PREDICTING INFLATION WITH THE TERM STRUCTURE SPREAD

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Abstract: It is tempting to interpret empirical evidence in a number of recent studies as suggesting that term structure spreads help predict future inflation over moderate horizons of 3 to 5 years. This paper argues that commonly measures of the predictive power of the term structure spread for future inflation are misleading. In particular, R^2s for estimated inflation-change equations can drastically overstate the predictive power of spreads. The paper explains why the overstatement is likely to be particularly large in countries whose monetary authorities have strong reputations for credibly targeting a stable inflation rate. Results from an empirical analysis of data from eleven industrialized countries suggest that the level of the short-term real rate may be more useful for predicting inflation than the term structure spread, possibly because changes in short-term real rates provide clearer measures of changes in the stance of monetary policy.

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1 Introduction

Forecasts of inflation play an important role in a wide range of economic decisions, including investment decisions, the government budgeting process, wage negotiations, and monetary policy decisions. One potentially useful source of information on expected inflation is the term structure of interest rates. In fact, a number of articles report that the spread between the yields on relatively long and relatively short government securities help predict inflation at moderate to long horizons. This paper revisits the question of whether the spread helps predict inflation.

Several hypotheses relate the term structure spread to future inflation. One hypothesis, as recently exposited in Estrella and Mishkin (1997), holds that the term structure spread is an indicator of the stance of monetary policy. According to this view, a low spread reflects relatively restrictive monetary policy because the spread is low when short-term interest rates are high relative to long-term interest rates. Thus, a low spread predicts that in response to the contractionary monetary policy, real activity will slow and inflation will decrease. A second hypothesis decomposes the term structure spread into three components: the expected real rate change, the expected inflation change, and the term premium. According to this view, if variation in the term structure spread is largely due to variation in expected inflation changes, then the term structure spread will help predict inflation changes. Fama (1990), Jorion and Mishkin (1991), Estrella and Mishkin (1997), and Day and Lange (1997) have found empirical evidence suggesting that at moderate to long horizons of 3 to 5 year, the term structure spread helps predict inflation changes.

One issue investigated in this paper is whether empirical results in some previously published studies actually provide evidence to support the claims that the term structure spread helps predict inflation. In particular, many empirical studies estimate inflation-change equations in which the difference between ex poste future inflation and current inflation is regressed on a constant and a term structure spread. These studies then use tests of the statistical significance of the estimated coefficient on the spread and the R^2 from the regression to provide evidence on the predictive power of the term structure spread for the future path of inflation. This paper argues that such an approach is flawed.

The flaw in the typical approach can be traced to the specification of the dependent variable in the standard inflation-change equation. The dependent variable is the difference between future inflation (measured ex poste) and current inflation. Consequently, regression R^2s describe the proportion of the variability in the *change* in inflation explained by the regressors, not the proportion of the variability in the *level* of future inflation. For moderate to long horizons, the less persistent is inflation, the less likely these R^2s are to reflect the ability of the regressors to predict future inflation and the more likely they are to reflect correlation between the regressors and current inflation. Similar problems arise when trying to interpret the meaning of statistical significance of coefficient estimates. For moderate to long horizons, the less persistent is inflation, the less likely statistically significant coefficient estimates are to reflect the ability of regressors to predict inflation and the more likely they are to reflect correlations between the regressors and current inflation.

The paper shows that the problems with standard interpretations are likely to be more severe for countries with historically stronger reputations as inflation hawks. This result obtains because monetary policies which credibly target a low and stable inflation rate tend to result in less persistent inflation processes. And, the less persistent is inflation, the greater are the problems with standard interpretations of estimated inflation-change equations.

The paper proposes alternative specifications of inflation equations that are not susceptible to misinterpretations. These specifications are used as the basis of a multi-country investigation of the predictive power of the spread for future inflation.

A second issue investigated in the paper is whether alternative indicators of the stance of monetary policy provide better information on future inflation. The spread is the difference between a long-term rate and a short-term rate. While the term structure spread may contain information on the stance of monetary policy, as noted by Estrella and Mishkin (1997), central banks can influence the term structure but cannot control it in any meaningful sense. Long-term rates fluctuate with market expectations about future short-term rates and variation in liquidity or term premiums. Short-term rates are more closely connected to the primary instrument of policy.[1] Thus, if the predictive power of the spread for inflation comes from its role as an indicator of monetary policy, variables such as short-term real rates that move more closely with policy changes may be better predictors of inflation. To investigate this issue, the paper compares the information contents of the spread and the short-term real rate for future inflation.

The next section of the article reviews two hypotheses of why term structure spreads might

[1] For more discussion see Estrella and Mishkin (1997).

help predict future inflation. The third section discusses the use of inflation-change equations to evaluate the predictive content of term structure spreads for future inflation. In particular, section 3 shows that regression R^2s and evaluations of the statistical significance of estimated coefficients from inflation-change equations can provide misleading assessments of the information in the term structure spread for future inflation. Empirical results from an analysis of quarterly data for Australia, Canada, France, Germany, Italy, Japan, the Netherlands, Sweden, Switzerland, the United Kingdom, and the United States are examined in the fourth section. The fifth section concludes.

2 Why might term structure spreads help predict inflation?

Economists often look to the term structure as a potential source of information about future economic conditions. A number of recent articles, including Harvey (1989), Estrella and Hardouvelis (1991), Estrella and Mishkin (1996 and 1997), Plosser and Rouwenhorst (1994), Bonser-Neal and Morley (1997), and Kozicki (1997), have investigated the predictive power of the spread for future real activity. The predictive power of the spread for future inflation has also been investigated in articles by Fama (1990), Mishkin (1990 and 1991), Jorion and Mishkin (1991), Abken (1993), Blough (1994), Frankel and Lown (1994), Engsted (1995), Gerlach (1995), Tzavalis and Wickens (1996), Alles and Bhar (1997), Davis and and Fagan (1997), Day and Lange (1997) and Kozicki (1997). This article contributes to the latter literature.

The term structure has attracted attention as an indicator because, according to the Expectations Hypothesis, the term structure is forward-looking, reflecting market expectations about future economic conditions. This section summarizes two theories based on the Expectations Hypothesis that justify why term structure spreads may help predict future inflation.

The Expectations Hypothesis relates the yield on longer-term financial instruments to expected future yields on short-term instruments. Following Campbell and Shiller (1991), in the case of pure discount bonds,

$$R_{n,t} = (1/k) \sum_{i=0}^{k-1} E_t R_{m,t+mi} + c_{n,m,t}, \qquad k = n/m \tag{1}$$

where $R_{n,t}$ is the nominal yield on an n-period bond and $c_{n,m,t}$ is a term premium that may vary with n, m, and t.[2] Often the Expectations Hypothesis is expressed with $m = 1$. Defining

[2]Campbell and Shiller assumed that the term premium was constant through time.

$c_{n,t} \equiv c_{n,1,t}$ and the short rate as $R_t \equiv R_{1,t}$, (1) states that the yield on an n-period bond is equal to a term premium plus the average of expected short rates up to $n-1$ periods in the future.

One theory argues that the term structure spread should help predict inflation because inflation responds to monetary policy actions and the term structure spread reflects the stance of monetary policy. The intuition behind this view can be explained by rewriting the Expectations Hypothesis in (1) with $m = 1$:

$$R_{n,t} = (1/n) \sum_{i=0}^{n-1} E_t R_{t+i} + c_{n,t}. \tag{2}$$

This expression equates long-term yields to an average of expected future short-term yields plus a risk premium. The first term on the right side of ths expression reflects what market participants expect short rates to average in the future. For large n this average should smooth most of the cyclical variation in expected short rates.[3] Intuitively, because long rates reflect average expected short rates over a relatively long time interval, long-term yields may provide a reasonable benchmark against which current short-term yields can be compared. The view that monetary policy is relatively tight may be reflected by short rates that are relatively high compared to long rates, i.e., by a term structure spread that is small or negative. Likewise, the view that monetary policy is relatively accommodative may be reflected by short rates that are relatively low compared to long rates, i.e., by a large term structure spread.

A second observation noted by analysts who believe the term structure spread reflects the stance of monetary policy is that the the spread usually falls when monetary policy is tightened. Short-term yields move closely with the interest rate that serves as an instrument of monetary policy (i.e., the federal funds rate in the United States). When monetary policy is tightened, short-term interest rates rise. Although long-term yields may react to policy, they rarely rise one-for-one with short-term interest rate increases. As a result, the term structure spread usually falls when monetary policy is tightened.

Under this policy-stance theory, the term structure spread helps predict inflation because it reflects the current stance of monetary policy and economic variables respond to monetary policy actions. Accordingly, a low or negative term structure spread predicts that in response to the relatively tight stance of monetary policy, real activity will slow and inflation will decrease.

[3] The variance of the average over $n > 1$ periods of a stationary random variable is less than the variance of the random variable before averaging. For I(1) random variables, however, this property does not hold.

Conversely, a large term structure spread predicts that in response to the relatively accommodative monetary policy, real growth will pickup and inflation will increase.

If the term structure spread holds predictive power for inflation primarily because it reflects the stance of monetary policy, then other measures of the stance of monetary policy should also help predict inflation. In particular, Bernanke and Blinder (1992) argue that changes in the federal funds rate reflect changes in the stance of monetary policy. More generally, short-term interest rates that move closely with the interest rate that serves as the instrument of monetary policy should also reflect the stance of monetary policy. While fluctuations in the term structure spread may reflect shifts in policy, they may also be caused by shifts in risk premium. Thus, short term interest rates may provide a better measure of the stance of monetary policy and may be a better predictor of future inflation. The question of whether the yeild spread or short term interest rates provide better forecasts of inflation is investigated in section 4.

A second theory of why term structure spreads should help predict inflation argues that the spread reflects the direction of future inflation changes. Support for this view also follows from the Expectations Hypothesis. To see how the expectations hypothesis implies a relationship between longer-term yields and expected inflation, replace short-term nominal rates in (1) with the sum of expected short-term real rates and expected inflation. The resulting expression for long-term yields is

$$\begin{aligned} R_{n,t} &= (1/n)\sum_{i=0}^{n-1} E_t(r_{t+i} + \pi_{t+1+i}) + c_{n,t} \\ &= (1/n)\sum_{i=0}^{n-1} E_t r_{t+i} + (1/n)\sum_{i=0}^{n-1} E_t \pi_{t+1+i} + c_{n,t} \end{aligned} \tag{3}$$

where r_t is a one-period real rate and π_{t+1} is a one-period inflation rate from t to $t+1$. Using (3), the term structure spread, constructed as the difference between the yield on an n-period bond and a one-period bond, is equal to

$$R_{n,t} - R_t = (1/n)\sum_{i=0}^{n-1} E_t(r_{t+i} - r_t) + (1/n)\sum_{i=0}^{n-1} E_t(\pi_{t+1+i} - \pi_{t+1}) + c_{n,t} \tag{4}$$

Thus, the term structure spread indicates the direction of expected real rate changes $((1/n)\sum_{i=0}^{n-1} E_t(r_{t+i} - r_t))$, the direction of expected inflation changes $((1/n)\sum_{i=0}^{n-1} E_t(\pi_{t+1+i} - \pi_{t+1}))$ and the term premium $(c_{n,t})$. If variation in the term structure spread is largely due to variation in expected inflation changes, then the term structure spread will help predict inflation changes if market expectations are correct on average.

3 Evaluating Predictive Content

Several recent articles use simple regressions to assess the information content of the term structure spread for future inflation. This section outlines the standard format of the "inflation-change equations" estimated in these articles and reviews how researchers have typically interpreted regression results. Two subsections explore in detail the pitfalls associated with these interpretations. The third subsection shows that problems are likely to be more severe in countries that have a reputation for credibly targeting a low and stable inflation rate.

The standard "inflation-change equation" estimated by researchers interested in assessing whether the term structure spread helps predict the future path of inflation, takes the form

$$\pi_{k,t+1} - \pi_{m,t+1} = \alpha_{k,n,m} + \beta_{k,n,m}(R_{n,t} - R_{m,t}) + resid_{t+1} \qquad (5)$$

where $\pi_{k,t+1} \equiv (1/k)\sum_{i=0}^{k-1} \pi_{t+1+i}$ is the k-period inflation rate from t to $t+k$, and $resid_{t+1}$ is the regression residual.[4] Miskhin (1990) set $k = n$ and considered (n, m) pairings of 3 and 1 months, 6 and 3 months, 9 and 6 months, 12 and 6 months, and 12 and 9 months. Fama (1990) set m to 1 year, n to 5 years, and allowed k to take on values of 2, 3, 4 or 5 years. Mishkin (1991) also set $k = n$, but considered (n, m) pairings of 3 and 1 months, 6 and 1 months, 6 and 3 months, 12 and 1 months, 12 and 3 months, and 12 and 6 months. Jorion and Mishkin (1991) and Day and Lange (1997) set $k = n$, m to 1 year, and allowed n to take on values of 2, 3, 4, or 5 years. Gerlach (1995) set $k = n$, m to 2 or 3 years, and allowed n to take on values from $m + 1$ to 10 years. Tzavalis and Wickens (1996) added a monetary regime dummy to account for a break in October 1980 and examined (n, m) for $n > m$ where n and m took values of 1, 3, 6, and 12 months.

Most studies *interpret* regression R^2s and estimates of the coefficient $\beta_{k,n,m}$ as suggesting how much information the nominal term structure contains about future inflation and the real term structure. First, regression R^2s are interpreted as providing information on forecast power for changes in inflation. Second, for $k = n$, the amount of information for future inflation and real rates in the term structure spread is assessed from statistical tests of whether estimates of the

[4] To reduce notation, $resid_{t+1}$ will be used to represent the regression residual in most regression equations. This should not be taken to imply that all regression equations have the same residuals.

coefficient $\beta_{n,n,m} \equiv \beta_{n,m}$ are significantly different from zero or one.[5] In particular, Mishkin (1990) writes:

> ... a statistical rejection of $\beta_{n,m} = 0$ provides evidence that (a) the term structure contains significant information about the future path of inflation and (b) the slopes of the term structures of real and nominal interest rates do not move one-for-one with each other. On the other hand, a statistical rejection of $\beta_{n,m} = 1$ provides evidence that (a) the slope of the real term structure is not constant over time and (b) the term structure of nominal interest rates provides information about the term structure of real interest rates.

The remainder of the section will show that interpretations based on regression R^2s and estimates of $\beta_{n,m}$ from the standard inflation-change equation can provide misleading assessments of the information in the term structure spread for future inflation. The development will use simplified notation, taking $m = 1$ and letting $k = n$, although similar results would hold for other values of m and for $k \neq n$.[6] For $m = 1$ and $k = n$, the inflation-change equation becomes

$$\pi_{n,t+1} - \pi_{t+1} = \alpha_n + \beta_n(R_{n,t} - R_t) + resid_{t+1} \tag{6}$$

where $\alpha_n \equiv \alpha_{n,n,1}$, $\beta_n = \beta_{n,n,1}$, and $\pi_{t+1} \equiv \pi_{1,t+1}$. The next two subsections discuss, respectively, the use of regression R^2s to assess predictive power and information revealed by estimates of β_n. The third subsection illustrates why extra care should be taken in analyses of countries with longstanding reputations for seeking to maintain stable inflation rates.

3.1 Interpreting R^2 of the inflation-change equation

Assessments of the information in the term structure spread for future inflation based on regression R^2s may be misleading.[7] R^2s provide a measure of the proportion of variation in the dependent variable of a regression explained by variation in the regressors. Most studies examining the

[5] Tzavlis and Wickens (1996) assert that assuming rational expectations implies $\beta = 1$ and compare estimates of β to "its predicted value of unity."

[6] As long as n is large relative to m the same results would hold. Under realistic assumptions about the inflation process, similar results could also be derived for large k and n but with $k \neq n$.

[7] The term R^2 is used to refer to the ratio of the variance of regression predictions of the dependent variable to the variance of the dependent variable. The paper does not distinguish between the variance of a random variable and estimates of the variance. The arguments put forward in this section also hold for measures such as $\bar{R}^2$.

predictive power of the spread for future inflation base their assessments on R^2s for regressions that have inflation changes, $(\pi_{n,t+1} - \pi_{t+1})$ as the dependent variable. Thus, as written, the regressions do not distinguish between the ability of the explanatory variables to explain variation in the path of future inflation, $\pi_{n,t+1}$, from their ability to explain variation in current inflation, π_{t+1}. A high R^2 could be merely signalling significant correlation between the term structure spread and current inflation. In other words, a high R^2 could be obtained even in situations when the term structure spread does not help predict future inflation.

An example will be used to illustrate situations when the term structure spread does not help predict future inflation, but for which high R^2s could be obtained during estimation of inflation change equations. For simplicity assume inflation follows a mean-reverting AR(1) process,

$$\pi_{t+1} = (1-\gamma)\bar{\pi} + \gamma\pi_t + \epsilon_{t+1} \tag{7}$$

where $\bar{\pi}$ is the mean of the inflation process, $0 \leq \gamma < 1$ is a measure of inflation persistence, and ϵ_t is an iid disturbance distributed with mean equal to zero and variance σ_ϵ^2.[8] Then the optimal forecast of $\pi_{n,t+1}$, using information available in t, is

$$\begin{aligned} E_t\pi_{n,t+1} &= (1/n)\sum_{i=1}^{n}[\gamma^i\pi_t + (1-\gamma^i)\bar{\pi}] \\ &= \left[(1/n)\sum_{i=1}^{n}\gamma^i\right]\pi_t + \left[1-(1/n)\sum_{i=1}^{n}\gamma^i\right]\bar{\pi} \end{aligned} \tag{8}$$

Thus, optimal forecasts of $\pi_{n,t+1}$ are a weighted average of the most recently observed one-period inflation rate, π_t, and the mean of the inflation process, $\bar{\pi}$. For low inflation persistence, i.e., small values of γ, or long horizons n, the optimal forecast of $\pi_{n,t+1}$ approaches the constant mean $\bar{\pi}$ and the weight on π_t approaches zero.

In this simple example, the ability of any regressor to help predict inflation is very limited for low inflation persistence or long forecast horizons. In a regression of $\pi_{n,t+1}$ on a constant, π_t, and a set of other regressors uncorrelated with π_t, say X_t, the estimator of the constant will be distributed with mean $[1-(1/n)\sum_{i=1}^{n}\gamma^i]\bar{\pi}$, the estimator of the coefficient on π_t will be distributed with mean $(1/n)\sum_{i=1}^{n}\gamma^i$, and estimators of coefficients on other variables will be distributed with

[8] The next subsection will provide a simple model that relates $\bar{\pi}$ to the rate of inflation targeted by monetary policy. The model will allow for time variation in the inflation target.

mean zero.[9] For low inflation persistence or long horizons, the mean of the estimator of the constant will approach the mean of the inflation process, and the mean of the estimator of the coefficient on π_t will approach zero.

The regression R^2 from a regression of future inflation on a constant and current inflation reflects the ratio of the regression predictions of $\pi_{n,t+1}$ to the variance of $\pi_{n,t+1}$:

$$
\begin{aligned}
R^2(\pi_{n,t+1}) &= \frac{\sigma^2((1/n)\sum_{i=1}^{n}\gamma^i\pi_t + (1-(1/n)\sum_{i=1}^{n}\gamma^i)\bar{\pi})}{\sigma^2(\pi_{n,t+1})} \\
&= \frac{[\sum_{i=1}^{n}\gamma^i]^2[\sum_{i=0}^{\infty}\gamma^{2i}]}{\left[(\sum_{j=0}^{n-1}(\sum_{i=0}^{j}\gamma^i)^2) + (\sum_{i=1}^{n}\gamma^i)^2(\sum_{i=0}^{\infty}\gamma^{2i})\right]} \qquad (9)
\end{aligned}
$$

where $\sigma^2(x)$ is used to denote the variance of x. The last line follows after recognizing that, for the AR(1) inflation process in (7),

$$
\begin{aligned}
\sigma^2(\pi_t) &= \sigma_\epsilon^2 \sum_{i=0}^{\infty}\gamma^{2i} \\
\sigma^2(\pi_{n,t+1}) &= \frac{\sigma_\epsilon^2}{n^2}\left[(\sum_{j=0}^{n-1}(\sum_{i=0}^{j}\gamma^i)^2) + ((\sum_{i=1}^{n}\gamma^i)^2\sum_{i=0}^{\infty}\gamma^{2i})\right] \\
cov(\pi_{n,t},\pi_t) &= \frac{\sigma_\epsilon^2}{n}\left[(\sum_{i=1}^{n}\gamma^i)(\sum_{i=0}^{\infty}\gamma^{2i})\right] \qquad (10)
\end{aligned}
$$

For $0 \leq \gamma \leq 1$, the value $\gamma = 0$ minimizes $R^2(\pi_{n,t+1})$ with $R^2(\pi_{n,t+1}) = 0$ for $\gamma = 0$, and $R^2(\pi_{n,t+1})$ is monotonically increasing in γ with $\lim_{\gamma\to 1} R^2(\pi_{n,t+1}) = 1$. Furthermore, for given γ, $R^2(\pi_{n,t+1})$ is decreasing in n. In other words, when inflation is not persistent, $R^2(\pi_{n,t+1})$ is close to zero for long forecast horizons. This result would be obtained even if additional variables, such as the short term rate or the yield spread, were included as regressors. Intuitively, the result reflects that the non-constant regressors do little to help explain the variability of future long-horizon inflation ($\pi_{n,t+1}$ for large n) when inflationary shocks are not persistent.[10]

The previous paragraph illustrated situations in which low R^2s could be obtained from regressions with future inflation as the dependent variable. This paragraph shows that in these cases it is still possible to obtain a high R^2 during estimation of an inflation-change equation, i.e.,

[9] The constraint that X_t be uncorrelated with π_t is not restrictive since π_t is included as an explanatory variable. As long as π_t is included as an explanatory variable, if regressors correlated with π_t were to be included, a transformation of variables could be performed, without loss of generality, to generate new regressors including π_t and a set of regressors uncorrelated with π_t.

[10] According to (7), when $\gamma = 0$, inflation is an iid distributed random variable with mean $\bar{\pi}$ and variance σ_ϵ^2. In this case, the best forecast of future inflation is the constant $\bar{\pi}$.

a regression of $\pi_{n,t+1} - \pi_{t+1}$ on a constant and a collection of other regressors. Continuing to assume that inflation follows the AR(1) process in (7), suppose that the inflation-change variable, $(\pi_{n,t+1} - \pi_{t+1})$ is regressed on a constant, and a set of other variables, Y_t.

$$\pi_{n,t+1} - \pi_{t+1} = a + b\,Y_t + e_t \tag{11}$$

For large n and small γ estimates of the coefficients on the non-constant regressors would basically reflect only the correlations between these regressors and current inflation (actually, the negative of current inflation, $-\pi_{t+1}$), because, as derived above, non-constant regressors do little to help predict future inflation $\pi_{n,t+1}$. The regression R^2 from this regression, however, would reflect the proportion of the variation in $\pi_{n,t+1} - \pi_{t+1}$ explained by the regressors. Let $\sigma^2(\hat{e})$ be the variance of the residuals from this regression. Then, the R^2 from this regression is equal to

$$\begin{aligned} R^2(\pi_{n,t+1} - \pi_{t+1}) &= 1 - \frac{\sigma^2(\hat{e})}{\sigma^2(\pi_{n,t+1} - \pi_{t+1})} \\ &= 1 - \frac{\sigma^2(\hat{e})}{\sigma^2(\pi_{n,t+1}) + \sigma^2(\pi_{t+1}) - 2cov(\pi_{n,t+1}, \pi_{t+1})} \end{aligned} \tag{12}$$

and, could be sizable. For instance, if $\gamma = 0$, then $R^2(\pi_{n,t+1} - \pi_{t+1}) = 1 - \sigma^2(\hat{e})/\sigma^2_\epsilon$ which would be comparable to the R^2 from the regression of π_{t+1} on a constant and Y_t. In this case, if Y_t is correlated with π_{t+1} then a large value of $R^2(\pi_{n,t+1} - \pi_{t+1})$ could be obtained.

Summarizing, this section has illustrated that for long-horizon forecasts, when inflation persistence is low, a high $R^2(\pi_{n,t+1} - \pi_{t+1})$ and a low $R^2(\pi_{n,t+1})$ can be simultaneously obtained. Clearly, in such situations interpreting a large value of $R^2(\pi_{n,t+1} - \pi_{t+1})$ as suggesting that the regressors help predict future inflation, $\pi_{n,t+1}$ would be inappropriate. This exercise has shown that assessments on the usefulness of variables to help predict the future path of inflation can be misleading when based on R^2s from regressions with inflation changes as the dependent variable.

One approach to determine the usefulness of the variables to predict future inflation is to compare the variance of the residuals from the inflation-change regression (11), $\sigma^2(\hat{e})$, to the variance of $\pi_{n,t+1}$ instead of to the variance of $\pi_{n,t+1} - \pi_{t+1}$. Such a comparison can be based on an *alternative* R^2, calcuated as

$$Alt.\ R^2(\pi_{n,t+1}) = 1 - \frac{\sigma^2(\hat{e})}{\sigma^2(\pi_{n,t+1})}. \tag{13}$$

This alternative R^2 is equal to the R^2 from the regression

$$\pi_{n,t+1} = a + bY_t + c\pi_{t+1} + e_t \tag{14}$$

with the coefficient on π_{t+1} restricted to equal one, i.e., $c = 1$.

3.2 Interpreting estimates of β_n

Assessments on the usefulness of the term structure spread to help predict the future path of inflation also may be misleading when based on estimates of the coefficient β_n. This subsection shows that a statistical rejection of $\beta_n = 0$ does not necessarily provide evidence that the term structure contains significant information about the future level of inflation.[11]

Under rational expectations, n-period bond yields will equal

$$R_{n,t} = E_t r_{n,t} + E_t \pi_{n,t+1} + c_{n,t} \tag{15}$$

where $r_{n,t}$ is the n-period real yield defined according to

$$\begin{aligned} r_{n,t} &= E_t r_{n,t} + \nu_{n,t+1} \\ &= (1/n) \sum_{i=0}^{n-1} E_t r_{t+i} + \nu_{n,t+1}, \end{aligned} \tag{16}$$

the n-period inflation rate equals

$$\pi_{n,t+1} = E_t \pi_{n,t+1} + \epsilon_{n,t+1}, \tag{17}$$

and $\nu_{n,t+1}$ and $\epsilon_{n,t+1}$ are expectational errors. The inflation-change equation can be rewritten

$$\pi_{n,t+1} - \pi_{t+1} = \alpha_n + \beta_n \left[E_t(r_{n,t} - r_t) + E_t(\pi_{n,t+1} - \pi_{t+1}) + c_{n,t} \right] + resid_{t+1} \tag{18}$$

where the term structure spread has been decomposed into 3 components: the expected real rate change, $E_t(r_{n,t} - r_t)$, the expected inflation change, $E_t(\pi_{n,t+1} - \pi_{t+1})$, and the risk premium, $c_{n,t}$. A general expression for β_n can be derived as

$$\beta_n = \frac{\rho\tilde{\sigma} + \tilde{\sigma}^2 + \rho_c \tilde{\sigma} \tilde{\sigma_c}}{1 + \tilde{\sigma}^2 + \tilde{\sigma_c}^2 + 2\rho\tilde{\sigma} + 2\rho_c \tilde{\sigma} \tilde{\sigma_c} + 2\rho_{rc} \tilde{\sigma_c}} \tag{19}$$

where $\tilde{\sigma}$ is the ratio of the standard deviation of $E_t(\pi_{n,t+1} - \pi_{t+1})$ to the standard deviation of the expected real rate changes, $E_t(r_{n,t} - r_t)$, ρ is the correlation between $E_t(\pi_{n,t+1} - \pi_{t+1})$ and $E_t(r_{n,t} - r_t)$, $\tilde{\sigma_c}$ is the ratio of the standard deviation of $c_{n,t}$ to the standard deviation of $E_t(r_{n,t} - r_t)$,

[11] In some situations, a statistical rejection of $\beta_n = 0$ may provide evidence that the term structure contains information about the *change* in inflation.

ρ_c is the correlation between $E_t(\pi_{n,t+1}-\pi_{t+1})$ and $c_{n,t}$, and ρ_{rc} is the correlation between $E_t(r_{n,t}-r_t)$ and $c_{n,t}$.[12]

A non-zero β_n need not imply that the term structure spread helps predict future inflation for the same reasons that a high $R^2(\pi_{n,t+1}-\pi_{t+1})$ may be misleading. As in the previous subsection, difficulties arise when inflation is not highly persistent and n is large. Once again, suppose inflation follows the stochastic process described in (7). If inflation persistence is relatively low, then as maturity n lengthens, variation in $E_t(\pi_{n,t+1}-\pi_{t+1})$ will be dominated by variation in π_{t+1}. In such circumstances, non-zero estimates of β_n would signal correlations between current inflation and at least one of the the three components of the term structure spread, i.e., at least one of the expected real rate change, the expected inflation change, or the risk premium. One extreme case is particularly noteworthy. Estimates of β_n would be close to one, for instance, if variation in the expected real rate change and the risk premium were small relative to variation in π_{t+1}. In this situation, a unit coefficient estimate of β_n, however, could merely be signalling that variation in both the dependent variable $(\pi_{n,t+1}-\pi_{t+1})$ and the term structure spread $(R_{n,t}-R_t)$ are dominated by variation in $(-E_t\pi_{t+1})$. In other words, the unit coefficient estimate of β_n need not imply that the term structure spread helps predict inflation, it may be reflecting that the term structure spread is correlated with current inflation.

Sizable correlations between regressors and current inflation may lead to misinterpretations of estimation results for inflation-change equations. As acknowledged in previous studies, in the presence of correlation between the expected real rate change, the expected inflation change, and the term premium, a wide variety of values of β_n are consistent with the Expectations Hypothesis. Given the acknowledgement for such correlations, it is surprising that many previous studies didn't also allow for correlations between current inflation and one or more of the components of the term structure spread.[13] To allow for such correlations, the restriction on the coefficient on current inflation should be relaxed. A more general inflation equation, justified by such an argument is

$$\pi_{n,t+1} = \alpha_n + \beta_n(R_{n,t}-R_t) + \gamma_n\pi_{t+1} + resid_{t+1}. \tag{20}$$

[12] Mishkin (1990) derived a similar expression, but combined the terms for expected real rate changes and the term premium into one, $E_t(r_{n,t}+c_{n,t}-r_t)$, which he called the slope of the real term structure.

[13] Fama (1990) and Day and Lange (1997) allowed for such correlations by including current inflation as a regressor in inflation-change equations. However, because estimated regressions had the inflation change variable, $(\pi_{n,t+1}-\pi_{t+1})$ as the regressand, regression R^2 still report the variation in inflation changes explained by the regressors and don't directly address how useful the regressors are for predicting $\pi_{n,t+1}$.

This equation explicitly examines the ability of the term structure spread to predict future inflation. Under the restriction $\gamma_n = 1$, the R^2 from this equation is equal to the alternative R^2, referred to as *Alt.* $R^2(\pi_{n,t+1})$, introduced in section 3.1. The restriction $\gamma_n = 1$ can be relaxed and coefficients can be estimated using instrumental variables techniques. Given the high degree of autocorrelation in inflation, a natural instrument choice for π_{t+1} is π_t.[14]

3.3 Relating Regression Interpretations to Monetary Policy

This section will use a simple model to intuitively illustrate why inflation persistence may be related to monetary policy actions. The model will show that policy actions designed to achieve a stable rate of inflation over time result in inflation processes that are less persistent, while policy actions which allow inflation goals to change may result in more persistent inflation processes. In the current paper, the relationship between long-run constancy of the inflation target of policy and inflation persistence is important because, combined with the discussion in the previous two subsections, it suggests that researchers should be particularly mindful of the interpretation of inflation-change regression results when analysis is conducted on countries which have followed policies targeting constant inflation rates.

The model has two equations. The first represents the observed empirical relationship between inflation and monetary policy actions. Inflation from t to $t+1$, π_{t+1}, is assumed to increase (decrease) if the real interest rate in the previous period, $r_t - E_t\pi_{t+1}$ was less (greater) than the long-run equilibrium real rate $\bar{rr}$:

$$\pi_{t+1} = \pi_t - \theta(r_t - E_t\pi_{t+1} - \bar{rr}) + \upsilon_{t+1} \tag{22}$$

where υ_{t+1} is an independent zero mean shock, and $\theta > 0$ is the responsiveness of inflation to real interest rates.[15] The second equation describes monetary policy decisions. Policy makers are

[14]Generally, inflation is sufficiently persistent that the inflation equation,

$$\pi_{n,t+1} = \alpha_n + \beta_n(R_{n,t} - R_t) + \gamma_n\pi_t + resid_{t+1}, \tag{21}$$

which replaces π_{t+1} with π_t, will provide similar results. In fact, both formats were estimated during empirical preparation for this paper and similar results were obtained.

[15]This equation could be replaced with general specifications without changing the basic intuition. In fact, (22) is very similar to the slightly more general specfications in Svensson (1996) and Rudebusch and Svensson (1998), which have real rates directly affecting output, with inflation responding to excess demand. Other general specifications might have money growth negatively related to the real interest rate and inflation moving with money growth in the long run. Alternatively, money growth could be negatively related to the real interest rate, increases in money

assumed to set nominal interest rates, r_t, such that real rates exceed the long-run equilibrium real rate whenever expected inflation, $E_t\pi_{t+1}$ exceeds the inflation rate targeted by monetary policy, π_t^T:[16]

$$r_t = E_t\pi_{t+1} + \bar{rr} + \phi(E_t\pi_{t+1} - \pi_t^T) \tag{23}$$

where $\phi > 0$. This policy function is like a forward-looking Taylor (1993) rule, but without the output gap. Forward-looking rules have been examined by Clarida, Gali, and Gertler (1998), Rudebusch and Svensson (1998), and discussed by Stuart (1996) and Freedman (1996). Three descriptions of the inflation target will be compared below to illustrate that stability in the policy target for inflation leads to less persistence in the inflation process.

Case 1: Monetary policy targets a constant inflation rate. To represent the case in which the monetary authority has historically targeted and is continuing to target the same constant rate of inflation, assume $\pi_t^T = \bar{\pi}$. The rational expectations forecast of inflation can be obtained by substituting for real rates in (22) from (23), taking conditional expectations, and solving for $E_t\pi_{t+1}$. With a constant inflation target, the rational expectations forecast of inflation is:

$$E_t\pi_{t+1} = \frac{1}{1+\theta\phi}\pi_t + \frac{\theta\phi}{1+\theta\phi}\bar{\pi} \tag{24}$$

The reduced form inflation process, under rational expectations, is:

$$\begin{aligned}\pi_{t+1} &= \pi_t - \theta(r_t - E_t\pi_{t+1} - \bar{rr}) + \upsilon_{t+1} \\ &= \pi_t - \theta\phi(E_t\pi_{t+1} - \bar{\pi}) + \upsilon_{t+1} \\ &= \frac{1}{1+\theta\phi}\pi_t + \frac{\theta\phi}{1+\theta\phi}\bar{\pi} + \upsilon_{t+1}. \end{aligned} \tag{25}$$

Inflation persistence can be measured by the size of the AR(1) coefficient, which was represented by γ in the previous two subsections. For this case of constant inflation targeting, the measure of inflation persistence is $\gamma(1) = 1/(1+\theta\phi)$. For $\theta > 0$ and $\phi > 0$, $0 < \gamma(1) < 1$, and inflation follows a mean-reverting process with mean $\bar{\pi}$.

growth could lead to increases in demand, and inflation could increase with excess demand. In general, frictions such as price stickiness or adjustment costs would imply a lag between changes in real rates and changes in inflation. The implicit causal orderings in these sample generalizations are motivated by empirical regularities.

[16] As with the first equation of the model, more general forms of the policy rule could be used without changing the intuition of the model. For instance, in a model where inflation arises from excess demand and output responds to increases in real interest rates, the policy rule could be generalized so that nominal rates are also increased in response to excess demand.

Case 2: Monetary policy aims to keep the rate of inflation unchanged from its current rate. To represent the case in which the monetary authority is willing to let bygones be bygones, regards recent inflation rates as acceptable, and adjusts the inflation target accordingly, assume $\pi_t^T = \pi_t$. Under rational expectations, the resultant reduced form inflation process is:

$$\pi_{t+1} = \pi_t + \upsilon_{t+1}, \tag{26}$$

inflation persistence for this case is $\gamma(2) = 1$, and the inflation process contains a unit root. In this case, shocks υ_t have permanent effects on inflation because the monetary authority changes the target one-for-one with changes in inflation.

Case 3: Monetary policy targets an inflation rate between the current rate and a constant rate. In this case, the monetary authority partially adjusts the targeted rate of inflation to reflect recent deviations of inflation from a preferred constant rate. Assuming that the target is adjusted by the fraction $0 < \delta < 1$ of the deviation of recent rates from the preferred constant $\bar{\pi}$, the target rate is equal to $\pi_t^T = \bar{\pi} + \delta(\pi_t - \bar{\pi})$. For this description of the inflation target, under rational expectations, the reduced form inflation process is:

$$\pi_{t+1} = \frac{1 + \theta\phi\delta}{1 + \theta\phi}\pi_t + \frac{\theta\phi(1-\delta)}{1+\theta\phi}\bar{\pi} + \upsilon_{t+1} \tag{27}$$

inflation persistence is $\gamma(3) = 1/(1 + \theta\phi) + (\theta\phi\delta)/(1 + \theta\phi)$, and inflation follows a mean-reverting process with mean $\bar{\pi}$.

Comparing the three cases, it is easy to see that inflation is the least persistent when the policy target is a constant inflation rate. Because $0 < \delta < 1$ it follows that $0 < \gamma(1) < \gamma(3) < \gamma(2) = 1$. Inflation is more persistent in case 3 than in case 1, but, because the policy target is not adjusted to fully reflect recent deviations of inflation from the preferred constant target as was assumed in case 2, inflation continues to follow a mean-reverting process, with mean $\bar{\pi}$.

4 Empirical Results

This section examines whether the term structure helps predict inflation using monthly data for Canada, France, Germany, Italy, Japan, the Netherlands, Sweden, Switzerland, the UK, and the United States and quarterly data for Australia.[17] Also investigated is whether the short-term

[17] To simplify exposition, the text will refer to data for all countries as monthly even though data for Australia is quarterly. It is left to the reader to make the obvious text translations for Australia. For example, references to 12

real rate is a better predictor of future inflation than the spread. Data is described in the Data Appendix. Calculations are based on the longest sample period for which data was available for short-term interest rates, long-term interest rates, and consumer price inflation.[18] [19]

Interest rate data is similar to that used by Bonser-Neal and Morley (1997) and Kozicki (1997). The term structure spread was measured as the difference between the yield on a long-term government bond and a 3-month rate. For Australia, inflation was estimated as 100 times the change in the log consumer price index over the previous four quarters. For remaining countries, inflation was estimated as 100 times the change in the log consumer price index over the previous twelve months. Measuring inflation over a year reduces measurement difficulties that arise when inflation is calculated over shorter intervals. First, month-over-previous-month and quarter-over-previous-quarter measures of inflation tend to be quite volatile, with much of the volatility regarded as noise or temporary flucations. Second, rounding problems introduce spurious volatility in these short-period measures of inflation over much of the early sample.

Short-term real rates serve both as a measure of the stance of monetary policy and as a measure of the level of the yield curve. Short-term real rates were estimated as the difference between the short rate used in constructing the term structure spread and expected inflation. Inflation over the prior four quarters (Australia) or twelve months (all countries except Australia) is used to proxy for short-term expected inflation. In addition to the smoothing arguments listed above to justify using measures of inflation calculated over a year, variation in the smoother series is closer to that observed in survey data on inflation expectations of consumers, investors, and forecasters.

Table 1 contains summary statistics on the term structure spread, one-year inflation rates, and short-term real rates. For each series, the third column reports the sample mean, the fourth column reports the standard deviation, and the fifth through seventh columns report correlation coefficients with, respectively, the spread, one-year inflation rates, and short-term real rates in the same country. The term structure spread and inflation are generally negatively correlated. Inflation

months should be taken to mean 4 quarters for Australia, 12, 36, and 60 month forecasts should be taken to mean 4, 12, and 20 quarter forecasts, and autocorrelations at lags 1, 12, and 24 should be taken to mean autocorrelations at lags 1, 4, and 8.

[18] This implies that the longer the forecast horizon being assessed in a regression, the shorter the estimation sample.

[19] Qualitative results presented in this section do not appear to be sensitive to the sample period chosen. Analysis was also performed on a shorter sample—starting in 1975:1 for all countries except Italy—and similar results to those reported were obtained. For comparison purposes results for the most general model are reported for the full sample in Table 6 and the shortened sample in Table 7.

and short-term real rates are negatively correlated in all countries, and highly negatively correlated with correlation coefficients more negative than -0.5 in eight of the eleven countries. Short-term real rates and the term structure spread also exhibit substantial negative correlations, with correlation coefficients more negative than -0.45 in eight of the eleven countries. Because short-term real rates move closely with the real rate that reflects the stance monetary policy (i.e., the real federal funds rate in the United States) the latter result supports the hypothesis that the term structure spread also reflects the stance of monetary policy.

Table 2 reports autocorrelation coefficients for one-year inflation rates. Because data is one-year inflation rates, calculated as the annual rate of inflation over the previous twelve months, subsequent monthly observations embed a 11-month overlap.[20] This overlap occurs because 12-month inflation rates reported for subsequent months contain eleven months of common data.[21] Because of the eleven-month overlap, high estimates of autocorrelation coefficients at a 1-month lag are not unexpected. Autocorrelation coefficients at a 12-month lag, calculated using monthly data on 12-month inflation rates will be quantitatively similar to first-order autocorrelation coefficients calculated using annual data on one-year inflation rates.

Inflation rates in all countries exhibit considerable persistence as measured by autocorrelation coefficients. Results are consistent with the discussion in section 3.3. In particular, persistence is lower in countries such as Germany and Switzerland who have longstanding reputations for following monetary policies seeking low and stable inflation. Persistence is higher in countries such as Canada, France, and Italy, who, at various times in their histories, have been regarded as less vigilant in their attempts to keep inflation low. Policymakers in Canada began targeting explicit inflation rates in the 1990s and have been successful at keeping inflation under control. In Italy, success on the inflation front has been more recent. Italian inflation subsided to rates less than or equal to three percent only within the past few years as efforts to achieve European Monetary Union targets accelerated.

Data constraints make it impossible to estimate the basic inflation-change equations of Jorion

[20] For example, December 1996 12-month inflation is the average inflation rate over January 1996 through December 1996 and January 1997 12-month inflation is the average inflation rate over February 1996 through January 1997. Both December 1996 and January 1997 12-month inflation rates put a (1/12) weight on monthly inflation in each of the 11 months of February 1996 through December 1996.

[21] For Australia, inflation is calculated as the annual rate of inflation over the previous four quarters. Thus, subsequent quarters contain three quarters of common data.

and Mishkin (1991), Gerlach (1995), and Day and Lange (1997) that match the inflation horizons to the bond maturities in the term structure spread. Instead, inflation-change regressions of the form,

$$\pi_{n,t+1} - \pi_{3,t+1} = \alpha_n + \beta_n(R_{120,t} - R_{3,t}) + resid_{t+1} \tag{28}$$

were estimated, where $(R_{120,t} - R_{3,t})$ is term structure spread as described earlier, and $(\pi_{n,t+1} - \pi_{3,t+1})$ is the difference between $\pi_{n,t+1}$, the n-month ex-poste inflation rate from month t to month $t+n$ expressed at an annual rate and current inflation $\pi_{3,t+1}$. The 1-year inflation rate from month $t-9$ to month $t+3$ was used for $\pi_{3,t+1}$ instead of the 3-month inflation rate from t to $t+3$ to reduce volatility due to noise and rounding of the price index. The horizon, n, of the forecast was varied between 12, 36, and 60 months.[22] This generalization seems reasonable since term structure spreads constructed as the gap between yields on longer maturity bonds and a short-maturity instrument are highly correlated for different maturities of the longer-maturity bond. Also, use of this generalization enabled expansion of the data sample beyond the four countries considered by Jorion and Mishkin.

Table 3 contains regression results from estimation of the inflation-change equation (28).[23] Three results stand out. First, estimates of the coefficient β_n tend to increase in magnitude with n. Estimates are positive and significant based on asymptotic 5 percent critical values, at horizons of 36 and 60 months (at least) for Australia, Canada, Japan, Sweden, and the United States and at a 60-month horizon for Germany and Switzerland. Second, in general, the regression R^2s reported in the column labeled $R^2(\pi_{n,t+1} - \pi_{3,t+1})$ also increase with n.[24] At a 60-month forecast horizon, among the seven countries with significant estimates of β_{60}, R^2s vary between a low of 0.06 for Switzerland and a high of .38 for Japan. Third, estimates of the alternative R^2, reported in the column labeled $Alt.R^2(\pi_{n,t+1})$, tend to decrease with n.

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[22] Fama (1990) made a similar generalization, limiting consideration to a 5-year spread, but considering inflation forecasting horizons of 1 through 5 years. Estrella and Mishkin (1997) also relaxed the maturity-matching restrictions.

[23] Standard errors of coefficient estimates were corrected for heteroskedasticity and serial correlation following the procedure recommended by Newey and West (1987) with the Newey and West (1994) automatic lag selection routine.

[24] The table contains two columns with R^2s. The column labeled $R^2(\pi_{n,t+1} - \pi_{3,t+1})$ is the standard regression R^2 for a regression of the specification in (28), reporting the proportion of the variance in the future change in inflation explained by the spread. Entries in the column labelled *Alt.* $R^2(\pi_{n,t+1})$ contain the estimate of the proposed alternative R^2. This alternative R^2 is equal to the R^2 that would obtain from the following regression

$$\pi_{n,t+1} = \alpha_n + \beta_n(R_{120,t} - R_{3,t}) + \gamma_n \pi_{3,t+1} + resid_{t+1} \tag{29}$$

with the coefficient on $\pi_{3,t+1}$ constrained to equal unity, i.e., $\gamma_n = 1$.

Two qualitative results, that coefficient estimates of β tend to be positive and increase with forecast horizon and that R^2s tend to increase with forecast horizon, are consistent with the findings in Table 3 of Jorion and Mishkin (1991) and Table 2 of Day and Lange (1997).[25] However, these two qualitative results need not imply that the term structure spread helps predict inflation. If the term structure spread has information about *future* inflation that can be obtained from the inflation-change regressions, then the regression equation should explain variation in $\pi_{n,t+1}$ and the alternative R^2 defined in (13) of section 3.1 should be positive.

The third result, that *Alt.* $R^2(\pi_{n,t+1})$ tends to decrease with n, suggests a very different story of the information about future inflation provided by the inflation-change equations. At short horizons, the proportion of the variation in future inflation explained by the inflation-change equation is relatively large, at least 0.49 in all countries at a 12-month forecast horizon and higher than the explained variation in the inflation change variable, $\pi_{n,t+1} - \pi_{3,t+1}$. This result is intuitive. Inflation rates are quite persistent. As shown in Table 2, autocorrelations at a 12-month lag varied between a low of 0.60 for Sweden and Switzerland and a high of 0.85 for France. Thus, current inflation holds a lot of information about inflation over the next year. Current inflation should be a good predictor of future inflation, the more persistent is inflation. Autocorrelation statistics provide further evidence to support this view. *Alt.* $R^2(\pi_{n,t+1})$ are higher in those countries with more persistent inflation as indicated by higher autocorrelation coefficients in Table 2.

For long horizons, however, results are strikingly different. For four countries, Australia, Germany, Sweden, and Switzerland, *Alt.* $R^2(\pi_{n,t+1})$ is *negative* at a 60-month horizon. Returning to Table 2, these are the countries with the lowest autocorrelation coefficients at lag=24 months. Thus, these are countries with relatively less persistent inflation, for which the unit coefficient restriction on $\pi_{3,t+1}$ is most likely to be rejected by the data. The large negative values of *Alt.* R^2 can be interpreted as a rejection by the data of the implied unit coefficient on current inflation, i.e., a rejection of $\gamma_n = 1$. Furthermore, the large negative values of *Alt.* $R^2(\pi_{n,t+1})$ suggest that estimates of β_n are driven by correlations between current inflation, $\pi_{3,t+1}$, and the term structure spread, not by predictive information in the spread for future inflation, $\pi_{n,t+1}$.

These observations on *Alt.* $R^2(\pi_{n,t+1})$ suggest two reasons why $R^2(\pi_{n,t+1} - \pi_{3,t+1})$ increases

[25]Gerlach (1995) found that for Germany estimates of β_n and R^2s for inflation-change equations increased with n for n less than or equal to 6 years, but decreased with n for n greater than 6 years.

with n. First, at longer horizons, regression R^2s for inflation-change regressions may be picking up correlation between current inflation and the term structure spread, not between future inflation and the term structure spread. Second, at short horizons, inflation is relatively predictable given current inflation because inflation is persistent. Inflation changes, by contrast are less likely to be predictable at short horizons–implying low $R^2(\pi_{n,t+1} - \pi_{3,t+1})$s of estimated inflation-change equations.

The results in Table 3 suggest that better predictors of future inflation, $\pi_{n,t+1}$ can be obtained by relaxing the unit-coefficient restriction on π_t. Table 4 contains results from estimation of

$$\pi_{n,t+1} = \alpha_n + \beta_n(R_{120,t} - R_{3,t}) + \gamma_n \pi_{3,t+1} + resid_{t+1} \tag{30}$$

where γ_n is unrestricted.[26] Coefficients were estimated using instrumental variables techniques using a constant, $(R_{120,t} - R_{3,t})$, and $\pi_{3,t}$ as instruments. Three results are discussed below.

First, point estimates of γ_n are less than one for all countries and at all horizons. Improvements obtained as a results of allowing free estimation of γ_n are evident by comparing regression R^2s in the final column of Table 4 to entries in the the column labelled *Alt.* $R^2(\pi_{n,t+1})$ in Table 3. At short horizons, the gains in terms of variation in $\pi_{n,t+1}$ explained by regressors are relatively small. Only small gains are realized because, even when freely estimated, estimates of γ_n remain close to one. At long horizons, however, the gains are substantial as estimates of γ_{60} are significantly below one for all countries. Gains are particularly sizable in the four countries that had negative estimates of *Alt.* $R^2(\pi_{n,t+1})$ in Table 3. R^2s rise to positive values of 0.32, 0.08, 0.50, and 0.02 respectively, for Australia, Germany, Sweden, and Switzerland. Estimates of γ_n are smaller than or equal to 0.37 in these five countries, and insignificantly different from zero in two of the countries. These results support the view that large negative values of *Alt.* R^2 obtained during estimation of the inflation-change equations reflect a rejection by the data of the implied unit coefficient restriction on γ_n.

A second result evident in Table 4 is that regression R^2s for the unrestricted inflation equation (29) tend to decrease with forecasting horizon. This is similar to the finding in Table 3 from

[26] Fama (1990) estimated an inflation-change equation including current inflation as an explanatory variable. This approach controlled for potential correlation between the current inflation component of the regressand and the term structure spread. Fama obtained negative estimates of $(\gamma_n - 1)$, which implied estimates of γ_n similar to those reported in Table 4. However, because the regressand is the inflation change variable, these equations are still not the best approach for assessing the predictability of future inflation.

estimation of the inflation-change equation (28), that *Alt.* $R^2(\pi_{n,t+1})$ tended to decrease with n, but opposite to the finding that $R^2(\pi_{n,t+1} - \pi_{3,t+1})$ generally increased with n. However, the result that inflation becomes less predictable as the forecast horizon increases seems more reasonable than the reverse.

A third result revealed in a comparison of Table 4 to Table 3 is that estimates of β_n often decrease when the unit coefficient restrictions on γ_n are relaxed. For Germany, Japan, Switzerland, and the United States, findings of statistical significance were reversed. In particular, as shown in Table 3, positive statistically significant estimates of β_n were obtained for longer horizons when γ_n was constrained to equal unity. But, as shown in Table 4, when γ_n was unconstrained and freely estimated, estimates of β_n declined in considerably in magnitude and ceased to be significant. These results suggest that significance of β_{60} in the unrestricted regression may have been due to correlation between the term structure spread and current inflation. In other words, *the term structure spread may not help predict future 60-quarter inflation even when estimates of β_{60} in Table 3 are positive and significant.*

Regression equations estimated to this point are grounded in the first interpretation of the Expectations Hypothesis offered in section 2. This interpretation related the term structure spread to the expected real rate change, the expected inflation change, and the term premium, as represented in (4). However, a second hypothesis was offered in section 2–the term structure spread may help predict future inflation because it reflects the stance of monetary policy.

If the predictive power of the term structure spread derives solely from its role as an indicator of the stance of monetary policy, then the significance of the term structure spread may disappear when other measures of the stance of monetary policy are also included as regressors. Tables 5, 6, and 7 contain regression results paralleling those in Tables 3 and 4, but with the short-term real rate also included as an explanatory variable. Regression results from estimation of

$$\pi_{n,t+1} - \pi_{3,t+1} = \alpha_n + \beta_n(R_{120,t} - R_t) + \delta_n(R_{3,t} - \pi_{3,t}) + resid_{t+1} \quad (31)$$

are summarized in Table 5.

Comparison of results in Table 5 with those in Table 3, shows that when the short-term real rate is included as a regressor, estimated coefficients on the term structure spread increase relative to Table 3 estimates, becoming statistically significant at 5% in the Netherlands and the UK. In five countries, coefficient estimates on the short-term real rate were *positive* and statistically significant.

Given the correlations in Table 2, positive estimates of the coefficient on the real rate would be expected if the coefficient was reflecting correlation between the real rate and the current inflation component of the dependent variable. Negative coefficient estimates would have been expected if the coefficient was reflecting a relationship between monetary policy actions and subsequent inflation. Despite statistical significance of estimated coefficients on the spread and the real rate, once again the explanatory variables do little to help predict $\pi_{n,t+1}$ at longer horizons in several countries. At a 60-month forecast horizon, *Alt.* $R^2(\pi_{n,t+1})$ remains negative in Australia, Germany, Sweden, and Switzerland despite statistical significance of the spread and (except Australia) the real rate. These results suggest that for at least these four countries, at longer horizons, statistical significance of estimated coefficients likely reflects correlation between current inflation and the regressors, not between future inflation and the regressors. More generally, within the structure of the inflation-change equations, statistical significance of coefficient estimates need not imply that regressors help predict $\pi_{n,t+1}$.

Results from estimation of

$$\pi_{n,t+1} = \alpha_n + \beta_n(R_{120,t} - R_{3,t}) + \delta_n(R_{3,t} - \pi_{3,t}) + \gamma_n \pi_{3,t+1} + resid_{t+1} \tag{32}$$

with γ_n unrestricted, are summarized in Table 6 and Table 7.[27] Table 6 results, following the practice in the previous tables, are based on analysis using the longest sample available for each country. Table 7 results are for samples constrained to include only observations dated 1975:1 or later.[28] At longer horizons, coefficient estimates of γ_n are generally insignificant. This result reinforces the earlier claim that the implicit unit coefficient on $\pi_{3,t+1}$ in the inflation-change equations of Table 5 would be rejected by the data, and, consequently that estimates of β_n and δ_n in Table 5 are to a certain extent reflecting correlations between current inflation and the short-term real rate and the term structure spread.

Coefficient estimates on the short-term real rate are either negative and statistically significant or insignificantly different from zero in all cases. Since higher short-term real rates signal tighter

[27] Since both the real rate, constructed as the nominal short rate less inflation from $t-12$ to t, and inflation from $t-9$ to $t+3$ are both included as regressors, and the two inflation measures are highly correlated, it would be inappropriate to conclude that it is the short-term real rate and not the short-term nominal rate that matters for predicting inflation. If instead the regression equation $\pi_{n,t+1} = \alpha_n + \beta_n(R_{120,t} - R_{3,t}) + \delta_n R_{3,t} + \gamma_n^* \pi_{3,t+1} + resid_{t+1}$ was estimated, nearly identical coefficient estimates of δ_n and β_n would be obtained, and the estimate of γ_n^* would almost equal $\gamma_n - 1$ with estimates of γ_n as in Table 6.

[28] Since data for Italy starts in March 1976, results for Italy are identical in both tables.

monetary policy, a negative estimate of δ_n is consistent with the view that future inflation will fall in response to relatively tight monetary policy. Coefficient estimates are more frequently statistically significant at longer horizons than at shorter horizons. This result is consistent with the view that inflation responds to monetary policy with long and variable lags.

If the predictive power of the term structure spread is due to its role as a measure of the stance of monetary policy, then estimates of β_n may become statistically insignificant when short-term real rates are included as regressors. Empirical results generally support this view. In Table 4, coefficient estimates of β_n were positive and significant at the 60-month forecasting horizon for each of Australia, Canada, France, Sweden, and the UK. By contrast, in Table 6, when short-term real rates are also included as regressors, coefficient estimates of β_n are positive and significant only for Sweden and the United States.

One explanation for the dominating predictive power of the real short rate for inflation is that this variable provides a cleaner measure of the stance of monetary policy than the term structure spread. If inflation is primarily a monetary phenomenon in the long run, then better measures of the stance of monetary policy might be expected to be better predictors of future inflation at moderate or long horizons. The short rate moves very closely with monetary policy actions. By contrast, although the term structure spread may contain information on the stance of monetary policy, the spread also likely contains information on credit market conditions.

5 Conclusions

The discussion and results in this paper warn against traditional interpretations of estimation results from inflation-change equations when assessing whether the term structure spread helps predict future inflation. In regressions of the difference between ex poste future inflation and current inflation on a constant and the term structure spread, positive regression R^2s and statistically significant estimates of the coefficient on the spread do not always signify that the spread helps predict future inflation. In countries such as Australia, Germany, Sweden, and Switzerland, where the inflation process is less persistent, regression R^2s and estimates of the coefficient on the term structure spread likely reflect correlations between the spread and *current* inflation and should not be interpreted as evidence on the predictive power of the spread for *future* inflation. Results from an empirical analysis of data from eleven industrialized countries suggest that the level of the short-term real rate may be more useful for predicting inflation than the term structure spread, possibly because changes in short-term real rates provide cleaner measures of changes in the stance of monetary policy.

6 Data Appendix

Australia: Quarterly data, 1969:Q2 - 1997:Q4

Short Rate – Weighted average yield on 13-week treasury notes allotted at last tender of month. Missing observation in October 1997 replaced with average of observations for September 1997 and November 1997. Source: International Financial Statistics, International Monetary Fund (IFS).

Long Rate – Assessed secondary market yields on non-rebate bonds with maturity of at least 10 years. Yields are calculated before brokerage and on the last business day of the month. Source: IFS.

Consumer Price Index – Consumer Prices, all groups, referring to a weighted average of eight capital cities, base 1989-1990. Source: IFS.

Canada: Monthly data, January 1958 - December 1997

Short Rate – Weighted average of the yields on successful bids for 3-month bills. Monthly data refer to the tender rates of the last Wednesday of the month. Source: IFS.

Long Rate – Secondary market, average bond yields on Government of Canada Bonds over 10 years. Last Wednesday of the month. Source: Bank for International Settlements (BIS).

Consumer Price Index – Consumer Prices, data for all cities with a population of over 30,000, with weights corresponding to family expenditure patterns of 1982, base 1986. Source: IFS.

France: Monthly data, January 1970 - December 1997

Short Rate – 3-month paris Interbank Offer Rate, average of all days in month (Friday values repeated on weekends), rates calculated on 360 days. Break in January 1987: Previously rates practiced on money market (pensions between banks). Source: BIS.

Long Rate – Secondary market yield, public and semi-public bonds. Loans subject to witholing tax. Incl. tax credit allocated to recipients of interest. Calculation takes purchase commissions into account. Missing observation in April 1974 replaced with average of observations for March 1974 and May 1974. Missing observation in March 1979 replaced with average of observations for February 1979 and April 1979. Source: BIS.

Consumer Price Index – Base 1990. The index is a Laspeyres chain index which reflects the entire population. The weights are revised each year on the basis of continuing family expenditure surveys and national accounts consumption expenditure data. Source: IFS.

Germany: Monthly data, January 1961 - December 1997

Short Rate – Frankfurt 3-month Interbank loan rate. Source: Board of Governors of the Federal Reserve System (BOG).

Long Rate – From January 1961 through December 1982 and for June 1983 data is for 7-15 year public sector bonds. Source: OECD Main Economic Indicators. From January 1983 through December 1997, excluding June 1983, data is the yield on German Goverment Bellwether bond. Source: BOG.

Consumer Price Index – Laspeyres index, weighting base 1991. The weights are updated every five years. Data cover the former Federal Republic of Germany prior to 1991. Data cover the former Federal Republic of Germany and the former German Democratic Republic from 1991 onward. Source: IFS.

Italy: Monthly data, March 1976 - December 1997

Short Rate – Average of the allotment rates at public auction of ordinary Treasury Bills (compound yield) gross of tax. Calculation based on calendar year (365 days). Calc. by taking weighted average of yields at auctions during month. Source: BIS. Missing observations February 1979 through October 1979 are 3-month Interbank rate. Data represent the arithmetic averages of daily rates quoted at noon. Source: IFS.

Long Rate – Monthly data are arithmetic averages of daily gross yields to maturity of fixed-copupon Treasury bonds with residual maturities between 9 and 10 years, based on prices in the official wholesale market. Prior to 1991, the data are average yields to maturity on bonds with original maturities of 15 to 20 years, issued on behalf of the Treasury by the Consortium of Credit for Public Works. Source: IFS.

Consumer Price Index – Index base 1985. The series is the national index, which represents the weighted arithmetic average of 20 areas, weighted by their population on December 21, 1984. The commodity weights reflect 1984 consumption patterns. Source: IFS.

Japan: Monthly data, January 1969 - December 1997

Short Rate – Yields of bonds trading with repurchase agreement (3-month). Rates are those offered by clients to securities companies. Source: BIS.

Long Rate – yield at issue, to subscribers on 10-year interest bearing Government bonds (prior to January 1973, 7-year). Source: BIS.

Consumer Price Index – Base 1985, The index is compiled by the Statistics Bureau and covers the whole country excluding one-person households and those engaged mainly in agriculture, forestry, and fishing. The weights were derived from the Family Income and Expenditure Survey conducted in 1985. Source: IFS.

Netherlands: Monthly data, October 1972 - December 1997

Short Rate – 3-month Amsterdam Interbank Offer Rates (montly average) based on offer rates of seven banks. Prior to December 1985, Interbank Deposit Rate. Source: BIS.

Long Rate – Yield on most recent 10-year government bond. Source: IFS.

Consumer Price Index – The data, base 1985, refer to wage earners' families of husband and wife without children or with nonincome-earning children living in the household. In 1985 the gross income of these households was equal to or above the wage level at which health insurance is mandatory. Source: IFS.

Sweden: Monthly data, December 1962 - December 1997

Short Rate – Rate on 3-month Treasury discount notes. Source: IFS.

Long Rate – Starting in December 1986: data is secondary market yield on 9 or 10 year government bonds. Monthly averages. From January 1994, 9-year government bonds. Source: BIS. Data prior to December 1986: Until December 1979 data refer to yields on government bonds maturing in 15 years, For January 1980 through November 1986 data refer to yields on bonds maturing in 10 years. Source: IFS.

Consumer Price Index – Base 1980. The weights are derived from national account estimates of private consumption expenditure and are revised every December. The index covers 70 urban and rural areas. Source: IFS.

Switzerland: Monthly data, January 1958 - December 1997

Short Rate – Time deposits, 3-month, with large banks. Data prior to June 1989 applied by applied by agreement by 4 main banks (fixed deposit convention). Source: BIS.

Long Rate – Secondary market yield on Confederation bonds. Until December 1981: all loans with remaining maturity of between 5 and 12 years, calcuated on basis of final maturity. From January 1982 onwards, all loans with at least 5 years to maturity and at least 3 years to first call date, calculated on the basis of final maturity if market price below call price or to call date if above call price. Source: BIS.

Consumer Price Index – Base December 1982. Beginning January 1992 however, base is May 1993. The weights are based on a family expenditure survey in 1975 for the whole country, from a sample of 980 households. Source: IFS.

United Kingdom: Monthly data, January 1962 - December 1997

Short Rate – Daily 3-month interbank sterling figs. Source: BOG.

Long Rate – Theoretical gross redemption bond yields. Beginning June 1976, the calculations are based on a method described in the Bank of England, *Monetary and Financial Statistics*, June 1976. Issue at par with 20 years to maturity. Source: IFS.

Consumer Price Index – Data refer to general index of retail prices, all items, base January 1987.

The weights are revised at the beginning of each year on the basis of results of a continuing family expenditure survey. Source: IFS.

USA: Monthly data, January 1958 - December 1997

Short Rate – 3-month Treasury Bill Secondary Market Rate. Raw data is averages of the bid rates quoted on a bank discount basis by a sample of primary dealers who report to the Federal Reserve Bank of New York. The rates are reported based on quotes at the official close of the U.S. government securities market for each business day. Data converted to yields as follows: $\text{yield} = 100 \times (((1 - \text{rate}/400)^{-4.0555555556}) - 1)$. Source: BOG.

Long Rate – Market yield on U.S. Treasury securities at 10-year constant maturity, quoted on investment basis. Source: BOG.

Consumer Price Index – Consumer Price Index (All Urban), All Items, 1982-1984=100. Seasonally Adjusted. Source: Bureau of Labor Statistics.

7 References

Abken, Peter A., 1993, Inflation and the Yield Curve, Federal Reserve Bank of Atlanta Economic Review, 78, 13-31.

Alles, Lakshman A. and Ramaprasad Bhar, 1997, The information on inflation in the Australian term structure, Applied Financial Economics, 7, 721-730.

Bernanke, Ben S., and Alan S. Blinder, 1992, The Federal Funds Rate and the Channels of Monetary Transmission, American Economic Review, 82, 901-921.

Blough, Stephen R., 1994, Yield Curve Forecasts of Inflation: A Cautionary Tale, New England Economic Review, 3-15.

Bonser-Neal, Catherine, and Timothy R. Morley, 1997, Does the Yield Spread Predict Real Economic Activity? A Multicountry Analysis, Federal Reserve Bank of Kansas City Economic Review, 82, 37-53.

Campbell, John Y., and Robert J. Shiller, 1991, Yield Spreads and Interest Rate Movements: A Bird's Eye View, Review of Economic Studies, 58, 495-514.

Clarida, Richard, Jordi Gali, and Mark Gertler, 1998, Monetary Policy Rules in Practice: Some International Evidence, European Economic Review, 42, 1033-1067.

Davis, E. Philip and Gabriel Fagan, 1997, Are Financial Spreads Useful Indicators of Future Inflation and Output Growth in EU Countries, Journal of Applied Econometrics, 12, 701-714.

Day, Jim, and Ron Lange, 1997, The Structure of Interest Rates in Canada: Information Content about Medium-Term Inflation, Bank of Canada Working Paper 97-10, May 1997.

Engsted, Tom, 1995, Does the Long-Term Interest Rate Predict Future Inflation? A Multi-Country Analysis, The Review of Economics and Statistics, 77, 42-54.

Estrella, A. and Gikas A. Hardouvelis, 1991, The Term Structure as a Predictor of Real Economic Activity, The Journal of Finance, 46, 555-575.

Estrella, A. and Frederic S. Mishkin, 1997, The predictive power of the term structure of interest rates in Europe and the United States: Implications for the European Central Bank, European Economic Review, 41, 1375-1401.

Estrella, A. and Frederic S. Mishkin, 1996, Predicting U.S. Recessions: Financial Variables as Leading Indicators, Federal Reserve Bank of New York, research paper no. 9609.

Fama, Eugene F., 1990, Term-Structure Forecasts of Interest Rates, Inflation, and Real Returns, Journal of Monetary Economics, 25, 59-76.

Frankel, Jeffrey A., and Cara S. Lown, 1994, An indicator of future inflation extracted from the steepness of the interest rate yield curve along its entire length, The Quarterly Journal of Economics, 109, 517-530.

Freedman, Charles, 1996, What Operating Procedures Should be Adopted to Maintain Price Stability—Practical Issues, in Achieving Price Stability: A Symposium Sponsored by the Federal Reserve Bank of Kansas City. Kansas City, 209-227.

Gerlach, Stefan, 1995, The Information Content of the Term Structure: Evidence for Germany, BIS Working paper No. 29, September 1995.

Harvey, Campbell R., 1989, Forecasts of Economic Growth from the Bond and Stock Markets, Financial Analysts Journal, September/October, 38-45.

Jorion, Philippe, and Frederic Mishkin, 1991, A Multicountry Comparison of Term-Structure Forecasts at Long Horizons, Journal of Financial Economics, 29, 59-80.

Kozicki, Sharon, 1997, Predicting Real Growth and Inflation with the Yield Spread, Federal Reserve Bank of Kansas City Economic Review, 82, 39-57.

Mishkin, Frederic, 1991, A Multi-Country Study of the Information in the Shorter Maturity Term Structure About Future Inflation, Journal of International Money and Finance, 10, 2-22.

Mishkin, Frederic S., 1990, What Does the Term Structure Tell Us About Future Inflation, Journal of Monetary Economics, 25, 77-95.

Newey, Whitney K., and Kenneth D. West, 1994, Automatic Lag Selection in Covariance Matrix Estimation, Review of Economic Studies, 61, 631-653.

Newey, Whitney K., and Kenneth D. West, 1987, A Simple, Positive Semi-Definite, Heteroskedasticity and Autocorrelation Consistent Covariance Matrix, Econometrica, 55, 703-708.

Plosser, Charles I., and K. Geert Rouwenhorst, 1994, International Term Structures and Real Economic Growth, Journal of Monetary Economics, 33, 133-155.

Rudebusch and Svensson, 1998, Policy Rules for Inflation Targeting, NBER Working Paper No. 6512.

Stuart, Alison, 1996, Simple Monetary Rules, Bank of England Quarterly Review, 36, 281-287.

Svensson, Lars E.O., 1996, Commentary: How Should Monetary Policy Respond to Shocks while Maintaining Long-Run Price Stability?—Conceptual Issues, in Achieving Price Stability: A Symposium Sponsored by the Federal Reserve Bank of Kansas City. Kansas City, 209-227.

Taylor, John B., 1993, Discretion versus policy rules in practice, Carnegie-Rochester Conference Series on Public Policy, 39, 195-214.

Tzavalis, Elias and M.R. Wickens, 1996, Forecasting inflation from the term structure, Journal of Empirical Finance, 3, 103-122.

TABLE 1

SUMMARY STATISTICS

Country	Series	Mean	Standard Deviation	Correlation with spread	inflation	real rate
Australia	spread	1.00	1.73	1.00		
	inflation	6.93	3.90	-.20	1.00	
	real rate	2.43	4.36	-.46	-.58	1.00
Canada	spread	1.22	1.59	1.00		
	inflation	4.52	3.13	-.45	1.00	
	real rate	2.66	2.70	-.45	-.24	1.00
France	spread	1.09	1.49	1.00		
	inflation	6.18	3.84	.05	1.00	
	real rate	2.91	2.96	-.52	-.64	1.00
Germany	spread	1.32	1.74	1.00		
	inflation	3.28	1.77	-.48	1.00	
	real rate	2.80	1.92	-.79	-.01	1.00
Italy	spread	-0.03	1.67	1.00		
	inflation	9.11	5.35	-0	1.00	
	real rate	4.03	3.32	-.26	-.76	1.00
Japan	spread	0.42	1.95	1.00		
	inflation	4.39	4.52	-.78	1.00	
	real rate	1.50	2.69	.29	-.73	1.00
Netherlands	spread	1.15	1.86	1.00		
	inflation	3.94	2.85	.03	1.00	
	real rate	2.95	3.20	-.74	-.62	1.00
Sweden	spread	0.96	1.91	1.00		
	inflation	6.15	3.30	-.07	1.00	
	real rate	2.18	3.64	-.58	-.53	1.00
Switzerland	spread	0.90	1.45	1.00		
	inflation	3.36	2.29	-.24	1.00	
	real rate	0.29	2.16	-.57	-.58	1.00
United Kingdom	spread	0.78	2.15	1.00		
	inflation	6.96	4.88	.08	1.00	
	real rate	2.12	4.06	-.65	-.70	1.00
United States	spread	1.03	1.37	1.00		
	inflation	4.35	2.90	-.45	1.00	
	real rate	1.89	2.20	-.13	-.30	1.00

TABLE 2

INFLATION AUTOCORRELATIONS

Country	Autocorrelation Lag (months)		
	1	12	24
Australia	.94	.69	.44
Canada	.99	.82	.65
France	.99	.85	.71
Germany	.98	.70	.40
Italy	.99	.79	.62
Japan	.99	.69	.45
Netherlands	.99	.83	.61
Sweden	.97	.60	.38
Switzerland	.98	.60	.24
United Kingdom	.99	.71	.50
United States	.99	.78	.49

For Australia autocorrelation coefficients are calculated using quarterly data on 4-quarter inflation rates, and the second, third, and fourth columns report autocorrelations at (quarterly) lags of 4, 8, and 16. Autocorrelation coefficients for remaining countries are calculated using monthly data on 12-month inflation rates.

TABLE 3
ESTIMATION RESULTS – INFLATION-CHANGE EQUATIONS

$$\pi_{n,t+1} - \pi_{3,t+1} = \alpha_n + \beta_n[R_{120,t} - R_{3,t}] + resid_{t+1}$$

Country	Horizon (n)	β_n	SE (β_n)	*Alt.* $R^2(\pi_{n,t+1})$	$R^2(\pi_{n,t+1} - \pi_{3,t+1})$
Australia	4	0.18	0.15	0.68	0.01
	12	0.55*	0.21	0.33	0.11
	20	0.70*	0.24	-0.11	0.13
Canada	12	0.12	0.11	0.78	0.01
	36	0.58*	0.20	0.61	0.20
	60	0.84*	0.18	0.45	0.30
France	12	0.11	0.13	0.84	0.01
	36	0.38	0.20	0.65	0.06
	60	0.41	0.24	0.41	0.04
Germany	12	0.04	0.08	0.61	0.00
	36	0.22	0.14	0.02	0.06
	60	0.43*	0.16	-0.76	0.16
Italy	12	0.29	0.22	0.86	0.06
	36	0.41	0.32	0.69	0.06
	60	-0.01	0.41	0.47	-0.00
Japan	12	0.70*	0.17	0.71	0.23
	36	1.14*	0.19	0.52	0.38
	60	1.27*	0.27	0.31	0.38
Netherlands	12	-0.15*	0.06	0.83	0.05
	36	-0.18	0.12	0.52	0.03
	60	-0.08	0.13	0.08	0.00
Sweden	12	0.26	0.15	0.55	0.05
	36	0.52*	0.16	0.21	0.15
	60	0.57*	0.15	-0.17	0.16
Switzerland	12	0.01	0.13	0.49	-0.00
	36	0.38	0.20	-0.44	0.06
	60	0.45*	0.22	-2.36	0.06
United Kingdom	12	0.07	0.19	0.61	-0.00
	36	0.38	0.31	0.25	0.05
	60	0.57	0.33	0.01	0.09
United States	12	0.08	0.14	0.73	0.00
	36	0.67*	0.20	0.38	0.17
	60	0.87*	0.24	0.20	0.25

* Significant at 5% level

SE($\cdot$) denotes the standard error, corrected for heteroskedasticity and serial correlation, of the estimated coefficient.

$R^2(\pi_{n,t+1} - \pi_{3.t+1})$ is the R^2 for the inflation change equation.

Alt. $R^2(\pi_{n,t+1})$ is the R^2 for the equation $\pi_{n,t+1} = \alpha_n + \beta_n[R_{120,t} - R_{3,t}] + \pi_{3,t+1} + resid_{t+1}$

Results are for monthly data, except Australian results are for quarterly data.

TABLE 4
ESTIMATION RESULTS – INFLATION EQUATIONS

$$\pi_{n,t+1} = \alpha_n + \beta_n[R_{120,t} - R_{3,t}] + \gamma_n \pi_{3,t+1} + resid_{t+1}$$

Country	Horizon (n)	β_n	SE (β_n)	γ_n	SE (γ_n)	R^2
Australia	4	0.09	0.17	0.81*	0.09	0.69
	12	0.38	0.24	0.60*	0.12	0.47
	20	0.58*	0.24	0.37*	0.17	0.32
Canada	12	0.00	0.14	0.87*	0.08	0.79
	36	0.39	0.20	0.78*	0.11	0.64
	60	0.60*	0.17	0.68*	0.10	0.56
France	12	0.13	0.13	0.91*	0.06	0.84
	36	0.44*	0.18	0.76*	0.07	0.69
	60	0.45*	0.23	0.64*	0.10	0.51
Germany	12	-0.11	0.08	0.71*	0.09	0.64
	36	-0.07	0.11	0.39*	0.13	0.30
	60	0.04	0.12	0.18	0.12	0.08
Italy	12	0.29	0.22	0.88*	0.06	0.86
	36	0.47	0.31	0.75*	0.06	0.75
	60	0.21	0.45	0.65*	0.09	0.62
Japan	12	0.59	0.38	0.94*	0.21	0.70
	36	0.90	0.49	0.87*	0.21	0.52
	60	0.68	0.40	0.67*	0.16	0.38
Netherlands	12	-0.15*	0.07	0.87*	0.07	0.84
	36	-0.14	0.11	0.67*	0.09	0.66
	60	-0.03	0.10	0.51*	0.07	0.57
Sweden	12	0.24	0.14	0.70*	0.10	0.59
	36	0.53*	0.13	0.50*	0.10	0.51
	60	0.61*	0.10	0.35*	0.10	0.50
Switzerland	12	-0.13	0.18	0.65*	0.14	0.54
	36	0.14	0.24	0.27	0.16	0.18
	60	0.15	0.14	-0.01	0.10	0.02
United Kingdom	12	0.09	0.15	0.75*	0.09	0.65
	36	0.46*	0.18	0.53*	0.09	0.49
	60	0.73*	0.17	0.39*	0.07	0.51
United States	12	-0.16	0.13	0.77*	0.09	0.74
	36	0.26	0.20	0.62*	0.12	0.48
	60	0.39	0.24	0.54*	0.13	0.43

* Significant at 5% level

SE($\cdot$) denotes the standard error, corrected for heteroskedasticity and serial correlation, of the estimated coefficient.

Estimated using instrumental variables with a constant, $R_{120,t} - R_{3,t}$, and $\pi_{3,t}$ as instruments.

Results are for monthly data, except Australian results are for quarterly data.

TABLE 5
ESTIMATION RESULTS – EXPANDED INFLATION-CHANGE EQUATIONS

$$\pi_{n,t+1} - \pi_{3,t+1} = \alpha_n + \beta_n[R_{120,t} - R_{3,t}] + \delta_n[R_{3,t} - \pi_{3,t}] + resid_{t+1}$$

Country	Horizon (n)	β_n	SE (β_n)	δ_n	SE (δ_n)	*Alt.* $R^2(\pi_{n,t+1})$	$R^2(\pi_{n,t+1} - \pi_{3,t})$
Australia	4	0.20	0.21	0.06	0.12	0.50	-0.00
	12	0.60*	0.19	0.08	0.16	0.16	0.07
	20	0.81*	0.24	0.14	0.16	-0.29	0.10
Canada	12	0.15	0.11	0.04	0.08	0.78	0.02
	36	0.58*	0.21	0.00	0.11	0.60	0.20
	60	0.86*	0.22	0.02	0.11	0.45	0.29
France	12	0.08	0.22	-0.04	0.12	0.84	0.01
	36	0.27	0.29	-0.10	0.15	0.66	0.07
	60	0.26	0.33	-0.14	0.14	0.42	0.06
Germany	12	0.28*	0.10	0.28*	0.11	0.64	0.09
	36	0.60*	0.17	0.44*	0.17	0.14	0.17
	60	1.03*	0.19	0.65*	0.18	-0.41	0.33
Italy	12	0.33	0.21	0.09	0.12	0.86	0.07
	36	0.47	0.29	0.09	0.12	0.69	0.07
	60	0.07	0.40	0.10	0.16	0.48	0.00
Japan	12	0.67*	0.16	0.06	0.21	0.71	0.23
	36	1.10*	0.26	0.09	0.21	0.52	0.38
	60	1.11*	0.27	0.26	0.18	0.34	0.41
Netherlands	12	-0.02	0.14	0.10	0.09	0.84	0.08
	36	0.16	0.19	0.25	0.14	0.57	0.14
	60	0.45*	0.21	0.39*	0.13	0.27	0.21
Sweden	12	0.49*	0.19	0.20*	0.10	0.59	0.11
	36	0.83*	0.22	0.25*	0.11	0.27	0.22
	60	0.93*	0.21	0.27*	0.11	-0.07	0.23
Switzerland	12	0.32*	0.12	0.36*	0.13	0.56	0.14
	36	1.04*	0.21	0.70*	0.18	0.03	0.37
	60	1.50*	0.20	1.07*	0.15	-0.49	0.59
United Kingdom	12	0.60*	0.19	0.43*	0.10	0.69	0.19
	36	1.21*	0.24	0.65*	0.14	0.48	0.33
	60	1.59*	0.22	0.76*	0.13	0.36	0.41
United States	12	0.08	0.15	0.02	0.12	0.73	0.00
	36	0.67*	0.21	0.00	0.17	0.38	0.17
	60	0.87*	0.25	-0.03	0.15	0.20	0.25

* Significant at 5% level

SE($\cdot$) denotes the standard error, corrected for heteroskedasticity and serial correlation, of the estimated coefficient.

$R^2(\pi_{n,t+1} - \pi_{3.t+1})$ is the R^2 for the inflation change equation.

Alt. $R^2(\pi_{n,t+1})$ is the R^2 for the equation $\pi_{n,t+1} = \alpha_n + \beta_n[R_{120,t} - R_{3,t}] + \pi_{3,t+1} + resid_{t+1}$

Results are for monthly data, except Australian results are for quarterly data.

TABLE 6
ESTIMATION RESULTS – EXPANDED INFLATION EQUATIONS

$$\pi_{n,t+1} = \alpha_n + \beta_n[R_{120,t} - R_{3,t}] + \delta_n[R_{3,t} - \pi_{3,t}] + \gamma_n \pi_{3,t+1} + resid_{t+1}$$

Country	Horizon (n)	β_n	SE (β_n)	δ_n	SE (δ_n)	γ_n	SE (γ_n)	R^2
Australia	4	-0.24	0.29	-0.21	0.18	0.63*	0.18	0.68
	12	-0.45	0.34	-0.53*	0.21	0.13	0.21	0.57
	20	-0.50	0.31	-0.68*	0.18	-0.24	0.21	0.52
Canada	12	-0.08	0.15	-0.07	0.09	0.83*	0.09	0.79
	36	0.14	0.23	-0.20	0.11	0.68*	0.12	0.66
	60	0.26	0.23	-0.26*	0.12	0.56*	0.11	0.59
France	12	-0.18	0.25	-0.32*	0.15	0.75*	0.08	0.85
	36	-0.34	0.23	-0.86*	0.18	0.29*	0.10	0.82
	60	-0.73*	0.22	-1.23*	0.15	-0.07	0.08	0.83
Germany	12	-0.02	0.16	0.08	0.15	0.75*	0.10	0.65
	36	-0.23	0.30	-0.13	0.24	0.30	0.17	0.28
	60	-0.24	0.28	-0.24	0.23	0.03	0.18	0.07
Italy	12	0.16	0.20	-0.23	0.19	0.76*	0.09	0.86
	36	0.09	0.19	-0.74*	0.21	0.35	0.12	0.82
	60	-0.27	0.28	-1.02*	0.21	0.08	0.13	0.79
Japan	12	0.63	0.41	0.04	0.22	0.97*	0.22	0.71
	36	0.67	0.59	-0.24	0.35	0.66	0.37	0.52
	60	0.07	0.48	-0.66	0.37	0.09	0.31	0.34
Netherlands	12	-0.40	0.26	-0.20	0.19	0.74*	0.14	0.84
	36	-0.84*	0.27	-0.55*	0.21	0.29*	0.14	0.69
	60	-1.20*	0.21	-0.89*	0.16	-0.12	0.12	0.70
Sweden	12	0.24	0.21	0.00	0.12	0.71*	0.11	0.59
	36	0.44	0.24	-0.08	0.13	0.45*	0.11	0.51
	60	0.42*	0.19	-0.15	0.11	0.26*	0.13	0.51
Switzerland	12	-0.11	0.33	0.01	0.23	0.67*	0.24	0.55
	36	-0.15	0.47	-0.24	0.32	0.08	0.32	0.13
	60	0.18	0.38	0.03	0.31	0.01	0.26	0.02
United Kingdom	12	0.60	0.33	0.44	0.26	1.00*	0.18	0.69
	36	0.72*	0.34	0.23	0.26	0.66*	0.17	0.51
	60	0.77*	0.27	0.03	0.24	0.41*	0.17	0.51
United States	12	-0.21	0.12	-0.10	0.11	0.74*	0.10	0.74
	36	0.14	0.17	-0.20	0.11	0.54*	0.13	0.50
	60	0.23	0.20	-0.28*	0.10	0.44*	0.13	0.48

* Significant at 5% level

SE(·) denotes the standard error, corrected for heteroskedasticity and serial correlation, of the estimated coefficient.

Estimated using instrumental variables with a constant, $R_{120,t} - R_{3,t}$, and $\pi_{3,t}$ as instruments.

Results are for monthly data, except Australian results are for quarterly data.

TABLE 7
ESTIMATION RESULTS – EXPANDED INFLATION EQUATIONS

(data from 1975:1 or later)

$$\pi_{n,t+1} = \alpha_n + \beta_n[R_{120,t} - R_{3,t}] + \delta_n[R_{3,t} - \pi_{3,t}] + \gamma_n \pi_{3,t+1} + resid_{t+1}$$

Country	Horizon (n)	β_n	SE (β_n)	δ_n	SE (δ_n)	γ_n	SE (γ_n)	R^2
Australia	4	0.13	0.35	0.07	0.19	0.87*	0.20	0.72
	12	0.16	0.37	-0.09	0.21	0.55*	0.19	0.59
	20	0.25	0.34	-0.13	0.20	0.43	0.23	0.58
Canada	12	-0.42	0.22	-0.12	0.13	0.67*	0.15	0.78
	36	-0.63*	0.31	-0.54*	0.19	0.30*	0.18	0.66
	60	-0.47	0.28	-0.66*	0.18	0.15	0.16	0.62
France	12	-0.13	0.17	-0.26	0.14	0.81*	0.08	0.91
	36	-0.41	0.21	-0.72*	0.19	0.44*	0.08	0.88
	60	-0.69*	0.22	-1.10*	0.17	0.08	0.07	0.87
Germany	12	0.09	0.16	0.18	0.17	0.75*	0.11	0.60
	36	-0.21	0.34	-0.17	0.30	0.19	0.19	0.18
	60	-0.53	0.29	-0.61*	0.24	-0.26	0.17	0.10
Italy	12	0.16	0.20	-0.23	0.19	0.76*	0.09	0.86
	36	0.09	0.19	-0.74*	0.21	0.35*	0.12	0.82
	60	-0.27	0.28	-1.02*	0.21	0.08	0.13	0.79
Japan	12	0.29	0.25	0.03	0.12	0.82*	0.14	0.74
	36	0.12	0.19	-0.19	0.16	0.46*	0.15	0.64
	60	-0.05	0.14	-0.48*	0.11	0.15	0.12	0.74
Netherlands	12	-0.07	0.21	0.02	0.16	0.79*	0.14	0.80
	36	-0.43	0.22	-0.24	0.15	0.38*	0.12	0.58
	60	-1.00*	0.22	-0.74*	0.15	-0.06	0.12	0.57
Sweden	12	0.27	0.28	0.01	0.18	0.73*	0.15	0.59
	36	0.37	0.29	-0.14	0.20	0.43*	0.25	0.53
	60	-0.01	0.34	-0.57*	0.22	-0.27	0.25	0.48
Switzerland	12	-0.83*	0.35	-0.25	0.30	-0.00	0.30	0.42
	36	-1.37*	0.43	-0.92*	0.39	-0.89*	0.36	-0.35
	60	-0.77*	0.37	-0.58	0.33	-0.80*	0.28	-0.29
United Kingdom	12	0.51	0.45	0.39	0.37	0.92*	0.26	0.68
	36	0.70	0.42	0.20	0.34	0.58*	0.26	0.64
	60	0.55	0.28	-0.11	0.25	0.25	0.20	0.76
USA	12	-0.54*	0.22	-0.30*	0.13	0.50*	0.13	0.73
	36	-0.47	0.25	-0.49*	0.16	0.17	0.20	0.47
	60	-0.23	0.26	-0.55*	0.15	0.02	0.21	0.46

* Significant at 5% level

SE(·) denotes the standard error, corrected for heteroskedasticity and serial correlation, of the estimated coefficient.

Estimated using instrumental variables with a constant, $R_{120,t} - R_{3,t}$, and $\pi_{3,t}$ as instruments.

Results are for monthly data, except Australian results are for quarterly data.