions to, forecasts, we work with the residuals from our model, rather than the forecasts *per se.* The practical aspects of the empirical analysis are discussed below.

5.2 Empirical proxies for news

It is assumed that each of the expectations terms can be proxied accurately by forecasts based on a VAR that incorporates a range of financial variables. The results presented subsequently are all conditional on the accuracy of this VAR as a proxy for agents' actual expectations generating mechanism. Underlying the choice of variables is the familiar proposition that all of the information used by market participants in forming their expectations is reflected in the prices and yields of the assets in which they trade. Thus the VAR consists, in the main, of market interest rates.

The VAR is used to provide proxies for revisions to expectations of *future* real rates and inflation, and for the unexpected *contemporaneous* excess return. In practice, revisions to expectations of *future* excess returns, the fourth element of equation (10), are not calculated directly but are found as the equation residual after substitution of the other three VAR-generated series. Any deficiencies in the VAR as a proxy for the true expectations generating process will be incorporated in this residual and may lead to its importance being overestimated in the empirical results.

The VAR state-vector w consists of six variables, some of which mirror those used in Campbell and Ammer (1993). The real rate, r^r, the change in the nominal one-month rate, $\triangle y_1$, and the ten-year-one-month spread, s_{120}, are all needed in order to construct the unexpected returns and the 'news' variables. The index-linked excess return, $igxs$, enters directly. Additional variables that aid the forecasting performance of the VAR are the three-month-one-month nominal yield spread, s_3, and an approximation to the slope of the ten-year real yield curve, $igslope$. The latter variable is equal to the difference between the approximate real yield to maturity of a ten-year index-linked bond and the contemporaneous *ex-post* one-month real rate.

For the first-order VAR,

$$
w_{t+1} = Aw_t + \epsilon_{t+1} \tag{26}
$$

where A is the coefficient matrix of the VAR, and ϵ is the error vector. The state vector is,

$$w_t = (s_{3,t}, r_t^T, \Delta y_{1,t}, s_{120,t}, igxs_t, igslope_t)' \tag{27}$$

The unexpected excess returns and revisions to expectations for all five decompositions can be obtained from the VAR error vector. If we define vectors e_i, for i=1 to 6, as the columns of a 6×6 identity matrix, then the unexpected excess return on an index-linked bond is simply $e_5'\hat{\epsilon}_{t+1}$. Unexpected returns on long-term bonds are approximated by the unexpected change in their yield, multiplied by the appropriate duration ϕ. Hence this return can be obtained from the VAR as $-\phi(e_3 + e_4)'\hat{\epsilon}_{t+1}$. The relative return on these two bonds follows directly. Similarly, the unexpected return on the short-bond portfolio is $-e_3'\hat{\epsilon}_{t+1}$, while that on the spread-portfolio is $-\phi e_4'\hat{\epsilon}_{t+1}$. Similar manipulations can be applied to revisions of expected future values of w, where these are calculated as

$$(E_{t+1} - E_t)w_{t+1+i} = A^i\hat{\epsilon}_{t+1} \tag{28}$$

The method used for estimating the VAR is exactly that of Campbell and Ammer (1993). The coefficients of A, and the elements of the error covariance matrix are estimated by the Generalized Method of Moments [Hansen(1982)] to produce a heteroskedasticity-consistent covariance matrix for the complete set of VAR parameters and for the subsequent variance decomposition.

6 Data

Conventional bond yields y_{120} are taken from the Bank of England's par-yield curve. The construction of index-linked bond data was more difficult because the index-linked market in the United Kingdom has neither the depth nor the variety of bonds of a typical market for conventional bonds; to date the number of index-linked bonds has never exceeded 20. In the absence of a ten-year bond, and given the obvious difficulties of estimating a yield curve from such a small number of assets, a simple representative ten-year-bond holding-period return was calculated as a weighted average of the actual returns on the two bonds either side of the ten-year maturity, with the weights set according to each bond's distance from the ten-year point. The inevitable disadvantage of the index-linked data is that they will be a rather less accurate measure of ten-year returns than those for conventionals. This however derives from the nature of the market and, as such, reflects the conditions faced by investors.

The other nominal interest rates are a three-month interbank rate for y_3 and a one-month interbank rate for y_1. The retail price index (RPI) is used for z since this is the index used by the issuing authorities for the indexation of real bonds. The *ex-post* one-month real interest rate was calculated as the one-month interbank rate minus the change in the general price index over the same month.

End-month data were used throughout for the results reported in the text. However, due to the potentially high level of noise in such data, some of the variance decompositions were also performed using month-average data. The results of the latter, which turned out to be very similar to those obtained using end-month data, are included in an appendix.

6.1 General characteristics of the data.

6.1.1 Summary statistics: means, standard deviations and AR(1) coefficients

Summary statistics for interest rates, asset returns, inflation, and the variables in the VAR (all on an annualised basis), are shown in Table 6.1.1.

Table 6.1.1
Summary statistics: 1983.3 to 1993.3

	mean	s.d.	AR(1)		mean	s.d.	AR(1)
y_1	11.195	2.216	0.985	s_3	0.017	0.276	0.456
y_3	11.212	2.183	0.977	r^r	6.079	5.822	0.125
y_{120}	10.481	1.160	0.950	$\triangle y_1$	-0.046	0.713	0.108
$inflation$	5.137	5.162	0.208	s_{120}	-0.713	1.914	0.940
$convxs$	2.316	35.760	0.116	$igxs$	-3.126	23.944	0.111
igy_{120}	3.898	0.790	0.790	$igslope$	-2.181	5.781	0.117

A common criticism of index-linked bonds is that they have not offered sufficient returns to attract large numbers of investors. The figures above offer some support for that position since they show that real bonds produced an annual return some 3% below that of the one-month rate, and some 5.5% below that of the equivalent nominal bond *convxs*. While it is to be expected that the average return on nominal bonds should exceed that on real bonds, since the latter offer considerable insurance against inflation, the 'inflation risk premium' appears to be quite substantial. Similarly, since the effects of news about inflation should be greater for nominal than for real bonds it is not surprising that the standard deviation of nominal-bond returns is the greater by about 50%.

Both the nominal and 'real' yield curves sloped down on average during the sample period. A simple expectations-model-based interpretation of this is that, on average, both rates were expected to fall, but that real rates were expected to fall further than nominal rates. On average, real rates were expected to decline by about 2 per cent from the high one-month *ex-post* level of 6% to a more reasonable 4%. This impression, that real rates were typically expected to revert to a 'norm', will emerge strongly from the results of the variance decomposition. The variability of excess returns on real bonds is very high, which suggests that there is a lot for the variance decomposition to explain. Casual inspection suggests that the real rate, with a standard deviation of 5.8%, should account for about 30% more of this variability as inflation. This conclusion will not,however, be supported by the variance decomposition.

6.1.2 The predictability of excess returns

Excess returns are clearly an important component of the total returns on both types of asset. To the extent that these are required *ex ante* as compensation for risk, they should be predictable. It is also possible that some of the factors that generate such *ex ante* premia are common to both assets.

Table 6.1.2 shows regressions of the excess returns for real and nominal bonds on the (lagged) variables included in the VAR[10].

Table 6.1.2
Regression equations for excess returns

	convxs	*igxs*
$s_{3,t-1}$	-3.611	-2.183
	(-0.301)	(0.237)
r^r_{t-1}	28.495	23.271
	(3.316)	(3.383)
$\Delta y_{1,t-1}$	6.829	0.275
	(1.437)	(0.0848)
$s_{120,t-1}$	2.287	0.829
	(1.236)	(0.698)
$igxs_{t-1}$	0.129	0.187
	(0.874)	(1.809)
$igslope_{t-1}$	28.667	23.871
	(3.330)	(3.507)
R^2	0.103	0.159

The R^2s indicate that excess returns on real bonds are rather more predictable than those on nominal bonds. This probably reflects the fact that they are less exposed to revisions to expected inflation. In both cases the principal sources of useful information are the lagged real rate variables r^r and *igslope*. Since the latter includes the lagged real rate, the equations were also estimated with the index-linked yield and the real rate entering separately; the results revealed that the more useful component of the slope was the yield. Additional useful information may be present in the lagged change in one-month rates and the lagged yield curve slope.

The correlation coefficient of the predictable elements of the excess returns was 0.83, which suggests that much of the variation in excess returns can be

[10] White-adjusted t-statistics in parentheses. All rates of return were measured in terms of per cent per month and were included in the equations in mean-deviation form.

interpreted in terms of a single, cross-market, time-varying risk premium. Nevertheless, there appear to be some predictable elements that are specific to each of the two types of asset.

6.1.3 Time-series properties and stationarity

All of the variables in the VAR were stationary at the 5% level (on the basis of Dickey-Fuller and augmented Dickey-Fuller tests, with and without trends) with the possible exception of the ten-year spread, which was stationary at the 10% level, depending upon the particular form of the test used, and on the sample period. While it is not absolutely clear whether this spread is I(1) or I(0) the balance of the evidence favours the latter, and is supported by stronger implications of stationarity when a longer sample period is analysed (1978 to 1993). In particular, it seems that the spread follows a highly persistent, stable, autoregressive process (with an AR(1) coefficient of 0.94 in both the long sample and our estimation sample). In such circumstances tests for stationarity are known to have low power.

An indication of the persistence of shocks to each of the variables in the VAR can be obtained from the first-order autoregression coefficients reported in Table 6.1.1. Innovations in each of the variables decay very quickly, with the exceptions of those to the two nominal spreads s_3 and s_{120}. It appears that once the slope of the nominal yield curve has moved it is relatively slow to revert to its long-run level, and that, in the course of this reversion, the curve 'bends' i.e. it reverts more quickly at the short end than over its entire length.

7 Empirical results

7.1 The VAR

The coefficient matrix for the first-order VAR employed in the variance decomposition is reported in Table 7.1, with corrected t-statistics in parentheses

Table 7.1

VAR(1) coefficient estimates

	$s_{3,t}$	r^r_t	$\triangle y_{1,t}$	$s_{120,t}$	$igxs_t$	$igslope_t$
$s_{3,t+1}$	0.358	-0.148	0.022	0.135	-0.001	-0.142
	(5.231)	(2.677)	(0.891)	(1.092)	(1.352)	(2.564)
r^r_{t+1}	-1.749	2.651	0.185	-0.022	0.028	2.558
	(0.739)	(1.919)	(0.286)	(0.077)	(1.172)	(1.835)
$\triangle y_{1,t+1}$	1.140	-0.157	0.055	0.011	0.001	-0.170
	(3.268)	(1.084)	(0.389)	(0.266)	(0.495)	(1.187)
$s_{120,t+1}$	-1.093	-0.213	-0.144	0.973	-0.003	-0.202
	(3.945)	(1.671)	(1.516)	(26.272)	(1.293)	(1.581)
$igxs_{t+1}$	-2.183	23.271	0.275	0.829	0.187	23.871
	(0.244)	(3.485)	(0.087)	(0.719)	(1.865)	(3.614)
$igslope_{t+1}$	1.772	-1.858	-0.196	0.014	-0.030	-1.762
	(0.752)	(1.357)	(0.309)	(0.048)	(1.246)	(1.275)

The VAR matrix contains a number of statistically insignificant coefficients but produces quite reasonable forecasting power, due in part to the inclusion of the additional variables that are not used in the variance decomposition *per se*.Among these, s_3 contributes significantly to the equation for the change in the one-month rate y_1 and is largely responsible for an R^2 of 0.3. The three-month spread contributes significantly to the forecast of the ten-year spread via its predictive power for the change in short rates, as shown by the opposing signs of its coefficients in the two equations (for $\triangle y_{1,t+1}$ and $s_{120,t+1}$). The equation for the excess return on real bonds includes two coefficients that are distorted by the inclusion of the *ex-post* real rate in *igslope* and as a separate variable. The strong collinearity between these variables is reflected in all of the equations. In the case of the slope variable the large positive coefficient simply reflects a duration effect. A positive slope at time t leads to a negative slope at t+1, as shown by the *igslope* equation itself. This change, which will probably incorporate a decline in the yield on real bonds, generates a positive holding period return on these bonds. The largest eigenvalue of matrix A is 0.93, indicating that the VAR is dynamically stable.

A second-order VAR was also estimated, with similar dynamic characteristics but at a significant cost in terms of degrees of freedom. Variance decomposition results are presented below for both systems, and show that parsimony in the VAR order does not significantly alter our conclusions.

7.2 Variance decompositions for conventional bonds

7.2.1 Innovations in long-term-bond returns

Table 7.2.1[11] presents a decomposition of the variance of conventional-bond returns. The most striking feature of the results is that news about inflation is clearly the dominant influence, although revisions to future excess returns also play a significant role. Conversely real rate news has only a small and imprecisely determined effect.

Table 7.2.1
Variance decomposition for innovations in long-bond returns

VAR(order) Sample	VAR(1) 83:3-93:3	VAR(2) 83:4-93:3
Share of		
$var(\tilde{x}_r)$	0.011	0.009
	(0.014)	(0.010)
$var(\tilde{x}_\pi)$	1.000	1.100
	(0.342)	(0.585)
$var(\tilde{x}_x)$	0.610	1.045
	(0.378)	(0.628)
$2cov(\tilde{x}_r, \tilde{x}_\pi)$	-0.042	-0.049
	(0.095)	(0.147)
$2cov(\tilde{x}_r, \tilde{x}_x)$	0.091	0.090
	(0.107)	(0.154)
$2cov(\tilde{x}_\pi, \tilde{x}_x)$	-0.670	-1.195
	(0.664)	(1.135)

The real-rate result is the most straightforward to interpret. While short-term *ex-post* real rates are quite volatile, the market's expectation appears to be that they will quickly settle down to some normal level. This is consistent with the earlier interpretation of the summary statistics of Table 1, and with the real rate's small AR coefficient in the VAR. It also suggests that investors are reasonably content to accept an approximate version of the Fisher hypothesis.

The variance of inflation news appears to be almost equal to that of unexpected excess returns. However, in addition to this, revisions to expected excess returns contribute almost 70% of the latter. The reason for this significant 'excess' variance is that news about inflation and excess returns are strongly negatively correlated. When investors revise their expectations of inflation upward they revise down their expectations of future excess returns. In terms of current excess returns, the inflation news generates a negative return as the

[11] Heteroskedasticity-consistent standard errors in parentheses.

price of conventional bonds declines, but this is offset to some degree by an increase in prices due to the reduction in future returns. The second-order VAR offers identical qualitative results but the loss of degrees of freedom appears to result in somewhat less precise coefficient estimates.

These results are very similar to those obtained by Campbell and Ammer(1993) using US data, and probably reflect the similar nature of both the assets and the investors in both markets.

7.2.2 Innovations in short-term returns

The decomposition of unexpected movements in one-month rates is given in Table 7.2.2. Once again the results are very similar to those of Campbell and Ammer, although not as well determined. The results suggest that revisions to both one-month-ahead inflation and real interest rates are more variable than the nominal one-month rate itself. However, the covariance component again implies that when expected inflation is revised up, expected real rates are revised down, and vice versa. Thus the revisions tend to offset each other.

Table 7.2.2
Variance decomposition for innovations in short-rates

VAR(order) Sample	VAR(1) 83:3-93:3	VAR(2) 83:4-93:3
Share of		
$var(\tilde{x}_r)$	1.247	6.685
	(1.640)	(7.987)
$var(\tilde{x}_\pi)$	1.673	6.472
	(1.846)	(7.848)
$2cov(\tilde{x}_r, \tilde{x}_\pi)$	-1.920	-12.157
	(3.248)	(15.766)

7.2.3 Innovations in yield spreads

As with the decomposition of short rates, the long-short-spread decomposition is not very well determined. The point estimates, however, are as one would expect from the separate decompositions for the long and short-rates. In particular, revisions to real rates, which had no noticeable effect on the long-bond but which had a substantial effect on short-rates, also have a major impact on the spread. This confirms our earlier interpretation i.e. that while real-rate revisions may be substantial from month to month they are expected to decay within the life of the long bond.

Table 7.2.3
Variance decomposition of innovations in yield spreads

VAR(order) Sample	VAR(1) 83:3-93:3	VAR(2) 83:4-93:3
Share of		
$var(\tilde{x}_r)$	1.385	7.067
	(1.768)	(8.280)
$var(\tilde{x}_\pi)$	0.784	0.759
	(0.484)	(0.330)
$var(\tilde{x}_x)$	1.020	5.818
	(1.246)	(7.403)
$2cov(\tilde{x}_r, \tilde{x}_\pi)$	0.278	0.284
	(1.246)	(1.619)
$2cov(\tilde{x}_r, \tilde{x}_x)$	-1.876	-12.232
	(2.838)	(15.658)
$2cov(\tilde{x}_\pi, \tilde{x}_x)$	-0.591)	-0.686
	(1.406)	(1.674)

Conversely, revisions to expected inflation have an effect on both ends of the yield curve, although they explain more of the variation in short rates than in long rates. The results are consistent with, for example, an upward revision to expected inflation that raises both ends of the curve but which also causes it to invert i.e. the upward movement at the short end is the greater of the two. This suggests that revisions to expected inflation are also inclined to revert but not sufficiently quickly to leave the long bond unaffected. Any such reversion would clearly be significantly slower than that for real rates.

7.2.4 Interpreting the conventional-bond results: an incredible monetary policy

It seems that short rates are sensitive to both inflation and real-rate innovations, while long rates are affected only by the former. Thus an inflation innovation will cause the slope of the curve to change, and this change, given that inflation *expectations* appear to be quite persistent, will also be persistent. These results are also consistent with the estimated AR(1) coefficient for the nominal yield curve slope reported in Table 6.1.1.

This interpretation is, however, at variance with the AR(1) coefficient for inflation *per se*. With a coefficient of 0.2, inflation does not appear to be markedly more persistent than real interest rates, and these do not appear to move the long rate. Despite this similarity, if inflation were significantly more variable that real rates it could still explain a more significant proportion of long-rate variablility. However, one-month inflation is actually *less* variable than one-month real rates.

The explanation for the potency of inflation probably lies in the difference between the markets' view of the inflation process and the actual process. The AR(1) coefficient for inflation from 1800 to 1992, based on annual data, was about 0.5. For the periods 1945 to 1983, and 1945 to 1992 it was about 0.8[12]. Thus inflation shocks were considerably more persistent in the period prior to the start of our sample. It is not unreasonable to suppose that the markets based their expectations of inflation's persistence on the long run of data that was available at the start of our sample. Thus the speed at which inflation left the economy during the 1980s appears to have come as a surprise to the bond markets and suggests that the monetary policy stance adopted from the late-1970s was not credible. If this was indeed the case, it suggests that the high levels of unemployment experienced in the United Kingdom in the early 1980s were due to unexpectedly severe reductions in inflation. While several factors may have been responsible for the recession(s), the evidence here suggests that one of them may have been the authorities' failure to convince the financial markets of their intentions, and, by extension, their probable failure to convince any other markets.

7.3 A variance decomposition for index-linked bonds

The results of a variance decomposition of returns on real bonds are reported in Table 7.3. The remarkable result here is that news about real rates contributes less than 3% of the total return variance, and that this proportion is very imprecisely estimated. This result may appear surprising at first sight. It is, however, consistent with the explanation given for the absence of real rate news from the determination of returns on nominal bonds i.e. real rates are typically not expected to vary very much over a ten-year horizon. If changes in real-rates cannot account for movements in the prices of real bonds they cannot reasonably be expected to play a significant role in those of nominal bonds.

[12]The annual data coefficient for 1983 to 1992 was 0.2, just as for monthly data.

Table 7.3
Variance decomposition of innovations in index-linked-bond returns

VAR(order)	VAR(1)	VAR(2)
Sample	83:3-93:3	83:4-93:3
Share of		
$var(\tilde{x}_r)$	0.025	0.021
	(0.036)	(0.025)
$var(\tilde{x}_\pi)$	0.082	0.084
	(0.025)	(0.027)
$var(\tilde{x}_x)$	0.923	0.974
	(0.173)	(0.215)
$2cov(\tilde{x}_r, \tilde{x}_\pi)$	0.006	-0.001
	(0.025)	(0.036)
$2cov(\tilde{x}_r, \tilde{x}_x)$	0.207	0.183
	(0.141)	(0.190)
$2cov(\tilde{x}_\pi, \tilde{x}_x)$	-0.244	-0.262
	(0.109)	(0.133)

The results for real rates should be contrasted with those for inflation. Despite the fact that we are explaining the returns on *real* bonds, inflation news has a small but very well determined effect, accounting for about 8% of the return variance. The reason for this was suggested earlier. Index-linked bonds are not perfectly indexed, and their dependence on inflation reflects their exposure to price movements in the eight months prior to each 'real' payment. The dominant factor is, however, revisions to future excess returns. This should not be surprising, of course, given that investors expect real rates to be stable, and real bonds were designed to be as independent of inflation as possible; there is simply nothing else for the returns to depend upon.

The covariance terms echo the results for conventional bonds. The small role played by inflation leads to a similarly small role for the covariance of inlation and returns news. This is consistent with our earlier description of index-linked bonds as a portfolio of nominal and real bonds.

7.4 A variance decomposition for relative returns

Table 7.4

Variance decomposition of relative real- and nominal-bond returns

VAR(order)	VAR(1)	Var(2)
Sample	83:3-93:3	83:4-93:3
Share of		
$var(\tilde{x}_r)$	0.000	0.000
	(0.000)	(0.000)
$var(\tilde{x}_\pi)$	0.924	0.997
	(0.299)	(0.501)
$var(\tilde{x}_x)$	0.114	0.244
	(0.172)	(0.283)
$2cov(\tilde{x}_r, \tilde{x}_\pi)$	-0.002	-0.009
	(0.008)	(0.017)
$2cov(\tilde{x}_r, \tilde{x}_x)$	0.002	0.004
	(0.006)	(0.010)
$2cov(\tilde{x}_\pi, \tilde{x}_x)$	-0.040	-0.236
	(0.428)	(0.744)

The decomposition of relative returns provides a very clear result. About 93% of the variance in unexpected relative returns is due to revisions to expected future inflation. The remaining 7% is due to revisions to expected future relative returns. Real rate news has almost no effect.

The substantial role played by inflation news suggests that the majority of the movement in 'break-even' inflation rates at the ten-year horizon is in fact due to changes in expected inflation. Nevertheless, up to 7% of these movements is due to changes in required returns on index-linked and conventional bonds. This suggests that while such signal-extraction exercises can generate useful indications about how expectations of inflation are changing, there remains considerable scope for research into the behaviour of risk premia to improve the signal-to-noise ratio.

It is also interesting to note that while break-even rates at the ten-year maturity reflect a significant element of expected inflation, they do not do so for the intuitively attractive reason that the real-rate influences on nominal and real bonds 'cancel out'. The size of these influences is so small that whether or not they cancel is immaterial. Break-even rates 'work' because the *asset-specific premia* are strongly positively correlated.

8 Conclusions

This paper has used a vector autoregression, coupled with a development of Campbell and Ammer's (1993) dynamic accounting identity, to investigate the causes of unexpected movements in the prices of real and nominal bonds. In both cases, the candidate causes are revisions to expected real interest rates, inflation and required excess returns over the life of the bonds. This approach also allowed us to test the extent to which relative yields on real and nominal bonds can be used to draw inferences about market expectations of inflation. It is quite possible, in theory, that both the standard methods of extracting these expectations from yield data, and the more sophisticated method employed by the Bank of England, could be seriously undermined by the presence of asset-specific risk or liquidity premia.

Unexpected movements in the prices of long-term conventional bonds are dominated by revisions to expected inflation. Revisions to expected real rates have almost no role to play. Changes to expected premia contribute significantly to price changes but are strongly negatively correlated with news about future inflation. Since it is more likely that inflation causes premia, rather than the reverse, the dominant role of inflation news appears to be robust. The results suggest that if inflation expectations could be stabilised by a consistent and credible anti-inflation policy, the variance of bond excess returns could be reduced by up to 40%.

The results for short-bond returns and for the spread-portfolios suggest that bond market participants had little faith in the anti-inflation resolve of the Thatcher government. This lack of confidence was misplaced, however, as the persistence of inflation shocks appears to have been significantly lower than the bond markets expected. It seems that the markets continued to form their expectations of inflation in ignorance of the fact that there had been a regime change. It is reasonable to suppose that if the financial markets could make such a mistake a similar error could have been made by participants in the goods and labour markets. Thus the behaviour of the financial markets offers some support for the 'inflation surprise' interpretation of the recession of the early 1980s.

A surprising result is that real-rate news fails to influence the price of real bonds. This is probably due to an expectation on the part of investors that, whatever real rates may do in the short term, they will revert to a 'normal' level well within the life of the asset. Thus while a *permanent*,or at least, a *persistent*,change in real rates could have a substantial effect on the prices of real bonds, market participants do not expect such a change to occur. In short, there is not enough real-rate news available. The persistent absence of real-rate effects suggests that investors are content to use the Fisher hypothesis as the basis of their long-range interest rate forecasts.

Relative yields on real and nominal bonds appear to offer a reliable source of information about the markets' expectations of inflation. Approximately 95%

of the variance of previously unexpected changes in these relative yields is due to revisions to expected inflation. Thus, it seems that real bonds do offer a rich source of information about the way in which expectations of inflation change and, therefore, about the credibility of monetary policy.

Appendices

A Log-linear equations.

A.1 The general equation

The linearisation constants in equation (1) are $\rho = 1/(1+D^r/P^r)$, where D^r and P^r are the fixed points in the linearisation, and $k = -\{ln(\rho) + (1-\rho)ln(1/\rho - 1)\} * (1-\rho^m)/(1-\rho)$. The fixed points may be chosen in a number of different ways. The objective in this paper is to approximate the average values of D^r and P^r over the life of the asset.

A.2 Conventional bonds

In practice ρ was calculated as $\rho_c = 1/(1+\bar{Y})$, which is a reasonable approximation for bonds trading around par. The *nominal* yield is appropriate in this case as the real coupon and real price are each calculated by reference to the same goods price index. Further, with respect to the average real price of the bond over its life, this declines as expected inflation increases. Hence the nominal yield, which also reflects inflation, is the appropriate rate.

Since the nominal coupon, c_c^n, is a known constant in the case of conventional bonds, it can be added to the linearisation constant as $k_c' = k_c + c_c^n(1-\rho_c^m)$.

A.3 Index-linked bonds

The constant ρ can again be approximated by the sample average par yield except that for index-linked bonds the real yield is used (recall that we require an approximation to the ratio D^r/P^r). A consequence of indexation is that changes in the goods price level alter both the nominal coupon and the nominal price of the bond. Thus the real coupon-price ratio is approximately constant regardless of inflation, which would alter the bond's nominal yield.

The nominal coupon in the case of index-linked bonds changes over time. However, it can still be added to the constant to generate a price equation as,

$$p_{g,t}^m = k_{g,t}^* - E_t \sum_{i=0}^{m-1} \rho_g^i h_{g,m-i,t+1+i}^n + E_t \sum_{i=0}^{m-1} \rho_g^i \pi_{g,t+1+i-l} \tag{29}$$

where $k_{g,t}^* = k_g + c_g(1-\rho_g) + z_{t-l} - \bar{z}$ which clearly changes over time. A further simplification derives from the fact that the first l elements of the π summation are known at time t. Hence,

$$p_{g,t}^m = k_{g,t}' - E_t \sum_{i=0}^{m-1} \rho_g^i h_{g,m-i,t+1+i}^n + E_t \sum_{i=l}^{m-1} \rho_g^i \pi_{g,t+1+i-l} \tag{30}$$

where

$$k'_{g,t} = k^*_{g,t} + \sum_{i=0}^{l-1} \rho_g^i \pi_{t+1+i-l} \tag{31}$$

Since $(E_{t+1} - E_t)k'_{g,t+1} = 0$ this 'constant' will not contribute to the unexpected excess return calculations.

B Data sources

All end-month interest rates and yields were provided by the Bank of England, as were the prices of index-linked bonds.

C Month-average data

C.1 Sources

Nominal interest rates are those publised in Financial Statistics. The codes are; ten-year yield (AJLW), three-month Treasury bill yield (AJNC), one-month interbank rate (VNEA). The prices of index-linked bonds were supplied by the Bank of England; their returns are calculated on an end-month basis.

C.2 Variance decompositions

Table 6
Variance decomposition for month-average data

	Conventional		Indexed		Relative	
VAR order	(1)	(2)	(1)	(2)	(1)	(2)
Sample	83:3-93:3	83:4-93:3	83:3-93:3	83:4-93:3	83:3-93:3	83:4-93:3
Share of						
$var(\tilde{x}_r)$	0.033	0.028	0.049	0.037	0.000	0.000
	(0.042)	(0.026)	(0.066)	(0.038)	(0.000)	(0.000)
$var(\tilde{x}_\pi)$	1.074	1.581	0.070	0.064	0.850	1.192
	(0.337)	(0.642)	(0.022)	(0.019)	(0.262)	(0.494)
$var(\tilde{x}_x)$	0.833	1.519	0.994	1.013	0.213	0.285
	(0.481)	(0.872)	(0.195)	(0.223)	(0.210)	(0.315)
$2cov(\tilde{x}_r, \tilde{x}_\pi)$	0.001	-0.068	0.006	-0.014	-0.009	-0.014
	(0.166)	(0.214)	(0.027)	(0.026)	(0.015)	(0.020)
$2cov(\tilde{x}_r, \tilde{x}_x)$	-0.039	-0.014	0.093	0.107	0.008	0.009
	(0.225)	(0.285)	(0.206)	(0.225)	(0.013)	(0.015)
$2cov(\tilde{x}_\pi, \tilde{x}_x)$	-0.9031	-2.046	-0.212	-0.205	-0.061	-0.472
	(0.764)	(1.413)	(0.101)	(0.107)	(0.404)	(0.774)

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