Inflation Uncertainty and the Recent Low Level of the Long Bond Rate

Yash P. Mehra

Many analysts and policymakers have been intrigued by the recently observed low levels of long-term interest rates. Figure 1 charts the actual and predicted levels of the nominal yield on ten-year U.S. Treasury bonds over 1994Q1 to 2005Q1; the predicted values were generated using the historical relationship that had existed between the long bond yield and several of its macroeconomic determinants including long-term inflation expectations, near-term outlook for the economy, and the stance of monetary policy. The prediction errors are also charted there. As one can see, for the past few years the actual long bond rate has remained consistently below what is predicted using these standard economic determinants.[1] Other analysts using somewhat different economic determinants have come to the same conclusion that the long bond rate has recently been substantially lower than can be explained by macroeconomic conditions.[2]

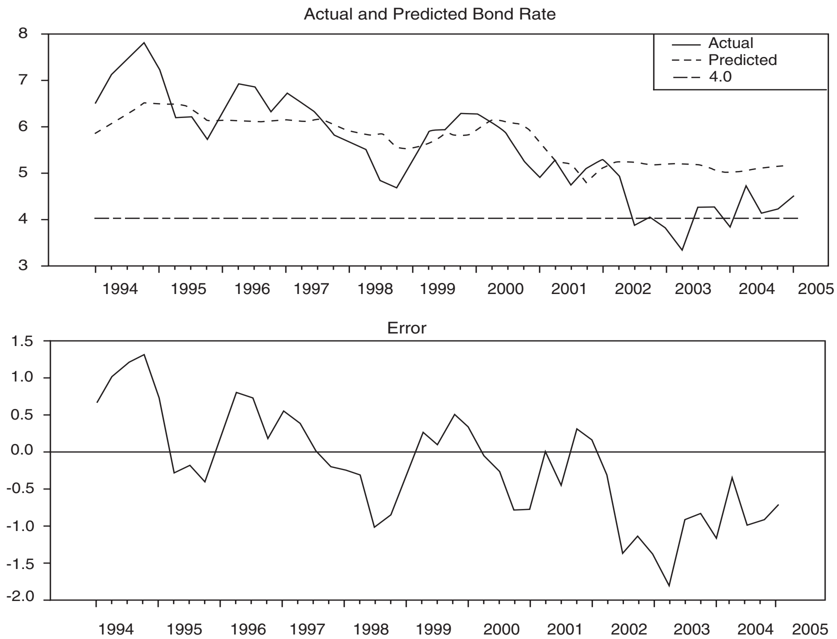
In order to explain the recent puzzling behavior of long-term interest rates, two alternative hypotheses have received prominent attention in the financial

The author thanks Juan Carlos Hatchondo, Hubert Janicki, Roy Webb, and John Weinberg for helpful comments. The views expressed are those of the author and do not necessarily represent those of the Federal Reserve Bank of Richmond or the Federal Reserve System.

[1] As discussed fully later, the reduced-form long bond equation used to generate the predicted values relate the long bond rate to long-term inflation expectations, near-term forecasts of real growth and inflation, and the surprise component of change in the fed funds rate target, denoted here as the baseline bond rate equation. This equation is estimated over 1984Q1 to 2004Q3 and simulated dynamically over 1994Q1 to 2005Q1, conditional on actual values of macroeconomic determinants and assuming the Fisher coefficient is unity. The predicted values charted in Figure 1 are the simulated values.

[2] See, for example, Warnock and Warnock (2005). Chairman Bernanke (2006) in his recent testimony to the U.S. Congress also notes that long-term interest rates have remained relatively low given recent strong real growth and rising short-term interest rates.

Figure 1 Baseline Bond Equation



press.[3] The first one attributes the current low level of the long bond rate to the lowering of the inflation risk premium. In particular, this hypothesis posits that as a result of the improved inflation performance of the U.S. economy, inflation uncertainty has declined, leading to lowering of inflation risk premiums, which is reflected in lower real and nominal bond yields.[4] The other hypothesis

[3] Some other hypotheses that have surfaced in the financial press have not been considered serious enough to warrant much attention. For example, one hypothesis involves the behavior of pension funds. This hypothesis attributes the recent decline in the long bond rate to increased demand for longer-term bond portfolios by pension funds and insurance companies that are needed to replenish their underfunded retirement plans. However, these funding shortfalls are not considered large enough to be able to explain the recent behavior of long-term interest rates. Another hypothesis posits that the current low level of the long bond rate may be signaling economic weakness. Most reduced-form interest rate models usually control for the influence of future real growth on current bond yields, yet those models still cannot account for the recent low level of the long bond rate.

[4] See, for example, Greenspan (2005), Kim and Wright (2005), Dudley (2006), and Bernanke (2006). Although several analysts attribute the low level of the long bond rate to lower bond risk premiums, they differ with respect to reasons for the collapse in risk premiums. Chairman Greenspan has focused on increased globalization and integration of financial markets as sources of the favorable inflation performance in many countries including the United States, whereas others (for example, Dudley 2006) attribute the favorable inflation performance to monetary policy. In contrast, Kim and Wright have emphasized the potential role of increased demand for U.S. Treasury

attributes recent declines in long-term interest rates to increases in purchases of U.S. Treasury securities by foreign central banks.[5]

This article develops an empirical test of the first hypothesis, using a reduced-form interest rate equation that links the long bond rate directly to macroeconomic variables, including an empirical proxy for inflation uncertainty. I focus on the first hypothesis for two reasons. First, despite the popularity of the first hypothesis in the financial press, it has not yet been formally investigated. In most previous research, the evidence in favor of the first hypothesis comes from the term structure model, indicating that term premiums have declined and that part of this decline is attributed to a decline in the inflation risk premium. This article, however, constructs a direct empirical measure of inflation uncertainty and examines whether the recent behavior of the long bond rate can be linked to the recent reduction in inflation uncertainty. Second, some previous research has indicated that the empirical evidence favoring the second hypothesis is fragile in the sense that the empirical evidence—the long bond rate is influenced by direct foreign capital inflows—is due to the most recent data.[6] In view of these considerations, I focus on the first hypothesis, but I do examine the robustness of results with respect to inclusion of foreign official purchases of U.S. Treasury securities in the list of macroeconomic determinants.

It is widely understood that investors holding long-term U.S. Treasury bonds bear an inflation risk, because actual inflation that is higher or lower than what they forecasted when they bought bonds would make their holding of bonds significantly less or more valuable. Hence, if there is considerable uncertainty about long-term inflation forecasts in the sense that the probability distribution of long-term inflation forecasts is widely dispersed, investors demand compensation for bearing the inflation risk, and hence long bond rates contain risk premiums.

Since we do not have a direct empirical measure of uncertainty about long-term inflation forecasts, this article constructs an empirical proxy making two identifying assumptions. The first assumption is that uncertainty about long-term inflation forecasts is positively correlated with uncertainty about short-term inflation forecasts, so that when investors become more uncertain about their short-term inflation forecasts, their uncertainty about long-term inflation forecasts also increases. The second assumption is that uncertainty about short-term inflation forecasts can be approximated by the mean squared error

---

securities relative to supply. The empirical work here focuses on domestic factors that might be at the source of the favorable inflation performance.

[5] See, for example, Wu (2005) and Warnock and Warnock (2005). Chairman Bernanke (2006) has focused instead on increased capital inflows arising as a result of an excess of desired global savings over the quantity of global investment opportunities that pay historically normal returns. The examination of the global savings glut hypothesis is beyond the scope of this article.

[6] See, for example, the evidence in Wu (2005) and Warnock and Warnock (2005).

(MSE) of short-term inflation forecasts, so that uncertainty about short-term inflation forecasts rises when the variance (in particular, the MSE) of ex-post short-term inflation forecast errors increases. Given these two assumptions, I examine the MSE of short-term inflation forecasts, using survey data on private-sector GDP inflation expectations. In particular, the article creates a time series on uncertainty about short-term inflation forecasts, using rolling three-year windows on the MSE of short-term inflation forecasts over 1984Q1 to 2004Q3.[7]

The resulting time series on uncertainty about short-term inflation forecasts has a clear downward trend over 1984 to 2004, which is consistent with the downward trend in mean and variance of short-term inflation forecasts. This trend suggests that reduction in short-term inflation uncertainty may reflect the good inflation performance of the U.S. economy; namely, short-term inflation uncertainty declined because inflation both steadily declined and became more predictable.

The article then estimates a reduced-form bond rate equation that links the long bond rate to macroeconomic variables, including the aforementioned empirical measure of uncertainty about short-term inflation forecasts. The results indicate the long bond rate is positively correlated with short-term inflation uncertainty over the full sample period of 1984Q1 to 2004Q3, suggesting that an increase in uncertainty about short-term inflation forecasts raises uncertainty about long-term inflation forecasts and hence may account for the presence of the inflation risk premium in the bond rate. However, the results also indicate that the estimated coefficient that measures the response of the long bond rate to short-term inflation uncertainty has declined since 2001Q4, implying that in recent years an increase in short-term inflation uncertainty is associated with a small-to-negligible increase in uncertainty about long-term inflation forecasts. In fact, the results are consistent with the hypothesis that the inflation risk premium embedded in the long bond rate has disappeared, thereby accounting in part for the current low level of the long bond rate.

As stated above, one of the identifying assumptions in the empirical work here is that uncertainty about long-term inflation forecasts is positively correlated with uncertainty about short-term inflation forecasts and that the magnitude of this positive correlation is stable over the sample period being studied. However, the result above—the correlation of the long bond rate with short-term inflation uncertainty has weakened in recent years—may be interpreted to mean that the identifying assumption made above does not hold for the complete sample period of 1984 to 2004; namely, while in the past an increase in short-term inflation uncertainty may have increased uncertainty about long-term inflation forecasts, it no longer does so. This development may be the

[7] Tulip (2005) uses this approach to investigate whether output has become predictable, using Greenbook forecasts.

consequence of increased Fed credibility. It is only recently that investors have become more confident that the current low and stable short-term inflation will continue in the long run so that a given increase in short-term inflation uncertainty now leads to a small-to-negligible increase in uncertainty about long-term inflation forecasts, and hence investors demand lower inflation risk premiums than before. This consequence of increased Fed credibility can be seen in the fact that it is only recently that both short- and long-term inflation forecasts have become fully anchored, in contrast to the early part of the sample period when they were not anchored.

The empirical work here that attributes the current low level of the long bond rate to a lower inflation risk premium is robust to the inclusion of foreign official capital inflows in the list of macroeconomic determinants of bond yields. The results do indicate the long bond rate is negatively correlated with this measure of foreign official capital inflows, however, this correlation is marginally significant and fragile, being absent in the period prior to the recent episode of increased capital inflows. Together, these results favor the hypothesis that attributes the recent low level of the long bond rate mostly to lowering of inflation risk premiums.

The rest of the article is organized as follows. In Section 1, I examine the behavior of uncertainty about short-term inflation forecasts, constructed using private-sector, ex-post inflation forecast errors. Section 2 contains discussion of a reduced-form interest rate equation that relates the long bond rate to macroeconomic variables. Section 3 presents empirical results, and concluding remarks are in Section 4.

1. A PRELIMINARY ANALYSIS: SOURCES OF DECLINE IN UNCERTAINTY ABOUT SHORT-TERM INFLATION FORECASTS

As indicated at the outset, if there is considerable uncertainty about long-term inflation forecasts, holders of long-term U.S. Treasury bonds bear an inflation risk and hence long bond yields have embedded in them inflation risk premiums. Since one does not have a direct empirical measure of uncertainty about long-term inflation forecasts, the article proceeds under the assumption that uncertainty about long-term inflation forecasts is positively correlated with uncertainty about short-term forecasts. This section constructs the empirical measure of uncertainty about short-term inflation forecasts and analyzes its behavior over the sample period of 1984Q1 to 2004Q3.

Measuring Uncertainty about Short-Term Inflation Forecasts

If inflation had been harder to forecast in the past, then it is likely to raise uncertainty about agents' current forecasts of expected future inflation rates.

Given this basic idea, the article examines ex-post inflation forecast errors, focusing on the MSE of one-to-four-quarters-ahead inflation forecasts. If the MSE of inflation forecasts increases over time, then it is likely to raise the variance of agents' current forecasts of expected future inflation rates and hence will lead to increased uncertainty about their mean inflation forecasts. For inflation forecasts, I use private-sector GDP inflation forecasts from the Philadelphia Fed's Survey of Professional Forecasters (denoted hereafter as SPF).[8] I use survey data because recent evidence indicates that surveys perform much better than some standard reduced-form inflation forecasting models in predicting future inflation.[9] Despite the evidence in Romer and Romer (2004) that Greenbook inflation forecasts are more accurate relative to private-sector forecasts, I use the latter because Greenbook forecasts are released to the public with a five-year delay, and hence bond yields are likely to reflect private-sector inflation expectations. Since surveys are used, I compute forecast errors using real-time data on actual inflation as in Romer and Romer (2004). I create time series on the MSE of one-to-four-quarters-ahead inflation forecasts, using rolling three-year windows over 1984Q1 to 2005Q3.[10] This time series is an empirical proxy measuring uncertainty about short-term inflation forecasts, denoted hereafter as short-term inflation uncertainty.

Figure 2 charts the rolling MSE of contemporaneous, one-quarter- and four-quarter-ahead inflation forecasts over 1984Q1 to 2004Q3.[11] As can be seen, the evidence of a decline in short-term inflation uncertainty is quite clear, as the MSE of inflation forecasts has drifted down intermittently since 1984. In particular, focusing on the MSE of the four-quarter-ahead inflation forecasts, short-term inflation uncertainty declined significantly first during the latter half of the 1980s, increased somewhat in the first half of the 1990s, and then again drifted lower beginning in the late 1990s.

Low Inflation, Great Moderation, and Short-Term Inflation Uncertainty

One plausible explanation of the decline observed in short-term inflation uncertainty over 1984Q1 to 2005Q3 is the good inflation performance of the U.S. economy due to Federal Reserve policy during this period. In particular, this explanation posits that, under Chairman Volcker and Chairman Greenspan,

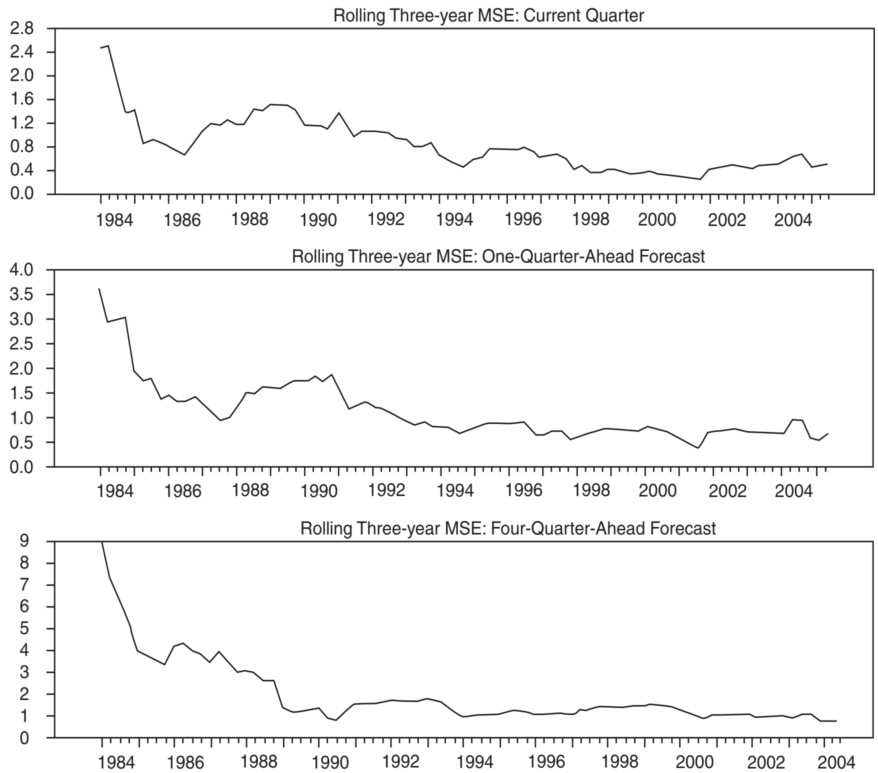
[8] Ideally, one needs to examine the MSE of ten-year-ahead Consumer Price Index (CPI) inflation forecasts. However, for the sample period 1984 to 2005Q3 studied here, it is not possible to generate enough observations on the forecast error. Hence, I focus on the MSE of short-term GDP inflation forecasts, assuming reduction in inflation uncertainty at short-term forecast horizons will lead to reduction in uncertainty at the long-term forecast horizon.

[9] Ang, Bekaert, and Wei (2006).

[10] I get qualitatively similar results using somewhat longer four-year rolling windows.

[11] Because I use lead data in generating forecast errors, the sample period ends in 2004Q3.

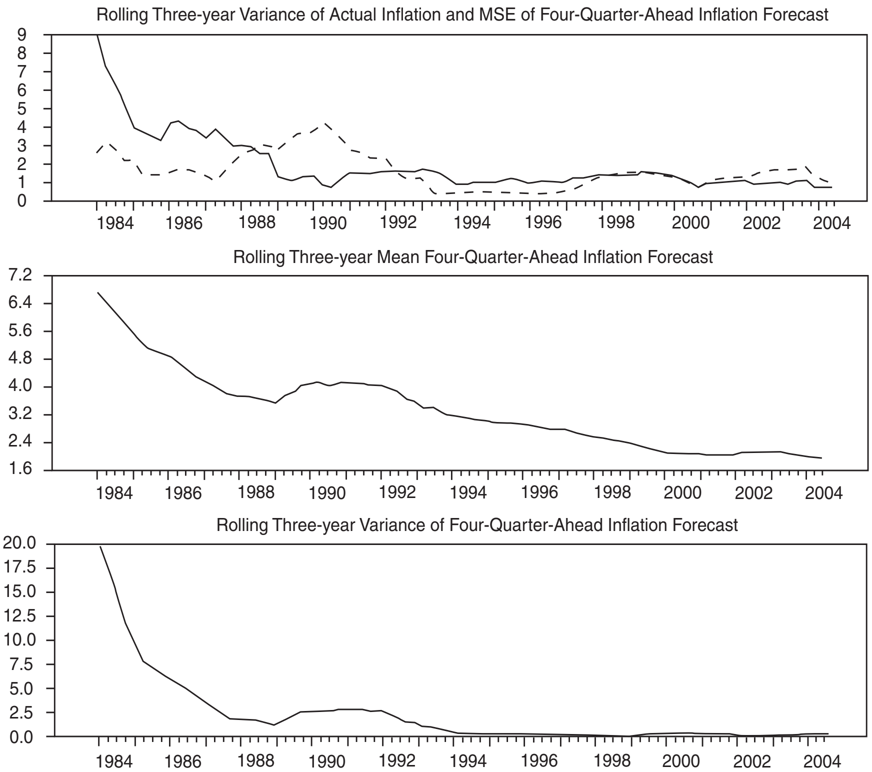
Figure 2 Uncertainty about GDP Inflation Forecasts



the Federal Reserve gradually had moved toward a policy framework that places a heavy weight on the requirement that the central bank keep inflation low and stable and hence the public's expectations of inflation under control. In addition, during this sample period the Fed has taken a number of steps toward increased transparency meant to reduce the public's uncertainty about the Fed's long-term inflation objective (Bernanke 2003, 2004). As a result, inflation has trended down and stabilized at low levels, thereby making inflation more predictable and contributing to lower short-term inflation uncertainty.

Figure 3 provides a visual confirmation of the hypothesis that decline in short-term inflation uncertainty is related to good inflation performance of the U.S. economy over 1984Q1 to 2005Q3. Focusing on the behavior of the four-quarter-ahead actual inflation and its forecast, the top panel in Figure 3 charts the variance of actual future inflation and the MSE of its forecast, calculated as before using rolling three-year windows. The middle panel charts the rolling mean of inflation forecasts, whereas the bottom panel charts the

Figure 3 GDP Inflation Volatility and Uncertainty



rolling variance of GDP inflation forecasts. The top and middle panels indicate that the series measuring the MSE of the inflation forecast has a downward trend that is shared by the series measuring the mean forecast but not by the series measuring the variance of actual inflation. This suggests that short-term inflation uncertainty declined not because inflation was less volatile but because inflation trended down.[12] Furthermore, the bottom panel indicates that variance of the predictable component of inflation also declined significantly during this period, suggesting increased predictability of inflation. Figure 3 thus provides a visual confirmation of the hypothesis that short-term inflation

[12] The argument that, over the sample period 1984Q1 to 2005Q3, the series measuring the variance of inflation does not depict a downward trend is not inconsistent with the evidence in previous research that volatility of inflation (measured by the variance of inflation) observed in the sample period since 1984 has been low relative to the one observed in the period before.

uncertainty declined because inflation both trended down and became more predictable.[13]

Current Low Short-Term Inflation Uncertainty and Anchoring of Long-Term Inflation Expectations

Figure 4 highlights another key feature of the recent favorable inflation performance: the current low level of short-term inflation uncertainty has accompanied decline in volatility of long-term inflation expectations. The top panel in Figure 4 plots the rolling MSE of four-quarter GDP inflation forecasts as before, and the other panel charts the rolling standard deviation of the ten-year-ahead CPI expected inflation. As one can see, during the past few years the standard deviation of the ten-year CPI inflation forecast has been zero, suggesting the recent stabilization and anchoring of long-term inflation expectations.

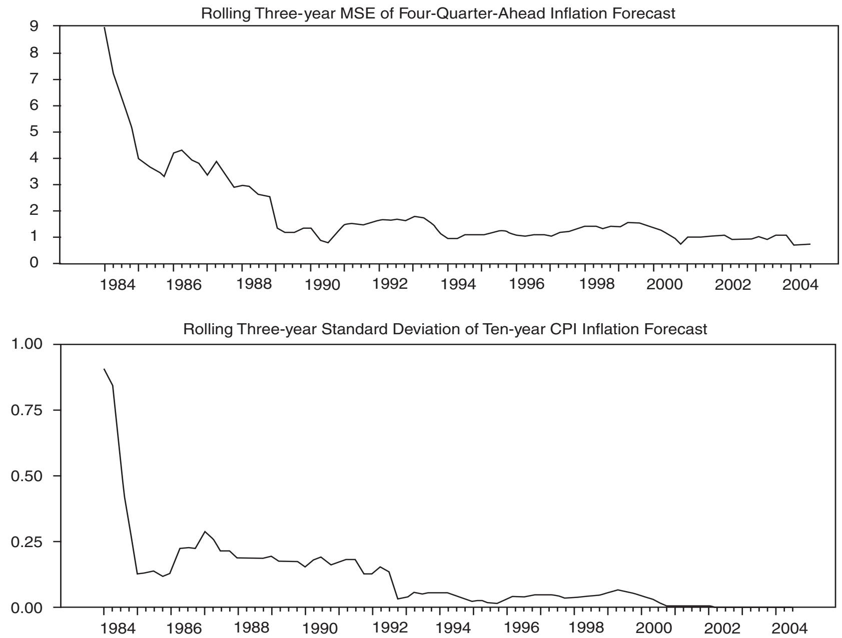
One simple explanation of this recent anchoring of long-term inflation expectations is that the recent period of low short-term inflation uncertainty has increased confidence that inflation will remain low and stable in the long run, which was absent before. This outcome may be the consequence of increased Fed credibility that occurred near the end of the sample period. During the early part of the sample period 1984 to 2005, though short-term inflation uncertainty declined to lower levels, long-term inflation expectations did not stabilize, reflecting the lack of Fed credibility. As one can see, during the early part of this sample period, both short-term and long-term inflation forecasts were not stabilized (see the bottom panel in Figure 3 and the lower panel in Figure 4). One implication of this different behavior of long-term inflation expectations is that the correlation of the long bond rate with short-term inflation uncertainty is likely to be weaker near the end of the sample period than it is during the early part, meaning a given rise in short-term inflation uncertainty is unlikely to raise uncertainty about long-term inflation forecasts as much as it did previously. This implication is confirmed by the empirical work in the following section, which attributes the recent decline

[13] As noted in Tulip (2005), the variance of actual future inflation is algebraically related to MSE as shown below.

$$\underbrace{\frac{1}{n}\sum_{n=1}^{12}(\pi_{t+4}-\bar{\pi})^2}_{Variance} = \underbrace{\frac{1}{n}\sum_{n=1}^{12}(e_{t+4})^2}_{MSE} + \underbrace{\frac{1}{n}\sum_{n=1}^{12}(f_{t+4}-\bar{\pi})^2}_{Predicted\ Variation} + \underbrace{\frac{2}{n}\sum_{n=1}^{12}(f_{t+4}-\bar{\pi})e_{t+4}}_{Covariance},$$

where π_{t+4} is actual four-quarter-ahead inflation, f is the survey forecast, e is the forecast error, and $\bar{\pi}$ is the sample mean. Hence, in the top panel, the distance between the line plotting variance and the line plotting MSE equals the sum of the last two terms. If we ignore the last term, the second term on the right-hand side of the equation above measures variance of the predictable component of inflation. The bottom panel in Figure 3 has charted the second term.

Figure 4 Stabilization of Inflation Expectations



in the inflation risk premium to reduced sensitivity of the long bond rate to uncertainty about long-term inflation forecasts.[14]

2. A REDUCED-FORM EMPIRICAL MODEL OF THE LONG BOND RATE

In this section, I discuss a reduced-form empirical equation that links the long bond rate to macroeconomic variables, including the empirical proxy for short-term inflation uncertainty. I also describe the data used to estimate the reduced-form equation.

[14] Figure 4 indicates that, for most of the 1990s, short-term inflation uncertainty remained low and stable, while long-term inflation expectations were stabilized. In order to uncover the relationship between the long bond rate and inflation uncertainty, one needs a period during which the potential explanatory variables, including the empirical measure of short-term inflation uncertainty, have varied considerably, as was the case during the early part of the sample period.

Long Run: The Fisher Equation

The reduced-form interest rate equation that underlies the empirical work here has two parts: a long-run and a short-run part. The long-run part, based on the Fisher equation, relates the level of the bond rate to long-term inflation expectations, risk premiums, and a risk-free long real rate, as in (1.3).

$$(1 - T_t)BR_t = rr_t + a_\pi \pi^e_t;\; a_\pi = 1, \tag{1.1}$$

$$rr_t = rr^* + a_r RP_t + \mu_t, \tag{1.2}$$

$$BR_t = (1/1 - T_t)[rr^* + a_r RP_t + a_\pi \pi^e_t + \mu_t];\; a_\pi = 1, \tag{1.3}$$

where BR is the long bond rate; T_t is the marginal tax rate on interest income in period t; rr_t is the after-tax expected long real rate; rr^* is the after-tax, risk-free expected long real rate; RP is a risk premium variable; π^* is long-term inflation expectations; and μ is the stationary disturbance term. Equation 1.1 is just the long-run Fisher equation that relates the after-tax long bond rate to the expected long real rate and inflation expectations. Equation 1.2 says the expected long real rate is mean stationary once we account for the presence of risk premiums in bond yields. If we substitute (1.2) into (1.1), one gets equation (1.3), which relates the level of the bond rate to long-term inflation expectations, risk premiums, and a risk-free long real rate.

The coefficient a_π is the after-tax Fisher coefficient that measures the response of the after-tax bond rate to inflation expectations and is generally assumed unity. The key point to note is that in the presence of taxes on interest income, the long bond rate should rise during an inflation episode by an amount that exceeds expected inflation sufficiently to compensate lenders both for their loss of capital due to inflation and for the taxation of interest income. Hence, in the presence of the tax effect, the before-tax Fisher coefficient $(a_\pi/(1-T_t))$ is likely to exceed unity, its exact magnitude varying with the marginal tax rate on interest income.[15] Furthermore, a significant component of risk premiums embedded in bond yields is likely to be inflation risk, arising as a result of unpredictable movements in long-term expected inflation.

Short Run: Short-Run Changes in the Bond Rate are Dominated by Changes in the Outlook for the Economy and the Stance of Monetary Policy

The bond rate equation given in (1.3) is long run and is motivated using the Fisher equation, in which the level of the long bond rate is related to the risk-

[15] Tanzi (1980).

adjusted expected long real rate and expected inflation. The expected long real rate is, however, unobservable. Recent research that has expanded term structure models of bond yields to include macroeconomic factors suggest that changes in the expected long real rate reflect changes in expected future short rates, which in turn are likely to be correlated with changes in the outlook for the economy and changes in the current and future stance of monetary policy.[16] In order to control for influences of other macroeconomic variables on the long bond rate, I consider the following short-run, error-correction specification of the bond rate equation (Mehra 1984, 1994):

$$\Delta(1-T_t)BR_t = f_0 + f_{\Delta rp}\Delta RP_t + f_{\Delta\pi}\Delta\pi_t^e + \sum_{h=1}^{k} f_{1rs}\Delta\dot{y}_{t+h}^e \qquad (2)$$

$$+\sum_{h=1}^{k} f_{2rh}\Delta\dot{P}_{t+h}^e + f_{3r}u\Delta FFR_t - f_{ec}\mu_{t-1} + \varepsilon_t,$$

$$\text{where } \mu_{t-1} = (1-T_{t-1})BR_{t-1} - rr^* - a_r RP_{t-1} - a_\pi \pi_{t-1}^e$$

where h is the forecast horizon, $\Delta\dot{y}_{t+h}^e$ is change in the h-quarter-ahead forecast of real growth, $\Delta\dot{P}_{t+h}^e$ is change in the h-quarter-ahead forecast of the inflation rate, and $u\Delta FFR$ is the surprise component of the change in the federal funds rate. Equation 2 relates short-run changes in the after-tax bond rate to three sets of economic variables: the first set contains first differences of economic variables that enter the long-run Fisher equation here (ΔRP_t, $\Delta\pi_t^e$); the second set contains variables measuring changes in the outlook for the economy and stance of monetary policy ($\Delta\dot{y}_{t+h}^e$, $\Delta\dot{P}_{t+h}^e$, $u\Delta FFR_t$); and the third set contains only a lagged error-correction variable (μ_{t-1}), measured as a gap between the actual level of the long rate and the level consistent with the long bond equation. The coefficient on the error-correction variable in (2) is hypothesized to be negative, meaning the bond rate declines if in the previous period the actual bond rate was high relative to the level consistent with its long-run determinants specified in (1.3).

In the empirical bond equation (2), changes in the outlook for the economy are measured as changes in private-sector forecasts of real growth and inflation. The expected signs of coefficients that appear on changes in anticipated real

[16] The reduced-form empirical bond rate equation estimated here is in spirit based on the recent empirical work that links bond yield dynamics to macroeconomic variables. To explain it further, as in finance literature, bond yields are modeled as risk-adjusted averages of expected future short rates. Expectations of future short rates, however, depend in part on expectations of future macroeconomic variables, which are generated using either a structural or a VAR model of the economy. This methodology thus relates bond yield dynamics to macroeconomic variables. See Clouse (2004) and Hordahl, Tristani, and Vestin (2006) for an empirical illustration of this joint econometric modeling of macroeconomic and term-structure dynamics and Diebold, Piazzesi, and Rudebusch (2005) for a summary of this literature.

growth and inflation variables in (2) are positive, suggesting that accelerated future real growth or inflation is likely to lead to higher future short real rates and hence to a higher long real rate. The positive correlation between the long real rate and higher anticipated real growth or inflation may arise as a result of "lean-against-the-winds" monetary policy strategy; namely, the private sector expects the Federal Reserve to raise the funds rate target when real growth or inflation is anticipated to accelerate, leading to higher future short real rates.

The impact of monetary policy actions on the expected long real rate is captured by the "surprise" component of changes in the funds rate target. Recent research indicates that bond yields respond to this surprise component and that the nature of the yield curve response depends crucially on the interpretation of market participants' reasons behind the policy move. If the policy move is interpreted to reveal "new" information about the outlook for inflation and real growth, interest rates of all maturities, including the long end, move in the same direction as the funds rate target. If, on the other hand, market participants view the policy move as driven by changes in the central bank's preferences (such as a shift to a more inflation-averse policy), long and short rates move in opposite directions (Ellingsen and Soderstrom 2001, 2004; Gurkaynak, Sack, and Swanson 2005). Thus, this literature suggests that the response of the long bond rate to policy is time varying, and the bond rate may actually fall if bond market participants interpret policy tightening as resulting in lower inflation in the long run.

Combining Long- and Short-Run Parts

Equation (2) is the short-term bond equation that relates changes in the bond rate to (a) "changes" in the private-sector outlook for real growth and inflation; (b) the surprise component of changes in the funds rate target; (c) changes in long-term inflation expectations and risk premiums; and (d) the lagged value of an error-correction variable, measuring discrepancies between the actual level of the bond rate and the level consistent with the long-run Fisher equation (1.3). If we substitute the expression for the error-correction variable into (2), we get a reduced-form long bond equation as in (3).

$$\Delta(1-T_t)BR_t = f_0 + f_{\Delta rp}\Delta RP_t + f_{\Delta\pi}\Delta\pi^e_t + \sum_{h=1}^{k} f_{1rh}\Delta\dot{y}^e_{t+h} + \sum_{h=1}^{k} f_{2rh}\Delta\dot{P}^e_{t+h} \tag{3}$$

$$+ f_{3r}u\Delta FFR_t - f_{ec}(1-T_{t-1})BR_{t-1} + f_{ec}rr^* + f_{ec}a_r RP_{t-1} + f_{ec}a_\pi\pi^e_{t-1} + \varepsilon_t$$

$$\Delta BR = (1/(1-T_t))[\delta_0 + f_{\Delta rp}\Delta RP_t + f_{\Delta\pi}\Delta\pi^e_t + \sum_{h=1}^{k} f_{1rh}\Delta\dot{y}^e_{t+h} + \sum_{h=1}^{k} f_{2rh}\Delta\dot{P}^e_{t+h}$$

$$+ f_{3r}u\Delta FFR_t - f_{ec}(1-T_{t-1})BR_{t-1} + f_{ec}a_r RP_{t-1} + f_{ec}a_\pi\pi^e_{t-1}] + \varepsilon_t$$

$$\text{where } \delta_0 = f_0 + f_{ec} rr^* \ .$$

Three key features of the short-term bond equation (3) need to be highlighted. The first is the equation relates changes in the bond rate to changes and levels of some macro variables, in particular long-term inflation expectations. As a result, it is possible to recover estimates of the coefficients of the long-run Fisher equation from the short-run, reduced-form equation. Thus, if we estimate the unrestricted reduced-form (3), the after-tax Fisher coefficient a_π is recovered as the estimated coefficient ($f_{ec} a_\pi$) on lagged inflation expectations (π^e_{t-1}) divided by the absolute value of the estimated coefficient (f_{ec}) on the lagged bond rate (BR_{t-1}).[17] The second feature to highlight is that the short-run response of the long bond rate to macroeconomic variables is likely to vary over time, as the marginal tax rate on interest income is not constant over time. The third feature to note is that in a steady state where the private sector's near-term real growth and inflation expectations are stabilized and where there are no monetary policy surprises, the long bond rate will converge to the level determined by the Fisher equation.[18]

Estimating the Bond Rate Equation: Description of the Data

The long bond equation (3) is estimated using quarterly data over 1984Q1 to 2005Q3. The long bond rate (*BR*) is the nominal yield on ten-year U.S. Treasury bonds observed in the third month of the quarter. The measure of monetary policy is the funds rate observed in the third month of the quarter. The survey forecast of the ten-year-ahead CPI expected inflation rate (π^{10}_t) is used as a proxy for long-term inflation expectations. The private-sector outlook for the economy is measured by the Survey of Professional Forecasters' (SPF) near-term forecasts of real growth and inflation, currently conducted by the Philadelphia Fed and released by the end of the second month of the quarter. Inflation uncertainty is measured by the series on inflation unpredictability, discussed in the previous section. The tax rates used are from the series on the (average) marginal tax rate on interest income given in the NBER's TAXSIM model.[19]

In some previous research, the surprise component of the change in the funds rate has been calculated using data from the fed funds futures market (Kuttner 2001). I, however, follow the strategy in Romer and Romer (2004)

---

[17] Estimate of the constant term in the long Fisher equation is not identified.

[18] To be specific, consider a steady state in which coefficients in (3) assume values given below: $f_{\Delta rp} = f_{\Delta\pi} = f_{1rs} = f_{2rs} = f_{3r} = a_{rp} = 0$, $f_{ec} = 1$, then the long bond rate equals the risk-free long expected real rate and expected inflation.

[19] See Feenberg and Coutts (1993) for more details. The tax series used is the one that measures the federal marginal tax rates on interest income.

and construct a different measure of monetary policy surprise. Romer and Romer develop a measure of policy shocks by removing the component of changes in the funds rate target that are due to past and anticipated developments in the economy, and they capture the effect of anticipated developments on the funds rate target using Greenbook forecasts of real growth and inflation. So, Romer and Romer's measure of policy shocks is free of movements anticipated by the Federal Reserve.

However, what one needs here is a measure of policy shocks that are free of movements anticipated by bond market participants. Hence, I purge the funds rate target of anticipated movements by using private-sector forecasts of real growth and inflation. In particular, I purge the endogenous and anticipated movements in the funds rate by running the following regression.

$$\Delta FFR_t = \alpha_0 + \sum_{h=1}^{k} \alpha_{1s} \Delta \dot{y}^e_{t+h} + \sum_{h=1}^{k} \alpha_{2h} \dot{P}^e_{t+h} + \alpha_3 \dot{y}^e_t + \alpha_4 \dot{p}^e_t + \sum_{s=1}^{k} \alpha_{3s} \Delta y_{t-s} \tag{4}$$

$$+ \sum_{s=1}^{k} \alpha_{6s} \Delta P_{t-s} + \alpha_7 FFR_{t-1} + u\Delta FFR_t,$$

where *FFR* is the actual funds rate, y is actual real growth, p is actual inflation rate, $u\Delta FFR$ is the residual, and the rest of the variables are defined as before. The residual $u\Delta FFR$ from the estimated regression (4) is the measure of the surprise component of changes in the funds rate target. Since the funds rate target is the average value of the actual funds rate observed in the third month of the quarter, the regression (4) provides estimates of changes in the funds rate anticipated based on the latest information available to bond market participants.

The funds rate equation (4) is estimated over 1983Q1 to 2005Q3 and is reproduced below:

$$\Delta FFR_t = \underset{(3.4)}{-.63} + \sum_{h=0}^{4} \underset{(0.2)}{.03} \Delta \dot{y}^e_{t+h} + \sum_{h=0}^{4} \underset{(2.6)}{.63} \Delta \dot{P}^e_{t+h} + \underset{(4.4)}{.19} \dot{y}^e_t \tag{5}$$

$$+ \underset{(2.0)}{.17} \dot{P}^e_t + \underset{(1.6)}{.04} \Delta y_{t-1} + \underset{(2.2)}{.10} \Delta P_{t-s} - \underset{(1.9)}{.06} FFR_{t-1} + u\Delta FFR_t$$

$$\text{Adjusted } R^2 = .44,$$

where all variables are defined as before. As one can see, changes in the funds rate target are significantly correlated with changes in forecasts of GDP inflation, besides being correlated with changes in lagged inflation and real growth. Changes in the funds rate target are also correlated with forecast levels of GDP inflation and real growth. In the empirical work here, the residual from the estimated funds rate equation (5) is used as a proxy for the surprise component of change in the funds rate target.[20]

As indicated above, the bond equation (3) allows for the presence of the tax effect. Hence, the equation is estimated using data observations on variables that have been pre-multiplied by the time-varying tax series $(1/(1 - T_t))$.[21] The bond rate equation is estimated by ordinary least squares.

3. EMPIRICAL RESULTS

This section discusses estimates of the bond equation (3) over 1984Q1 to 2004Q3. In order to examine robustness of results, I also estimate the bond equation over a shorter sample period, 1984Q1 to 2000Q4, excluding observations pertaining to the most recent sub-period of low bond yields and increased foreign official inflows into U.S. Treasury securities.

Estimates of the Bond Rate Equation: With and Without Inflation Uncertainty

Table 1 contains estimates of the bond rate equation (3) over two sample periods, 1984Q1 to 2000Q4 and 1984Q1 to 2004Q3. The columns labeled (1.1) and (1.2) contain estimates of what is denoted hereafter as the "baseline" bond equation. In the baseline bond equation, the long-run part contains long-term inflation expectations and the short-run part includes macroeconomic variables measuring changes in the outlook for the economy and monetary policy. If we focus on estimates of the baseline equation for the shorter period of 1984Q1 to 2000Q4, they suggest the following observations. First, short-term changes in the bond rate are significantly correlated with changes in long-term inflation expectations and the short-term outlook for real growth and GDP inflation. The estimated coefficients that appear on these macroeconomic variables are statistically significant and correctly signed, indicating that accelerations in long-term expected inflation and short-term forecasts of real growth and inflation are associated with a higher bond rate.

Second, the long bond rate is positively correlated with the surprise component of the change in the funds rate, suggesting that policy tightening is associated with a rising bond rate. The estimated coefficient on policy surprises has a positive sign, suggesting that on average policy surprises have conveyed new information about the state of the economy.

---

[20] The first four estimated autocorrelation coefficients of the monetary policy surprise series are .20, .15, .06, and .02, which are insignificantly different from zero, suggesting that time series in fact do measure policy surprises.

[21] See Tanzi (1980) and Mehra (1984) for details.

Table 1 Estimates of the Bond Rate Equation

Independent Variables	Dependent Variable: ΔBR_t Sample Period Ending in (1.1) 2000Q4	(1.2) 2004Q3	(2.1) 2000Q4	(2.2) 2004Q3	(3.1) 2004Q3	(3.2) 2004Q3
$const.$						
BR_{t-1}	-.18 (2.6)	-.21 (3.2)	-.24 (3.4)	-.25 (3.8)	-.30 (4.4)	-.27 (4.3)
π^{10}_{t-1}	.28 (2.6)	.32 (3.0)	.23 (2.3)	.26 (2.5)	.27 (2.6)	.27 (4.3)
RP_{t-1}			.09 (2.4)	.09 (2.3)	.10 (2.8)	.10 (3.5)
$DU * RP_{t-1}$					-.34 (2.2)	-.10 (3.5)
$\Delta\pi^{10}_t$	.28 (1.9)	.30 (2.1)	.23 (1.6)	.24 (1.7)	.27 (1.7)	.24 (1.8)
Δy^e_{t+s}	.24 (1.8)	.18 (1.6)	.25 (1.9)	.18 (1.7)	.22 (2.1)	.20 (1.9)
$\Delta \dot{P}^e_{t+s}$	.45 (2.3)	.41 (2.2)	.48 (2.5)	.43 (2.4)	.45 (2.5)	.44 (2.5)
$u\Delta FFR_t$	.28 (2.4)	.19 (1.8)	.27 (2.4)	.20 (2.0)	.16 (1.6)	.19 (1.9)
a_π	1.57	1.54	1.0	1.0	.92	1.0
a_r			.39	.35		.35; 0.0[a]
R^2	.25	.20	.31	.25	.28	.28
SER	.530	.537	.509	.522	.509	.510

Notes: The reported coefficients (with t-values in parenthesis) are from the bond rate equation (4) of the text estimated over the sample period that begins in 1984Q1 but ends as indicated above. BR is the ten-year bond rate, π^{10} is the ten-year-ahead survey inflation forecast, RP is an inflation risk variable measured as the MSE of forecast errors, Δy^e_{t+s} is the average of zero-to-four-quarter-ahead (survey) real growth forecasts, $\Delta \dot{P}^e_{t+s}$ is the average of zero-to-four-quarter-ahead (survey) GDP inflation forecasts, $u\Delta FFR_t$ is the surprise component of change in the funds rate, DU is a dummy variable defined as unity over 2001Q4 to 2005Q3 and zero otherwise, R^2 is adjusted-R squared, and SER is the standard error of estimate. a_π is the long-term after-tax coefficient on ten-year expected inflation (Fisher coefficient) and a_r is the long-term coefficient on the inflation-related risk variable. All equations are estimated by ordinary least squares, using time series data pre-multiplied by (1/(1-Tax_t)), where Tax_t is the marginal tax rate on interest income.

a: post-break a_r

Third, the estimated after-tax Fisher coefficient a_π that measures the long-term response of the bond rate to inflation expectations is positive and far above unity. Since the baseline bond equation is estimated without controlling for the potential influence of inflation uncertainty on the long bond rate, the estimated Fisher coefficient may be biased upward, capturing in part the inflation risk premium embedded in the long bond yield.[22]

Finally, the above-noted three observations about the relationship between the long bond rate and macroeconomic variables continue to hold if we con-

[22] The sign of bias in the estimated Fisher coefficient is positive because inflation risk, which is omitted from the regression, is likely to be positively correlated with the level of expected inflation; namely, inflation uncertainty is large if expected inflation is high and variable.

sider estimates of the baseline equation over the full sample period given in the column labeled (1.2).

The columns labeled (2.1) and (2.2) in Table 1 contain estimates of the baseline equation augmented to include the empirical measure of short-term inflation uncertainty. Three results need to be highlighted. The first one is that the long bond rate is positively correlated with short-term inflation uncertainty, as the estimated coefficient on the pertinent variable is positive and statistically different from zero.[23] The estimated coefficient on short-term inflation uncertainty has a positive sign, suggesting that an increase in uncertainty about short-term inflation forecasts raises uncertainty about long-term inflation forecasts and hence may account for the presence of the inflation risk premium in the bond rate. The second result to note is that estimates of coefficients on other macroeconomic variables remain mostly unaffected when the bond equation is estimated controlling for the influence of inflation uncertainty, with the exception of the coefficient that appears on the lagged level of inflationary expectations (compare estimates across columns labeled [1.1] through [2.2]). The estimated after-tax Fisher coefficient is now close to unity (the p-value of the null hypothesis that $a_\pi = 1$ is .90, leading to the acceptance of the hypothesis), suggesting that failure to control for the presence of the inflation risk premium yields an unduly large estimate of the Fisher coefficient. Finally, the results appear robust across two sample periods considered here. In particular, the estimated coefficient on short-term inflation uncertainty remains positive and statistically significant in both sample periods, suggesting the result that inflation uncertainty matters in determining the long bond yield is not due to the most recent data.

Testing Stability of the Bond Rate Equation: Disappearance of the Inflation Risk Premium

Even though estimates of the baseline equation augmented with inflation uncertainty as reported in Table 1 appear similar across two sample periods, I now formally test parameter stability of the bond equation. As discussed earlier, one popular explanation of the current low level of the bond rate is that bond market participants are now demanding lower inflation risk premiums than before. Figures 2 and 3 indicate that uncertainty about short-term inflation forecasts declined steeply during the early part of the sample period 1984Q1 to 2004Q3 and so did variances of both GDP and long-term CPI inflation forecasts. However, during the early part, both short-term inflation

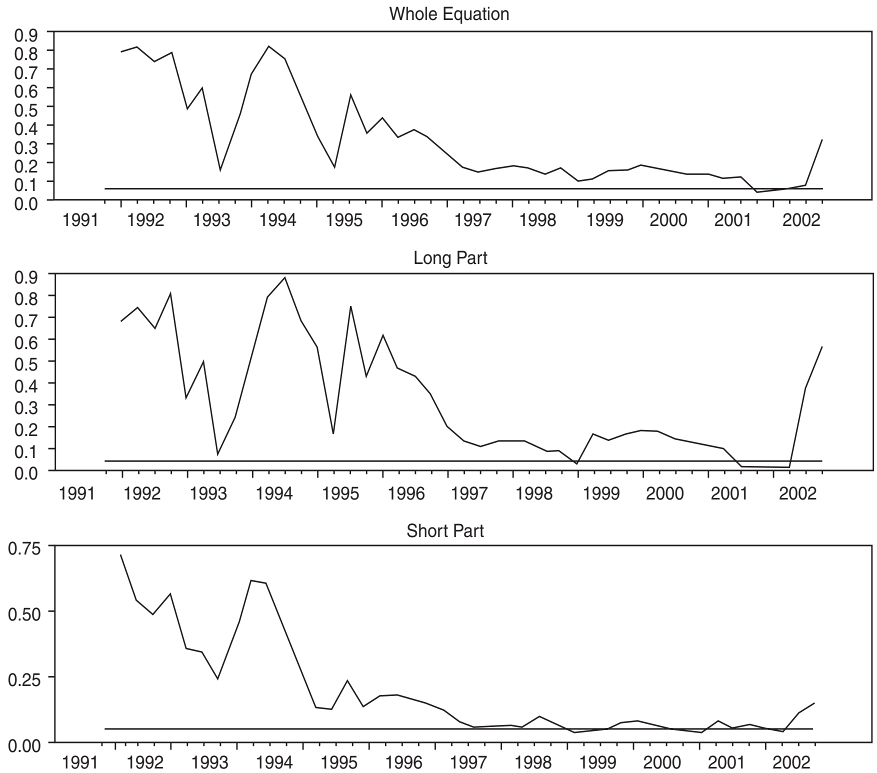
[23] The preliminary empirical work indicated that the long bond rate is positively correlated with the lagged level of the empirical measure of inflation uncertainty. First differences of this variable do not enter the bond equation. Together these results imply that inflation uncertainty enters the long-run part of the bond equation.

uncertainty and variances of both GDP and long-term CPI inflation forecasts remained fairly high, meaning uncertainty about long-term inflation forecasts remained high and long-term inflation expectations remained highly variable. Since then, short-term inflation uncertainty has declined, although modestly, and this modest decline in short-term inflation uncertainty has been accompanied by a significant reduction in the volatility of inflation expectations. In particular, the standard deviation of the ten-year-ahead CPI expected inflation has hovered around zero during the past few years, suggesting that market participants expect inflation to remain low and stable in the long run (see Figure 4). These considerations suggest that correlation of the long bond rate with short-term inflation uncertainty, which is a proxy for its correlation with uncertainty about long-term inflation forecasts, may not be stable over the sample period, 1984Q1 to 2004Q3. In particular, the coefficient a_r that measures the long-term response of the bond rate to short-term inflation uncertainty may have declined in recent years, because an increase in short-term inflation uncertainty may not raise uncertainty about long-term inflation forecasts as much as it did previously. Hence, I formally test parameter stability, using the Chow test with the break date treated as unknown over 1994Q1 to 2002Q4.

Figure 5 plots p-values of a Chow test for stability of different coefficients in the augmented bond equation as a function of the break date over 1994Q1 to 2002Q4. Panel A in Figure 5 plots the p-value of a Chow test where the null hypothesis is that all coefficients of the long bond rate equation are stable against the alternative that they have changed at the given date; panel B plots the p-value for stability of coefficients in the long-run part (coefficients on the constant term, inflation uncertainty, and long-term inflation expectations); and panel C plots the p-value for stability of coefficients in the short-run part (coefficients on changes in anticipated real growth and inflation and the surprise component of the change in the funds rate). The dashed line indicates a p-value of .05. In Figure 5, one main observation is that there is evidence of parameter instability only in the long-run part of the bond equation, suggesting that coefficients that appear on inflation uncertainty and long-term inflation expectations have changed, with the break date being 2001Q4. I assume the after-tax Fisher coefficient a_π has not changed and equals unity, because bond investors must be compensated for expected inflation even if they expect inflation to remain low and stable forever. Hence, I capture the break in the long-run part of the equation by allowing a different coefficient on short-term inflation uncertainty, because bond market participants may demand a lower inflation risk premium if they expect inflation to remain low and stable in the long run.[24]

[24] The alternative—that investors would not want to be compensated for expected inflation—is not reasonable.

Figure 5 P-Values for Chow Test: Baseline Equation With Inflation Uncertainty



Columns (3.1) and (3.2) in Table 1 present the estimated augmented bond equation that allows for the presence of a break in the coefficient on inflation uncertainty, captured here by including a dummy variable interacting with lagged inflation uncertainty. Column (3.1) contains unrestricted estimates, whereas column (3.2) contains estimates under the restrictions that the after-tax Fisher coefficient a_π is unity and that the coefficient on inflation uncertainty is positive before 2001Q4 but zero thereafter. The p-value for the null hypothesis that the Fisher coefficient a_π equals unity and the risk coefficient a_r is zero is .28, which is large, leading to the acceptance of the null. As shown, the estimated coefficient on the slope dummy variable is negative and statistically different from zero, suggesting that the long bond rate has become less sensitive to inflation uncertainty in recent years. In fact, estimates are consistent with the disappearance of the inflation risk premium in the long bond rate. In the pre-break period of 1984Q1 to 2001Q3, the average inflation risk premium is estimated to be about .98 of a percentage point, whereas, in the post-break period, the average risk premium is zero.[25]

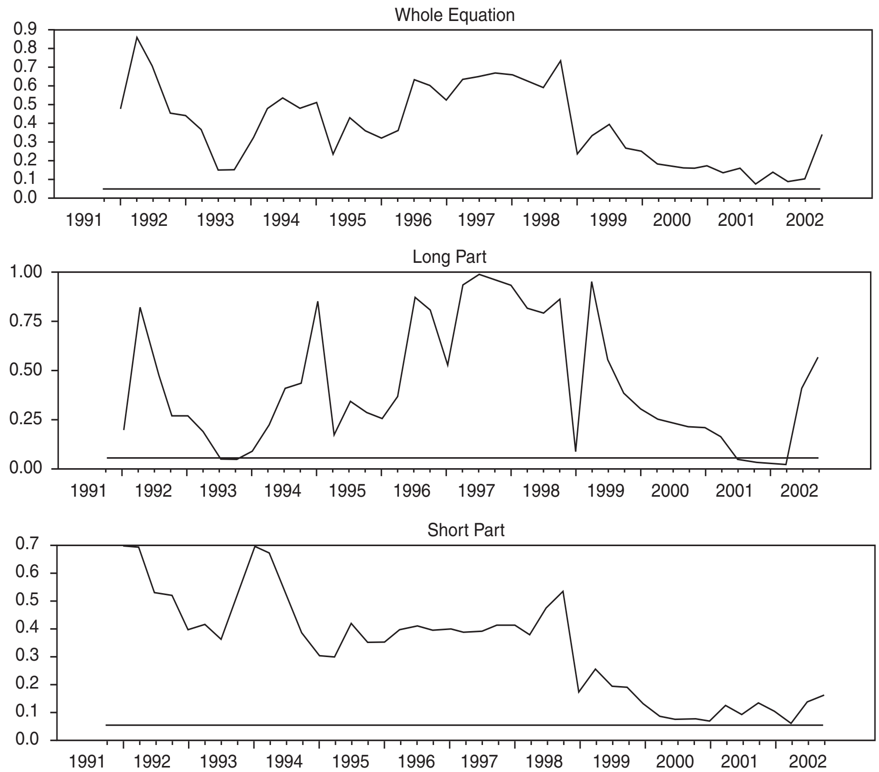
An Alternative Test of Lower Inflation Risk Premiums: Testing for a Shift in the Fisher Coefficient

The key result here is that the long bond rate is no longer correlated with the empirical measure of short-term inflation uncertainty, indicating the disappearance of inflation risk premiums from bond yields. But the aforementioned result is derived using the bond rate equation in which inflation uncertainty is measured by the MSE of short-to-medium-term GDP inflation forecasts. I now consider an alternative test of the hypothesis that inflation risk premiums have declined, using only the baseline bond equation. The basic idea behind the test is that if the bond rate equation is estimated without including a direct empirical measure of inflation uncertainty, then the estimated after-tax Fisher coefficient is likely to be above unity, because bond market participants must be compensated for inflation as well as for inflation-related risk. Hence the hypothesis inflation risk premiums that have declined can be tested by examining the temporal stability of the after-tax Fisher coefficient. Under the null hypothesis that inflation risk premiums have disappeared in recent years, the after-tax Fisher coefficient should now be closer to unity than it has been before.

For the full sample period 1984Q1 to 2004Q3, the estimated baseline bond equation is already reported in the column labeled (1.2) in Table 1. As one can see, the estimated after-tax Fisher coefficient is 1.5, far above unity, reflecting in part the presence of inflation-related risk premiums. Figure 6, which is similar to Figure 5, re-examines parameter stability of the baseline equation and plots p-values of a Chow test for stability of different coefficients as a function of the break date over 1994Q1 to 2002Q4. As can be seen, there is evidence of parameter instability not in the short-run part but in the long-run part of the bond rate equation, suggesting that the coefficient on long-term expected inflation has changed, with the break date being 2001Q4. Given such evidence of instability, I re-estimate the bond equation over 1984Q1 to 2004Q3, allowing the presence of a different Fisher coefficient since 2001Q4 and using a slope dummy. The estimated baseline bond equation is reported

[25] The magnitude of the inflation-related risk premium at time t is simply the long-term coefficient on inflation uncertainty times period t value of the time series measuring inflation uncertainty. In the pre-break period, the long-term coefficient on inflation uncertainty is .35 and the sample mean of the MSE of the four-quarter-ahead inflation forecast is 2.8 percentage points, suggesting that the average inflation risk premium over 1984Q1 to 2001Q4 is about 98 basis points. In the post-break period, however, the long-term coefficient on inflation uncertainty is not different from zero, suggesting that the inflation risk premium has disappeared.

Figure 6 P-Values for Chow Test: Baseline Bond Equation



below in (6).

$$\Delta BR_t = \underset{(0.4)}{.04} + \underset{(3.4)}{.32}\pi^{10}_{t-1} - \underset{(2.4)}{.10}(\pi^{10}_{t-1} * DU_{t-1}) - \underset{(3.5)}{.23}BR_{t-1} \tag{6}$$

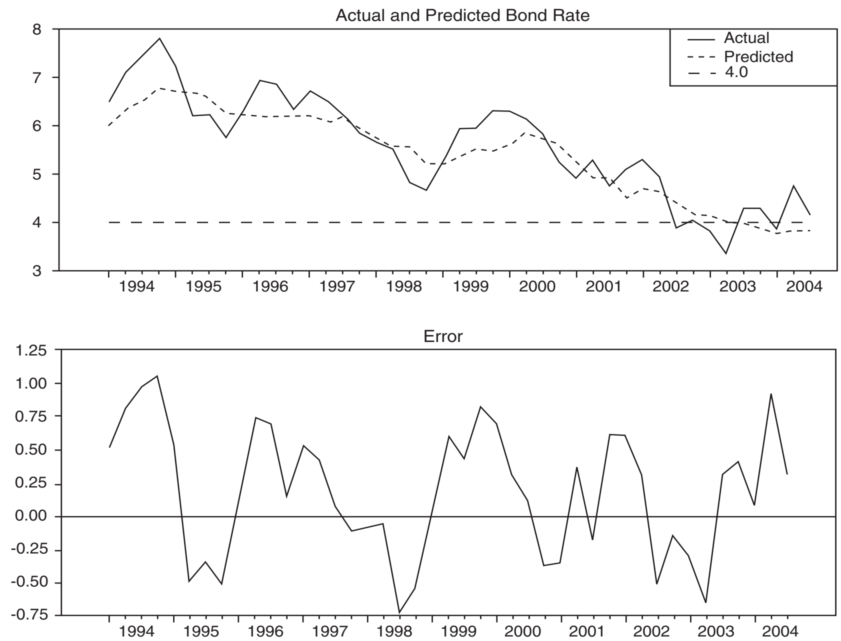
$$+\underset{(2.1)}{.29}\Delta\pi^{10}_t + \underset{(1.8)}{.19}\Delta\dot{y}^e_t + \underset{(2.1)}{.38}\Delta\dot{P}^e_t + \underset{(1.8)}{.19}u\Delta FFR_t.$$

Fisher Coefficient: $a_\pi = 1.42$(Pre-break) Adjusted $R^2 = .22$ $SER = .526$

=1.00(Post-break),

where DU is a dummy variable defined as unity over 2001Q4–2005Q1 and zero otherwise and where other variables are defined as before (see Table 1). As shown, the estimated after-tax Fisher coefficient is now unity and is consistent with the reduced magnitude of the inflation risk premium. Since the bond rate equation is estimated in first difference form, this reduction in the magnitude of the Fisher coefficient will result in reducing the level of the

Figure 7 Baseline with Lower Fisher Coefficient

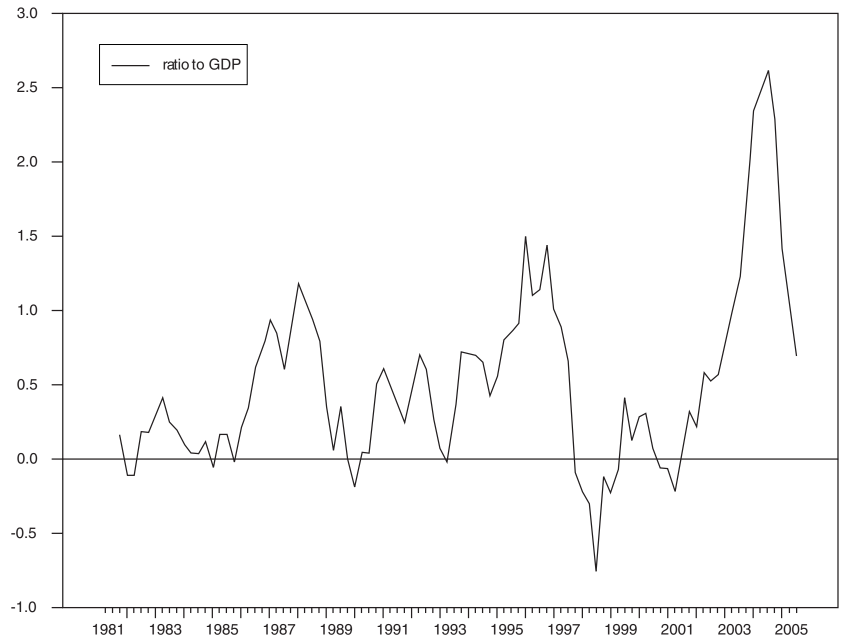


long real rate associated with long-term inflation expectations. The survey forecast of the ten-year-ahead CPI inflation rate has hovered around a narrow 2 percent to 2.5 percent range in recent years. Given that the magnitude of the Fisher coefficient declined by about 40 basis points, the reduction in after-tax real and nominal bond yields that can be attributed to reduction in inflation-related risk premiums may range between .8 of a percentage point to about 1.1 percentage points.

Predicting the Recent Low Level of the Long Bond Rate

I now present evidence that the bond equation that allows for the presence of a downward shift in the Fisher coefficient as in equation (6) is consistent with the actual behavior of the long bond rate in recent years. In particular, I estimate the bond equation (6) over 1984Q1 to 2004Q3 and simulate it dynamically over 1994Q1 to 2004Q3. Figure 7 charts the simulated values generated using actual values of right-hand-side explanatory variables. Actual values of the bond rate and the forecast errors are also charted there. This figure suggests two observations. First, this equation predicts reasonably well the

Figure 8 Foreign Official Purchase of U.S. Treasury Securities



actual path of the bond rate over 1994Q1 to 2004Q3. The mean prediction error is small and equals .20, and the root mean squared error is one-half of a percentage point. Second, during the past two-and-a-half years, the ten-year bond rate has hovered around 4 percent, and this behavior of the bond rate seems consistent with economic fundamentals, once we allow for a break in the Fisher coefficient.

Robustness: Assessing the Potential Role of Increased Foreign Purchases of U.S. Treasury Securities

As indicated at the beginning, another popular explanation of the current low level of long-term interest rates is that increased purchases of U.S. Treasury securities by foreign individuals and foreign central banks may have contributed to the recent declines in long bond rates. Figure 8 charts foreign official net purchases of U.S. securities (summed over four quarters) as a percentage of lagged U.S. GDP, and this chart clearly indicates a significant increase in foreign official net purchases during the past few years.

One preliminary test of the above-noted explanation is to augment the baseline bond equation (4) to include the level and/or change in foreign official

Table 2 Estimates of the Bond Rate Equation, Including Foreign Official Holdings of U.S. Treasury Securities

	Dependent Variable: $\triangle BR_t$ Sample Period Ending in			
Independent Variables	(4.1) 2000Q4	(4.2) 2004Q3	(5.1) 2000Q4	(5.2) 2004Q3
$const.$				
BR_{t-1}	-.19 (2.6)	-.22 (3.4)	-.25 (3.3)	-.27 (3.9)
π^{10}_{t-1}	.29 (2.5)	.34 (3.1)	.22 (1.9)	.28 (2.5)
RP_{t-1}			.10 (2.6)	.09 (2.4)
rk_{t-1}	.03 (0.2)	-.01 (0.1)	.10 (0.8)	.02 (0.2)
Δrk_t	-.20 (1.0)	-.28 (1.6)*	-.18 (0.9)	-.30 (1.8)*
$\Delta\pi^{10}_t$	.30 (2.0)	.33 (2.2)	.24 (1.7)	.28 (1.9)
Δy^e_{t+s}	.20 (1.5)	.14 (1.3)	.20 (1.6)	.15 (1.4)
ΔP^e_{t+s}	.42 (2.0)	.37 (1.9)	.41 (2.1)	.38 (2.1)
$u\Delta FFR_t$	.26 (2.2)	.17 (1.6)	.24 (2.1)	.18 (1.7)
a_π	1.53	1.50	.90	1.0
a_r			.41	.35
R^2	.25	.21	.31	.26
SER	.533	.535	.508	.517

Notes: rk is foreign official holdings of U.S. Treasury securities, expressed as a proportion of lagged GDP; other variables are defined as in Table 1. See notes in Table 1. * significant at the .10 level.

purchases and examine whether the long bond rate is negatively correlated with foreign official inflows over 1984Q1 to 2004Q3. In order to determine whether results are due to the most recent large foreign inflows, I also estimate the bond equation over a shorter sample period, 1984Q1 to 2000Q4.

Table 2 presents estimates of the augmented baseline bond equations over two sample periods. The columns labeled (4.1) and (4.2) present estimates of the baseline equation augmented to include foreign official capital inflows, whereas the columns labeled (5.1) and (5.2) contain estimates of the baseline equation augmented to include both foreign inflows and the empirical measure of inflation uncertainty. If we focus on estimates from the baseline equation with foreign inflows over the shorter sample period 1984Q1 to 2000Q4, they suggest the long bond rate is not significantly correlated with foreign official inflows. The estimated coefficients that appear on empirical measures of foreign inflows are not statistically different from zero (the p-value for the null hypothesis—coefficients on the level and change in foreign official inflows are zero—is .45, which is large and leads to the acceptance of the null hypothesis). The result that the long bond rate is not correlated with foreign capital inflows continues to hold if we augment the baseline equation to include both capital inflows and the empirical measure of inflation uncertainty. As can be seen, the estimated coefficient on foreign official inflows remains statistically insignificant, whereas the estimated coefficient on inflation uncertainty is correctly signed and statistically significant (compare coefficients across columns labeled [4.1] and [5.1] in Table 2).

If we consider estimates of the augmented baseline equations over the full sample period that spans the recent period of large foreign inflows, the results are mixed. The estimated coefficient that appears on the level of foreign inflows is still not statistically different from zero. However, the estimated coefficient that appears on the variable measuring change in foreign inflows turns negative and is marginally significant, suggesting part of the decline observed in the long bond rate in recent years may be due to increased foreign purchases (see the coefficient on foreign capital inflows in columns labeled [4.2] and [5.2] in Table 2). But these results also imply that negative correlation between changes in the long bond rate and changes in foreign official purchases found in the full sample period are mainly attributed to the most recent period and hence are not indicative of the presence of a consistent relation between bond yields and increased foreign purchases of U.S. Treasury securities. Thus, the hypothesis that the current low level of the long bond rate is in part due to increased foreign official purchases of U.S. Treasury securities must be considered tentative.[26]

4. CONCLUDING OBSERVATIONS

One suggested explanation of the current low level of the long bond rate is that inflation risk premiums have declined. This explanation posits that, as a result of the good inflation performance of the U.S. economy and increased confidence that the Federal Reserve will keep inflation low and stable, investors are now demanding lower inflation risk premiums than before. This lowering of inflation risk premiums is reflected in lower real and nominal yields on bonds. This article develops an empirical test of the aforementioned explanation.

Since we do not have a direct empirical measure of uncertainty about long-term inflation forecasts, the article develops an empirical proxy for uncertainty about short-term inflation forecasts, assuming uncertainty about long-term inflation forecasts is positively correlated with uncertainty about short-term ones. Another assumption is that if inflation had been harder to forecast in the past, it would raise the variance of current forecasts of expected future

[26] The evidence in previous research on the role of foreign official purchases of U.S. Treasury securities in explaining the current low level of the long bond rate is also mixed. Wu (2005) reports evidence indicating the long bond rate is not at all correlated with foreign official purchases. Warnock and Warnock (2005) report mixed evidence; they also find the estimated coefficient on the foreign official capital inflows in their reduced-form interest rate equation is not statistically different from zero over the estimation period that excludes the surge in inflows of the past few years.

inflation rates, leading to increased uncertainty about future expected inflation. Given these basic assumptions, the article examines the MSE of short-to-medium-term inflation forecasts, using survey data on private-sector GDP inflation expectations. In particular, the article creates a time series on the MSE of short-term inflation forecasts, using rolling three-year windows over 1984Q1 to 2004Q3. This time series can be viewed as measuring uncertainty about short-term inflation forecasts and hence may provide information on uncertainty about long-term inflation forecasts. The time series measuring uncertainty about short-term inflation forecasts has a downward trend that appears to be consistent with the downward trend in mean and variance of forecast inflation, suggesting inflation uncertainty declined over this period because inflation both steadily declined and became more predictable.

The results indicate the long bond rate is positively correlated with the empirical measure of short-term inflation uncertainty over the full sample period 1984Q1 to 2004Q3, which suggests that an increase in uncertainty about short-to-medium-term inflation forecasts raises uncertainty about long-term inflation forecasts and hence may account for the presence of the inflation risk premium in the bond rate. However, the results also indicate that the estimated coefficient that measures the response of the long bond rate to short-term inflation uncertainty has declined since 2001Q4, implying that an increase in uncertainty about short-term inflation forecasts has not raised uncertainty about long-term inflation forecasts as much as it did previously. In fact, the results are consistent with the hypothesis that the inflation risk premium embedded in the long bond rate has disappeared, thereby explaining the current low level of the long bond rate.

Another competing explanation of the recent low level of the long bond rate is increased purchases of U.S. Treasury securities by foreign central banks, which may have contributed to reducing nominal yields on long-term bonds. The empirical work here indicates the long bond rate is in fact negatively correlated with foreign capital inflows over the full sample period. However, this negative correlation between the long bond rate and foreign official inflows found in the data is marginally significant and fragile, arising mainly as a result of most recent capital inflows and hence may not be indicative of the presence of a consistent relation between foreign capital inflows and bond yields. Hence, the second hypothesis must be considered tentative. Together, these results by far favor the explanation that attributes the recent low level of the long bond rate mostly to the reduction in inflation uncertainty.

REFERENCES

Ang, Andrew, Geert Bekaert, and Min Wei. 2006. "Do Macro Variables, Asset Markets, or Surveys Forecast Inflation Better?" Finance and Economic Discussion Series, Federal Reserve Board, Washington, D.C.

Bernanke, Ben S. 2003. "A Perspective on Inflation Targeting." Remarks at the Annual Washington Policy Conference of the National Association of Business Economists, Washington, D.C.

__________. 2004. "Fedspeak." Remarks at the Meetings of the American Economic Association, San Diego, California.

__________. 2006. "Testimony of Chairman Ben S. Bernanke." Semiannual Monetary Policy Report to the Congress.

Clouse, Jim. 2004. "Reading the Minds of Investors: An Empirical Term Structure Model for Policy Analysis." Finance and Economic Discussion Series, Federal Reserve Board, Washington, D.C.

Diebold, Francis X., Monika Piazzesi, and Glen D. Rudebusch. 2005. "Modeling Bond Yields in Finance and Macroeconomics." *American Economic Review* 95 (2): 415–20.

Dudley, Bill. 2006. "Low Bond Risk Premia: The Collapse of Inflation Volatility." *U.S. Economics Analyst*, U.S. Economic Research Group, Goldman Sachs (January).

Ellingsen, Tore, and Ulf Soderstrom. 2001. "Monetary Policy and Market Interest Rates." *American Economic Review* 91 (5): 1594–1607.

__________. 2004. "Why Are Long Rates Sensitive to Monetary Policy?" Sveriges Riksbank Working Paper 160.

Feenberg, Daniel, and Elisabeth Coutts. 1993. "An Introduction to the TAXSIM Model." *Journal of Policy Analysis and Management* 12 (1): 189–94.

Greenspan, Alan. 2005. "Remarks by Chairman Alan Greenspan." Central Bank Panel Discussion, International Monetary Conference, Beijing, People's Republic of China. June 6.

Gurkaynak, Refet S., Brian Sack, and Eric T. Swanson. 2005. "Do Actions Speak Louder than Words? The Response of Asset Prices to Monetary Policy Actions and Statements." *International Journal of Central Banking* (May): 55–93.

Hordahl, Peter, Oreste Tristani, and David Vestin. 2006. "A Joint Econometric Model of Macroeconomic and Term-Structure Dynamics." *Journal of Econometrics* 131 (1–2): 405–44.

Kim, Don H., and Jonathan H. Wright. 2005. “An Arbitrage-Free Three-Factor Term Structure Model and the Recent Behavior of Long-Term Yields and Distant-Horizon Forward Rates.” Finance and Economics Discussion Series Paper No. 33.

Kuttner, Kenneth N. 2001. “Monetary Policy Surprises and Interest Rates: Evidence from the Fed Funds Futures Market.” *Journal of Monetary Economics* 47: 523–44.

Mehra, Yash. 1984. “The Tax Effect, and the Recent Behavior of the After-tax Real Rate: Is it Too High?” Federal Reserve Bank of Richmond *Economic Review* 70 (4): 8-20.

______________. 1994. “An Error-Correction Model of the Long-term Bond Rate.” Federal Reserve Bank of Richmond *Economic Quarterly* 80 (4): 49–67.

Romer, Christina and David Romer. 2004. “A New Measure of Monetary Shocks: Derivation and Implications.” A paper presented at Economic Fluctuations and Growth Research Meeting, NBER, July 17.

Tanzi, Vito. 1980. “Inflationary Expectations, Economic Activity, Taxes, and Interest Rates.” *American Economic Review* (March): 12–21.

Tulip, Peter. 2005. “Has Output Become More Predictable? Changes in Greenbook Forecast Accuracy.” Finance and Economic Discussion Series, Federal Reserve Board, Washington, D.C.

Warnock, Francis E., and Veronica Cacdac Warnock. 2005. “International Capital Flows and U.S. Interest Rates.” Finance and Economic Discussion Series Federal Reserve Board, Washington, D.C.

Wu, Tao. 2005. “The Long-Term Interest Rate Conundrum: Not Unraveled Yet?” Federal Reserve Bank of San Francisco *Economic Letter* (April).