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On the Need for a New Approach to Analyzing Monetary Policy*

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ABSTRACT

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The main instrument of monetary policy is the short term nominal interest rate. Standard analyses of monetary policy focus on modeling the link between this interest rate and the rest of the macroeconomy. Here we document two regularities about the behavior of short and long term interest rates and the macroeconomy that point to key failings of current approaches to modeling monetary policy. We argue that to account for these regularities one must address central questions about the links between monetary policy and the economy that are not addressed successfully in standard monetary models. We are led to call for a new approach to analyzing monetary policy.

Our first regularity concerns the changing behavior of long term and short term nominal interest rates, referred to as *long* and *short rates*, observed in the long run historical data. Most economists are familiar with the observation that long and short rates move closely together, at least at low frequencies, for much of the post World War II period. What we find striking is that this pattern is found only in this post war data. In the pre-World War II data, in the United States and in many of the countries in Europe, long rates were smooth even when short rates were highly volatile. In addition, at least in the United States, we find evidence that after 1990 long rates and short rates seem to be returning to the prewar pattern.

Standard theory based on the expectations hypothesis links movements in long rates to movements in agents' expectations of averages of future short rates over the long term. We argue, on the basis of the recent work of Cochrane and Piazzesi (2005) that, at least at low frequencies, the expectations hypothesis is a useful characterization of the data: movements in long rates do correspond mainly to movements in agents' expectations of averages of future short rates over the long term.

In light of this theory and these empirical results, we see two regimes in these low frequency data on short and long rates: a *stable regime* in which agents have stable long term expectations of short rates and a *volatile regime* in which agents have volatile long term expectations of short rates. The stable regime was in place in many countries for over 100 years prior to World War II. During this regime, it appears that agents' expectations of the average level of the short rate over the long term were firmly anchored. After World War II these countries switched to the volatile regime as early as the 1950s. During this period, it appears that agents' long run expectations became unanchored — it appears that agents interpreted most of the higher frequency movements in the short rate as nearly permanent changes in policy. Finally, data post 1990 suggests that at least the United States has switched back to a stable regime — perhaps agents' long run expectations may be anchored again.

These data on the changing behavior of long and short rates then raise a central question about monetary policy: What change in policy and institutions led to these changes in regimes? We argue that existing analyses of the post World War II data cannot account for the behavior of long term interest rates observed during that time period. Because these existing models fail to account for the volatility of agents' long run expectations observed over this time period, we are skeptical of the usefulness of these models for analyzing this central question about the design of monetary policy and monetary institutions. A new approach is needed.

Our second regularity concerns the comovements in the short rate and the macroeconomy at business cycle frequencies. Empirical work in finance indicates that at business cycle frequencies it is risk that moves when the short rate moves. This result is derived from regressions that show that the risk premium on long term bonds moves more than one-for-one with the *yield spread* between the long and the short rate.

Is this empirical work in finance a problem for standard analyses of monetary policy at business cycle frequencies? We argue that the answer is yes. Standard monetary models rule out movements in risk premia by assumption. Instead, in these models movements in the short rate, by assumption, correspond through the consumption Euler equation to movements in the expected growth in the marginal utility of consumption and expected inflation. This assumption is a serious failing of these models. In the data, movements in the expected growth in the marginal utility of consumption and expected inflation as computed from estimates of a standard model account for virtually none of the movements in the short rate observed at business cycle frequencies. We look to the empirical work in finance on time varying risk as offering the best answer to the question of what in the macroeconomy moves when the short rate moves? It is risk.

These data on the comovement of risk and the short rate at business cycle frequencies then raise a second central question about monetary policy and the macroeconomy: Is the Fed changing the short rate to passively accommodate changes in risk? Or does the Fed change risk when it changes the short rate? Until we can answer this question about the direction of causation, we cannot address key counterfactual questions about what would have happened to the macroeconomy if policy had been different. Since standard monetary models rule out time varying risk by assumption, they cannot be used to address these counterfactual questions. A new approach is needed.

We present our two regularities in the next section. In section 3, we interpret our first regularity in light of standard theory and argue empirically that movements in the long rate correspond mainly to movements in agents' expectations of long run averages of the future short rate, at least at low frequencies. In section 4, we interpret our second regularity in light of standard theory and review the failings of standard monetary models to account for movements in the short rate at business cycle frequencies. In section 5 we discuss our two regularities in light of the current frontier analyses of monetary policy. In section 6 we conclude.

1. Long and Short Term Rates

In this section we document the two regularities in the data on long and short term interest rates that we discuss in this paper. We consider a decomposition of the short rate i_t into the sum of a long rate y_t^L and the yield differential $i_t - y_t^L$ as in

$$(1) \qquad i_t = y_t^L + (i_t - y_t^L)$$

Our first empirical regularity concerns how the relative importance of fluctuations in the long rate and fluctuations in the yield spread in accounting for fluctuations in the short rate have changed over time. We document this regularity in three time periods examining the data post World-War II, pre World-War II, and post 1990. The second regularity concerns how risk moves with the term spread.

Fact 1A. After World War II, long rates are volatile relative to short rates

In Figure 1 we present our decomposition (1) of the short rate into the two components of the long rate and the yield spread for the post - World War II period in the United States. For the short rate we use the 3-month U.S. Tbill yield and for the long rate we use the yield on a 13-year zero coupon U.S. Treasury bond. For the period 1946:12 to 1991:2 we use the data from McCulloch and Kwon (1991) for both series. For the period 1991:3 to 2007:12 we use CRISP for the 3-month Tbill and Gurkaynak, Sack, and Wright (2006) for the 13-year zero coupon bond. In this figure we see that the vast bulk of the movements in the short rate correspond to movements in the long rate and only a small portion of the movements in the short rate correspond to movements in the yield spread. Using data from 1947 to 2007 we compute a variance decomposition which confirms that fluctuations in the long rate account for the bulk of the fluctuations in the short rate.

$$
\begin{aligned}
1 \quad = \quad & \frac{var(y_t^L)}{var(i_t)} + \frac{var(i_t - y_t^L)}{var(i_t)} + 2\frac{cov(y_t^L, i_t - y_t^L)}{var(i_t)} \\
& \quad .92 \qquad\quad + .20 \qquad\qquad - .12
\end{aligned}
$$

We note that our decomposition in Figure 1 is not that sensitive to the choice of the long rate or the choice of the short rate. We chose to use the yield on a 13 year zero coupon bond as our measure of the long rate because that is the longest yield available for the whole postwar sample in our data set. In Figure 2 we plot the yields on 13 year, 23 year, and 30 year zero coupon bonds for the dates at which these series are available[1]. We see that yields on these different bonds are remarkably similar where they overlap.

Following the suggestions in Piazzesi (2003) pages 41-43, we have chosen the 3 month Tbill rate as our measure of the short rate. In Figure 3 we plot the 3 month Tbill rate together with the Federal Funds target rate, which starts only in 1982. We see that these two series track each other closely over the period where they overlap.

Fact 1B. Before World War II, long rates are not volatile relative to short rates.

In Figure 4A we graph a short rate and a long rate for the United States from 1835 through the present. For the short rate we use the US 3-month Commercial Paper rate and for the long rate we use the yield on a 10 year U.S. Treasury bond (available at www.globalfinancialdata.com.). Clearly, in the pre-World War II period, fluctuations in the long rate are a much smaller component of overall fluctuations in the short rate than they are in the post-World War II period. In particular, in the data from 1835 to 1939 our variance

[1] Up to 1991:2 these series are McCulloch and Kwon (1991) and from 1991:3 they are from Gurkaynak, Sack, and Wright (2006).

decomposition is

$$\begin{array}{ccccc} 1 & = & \frac{var(y_t^L)}{var(i_t)} + \frac{var(i_t - y_t^L)}{var(i_t)} + 2\frac{cov(y_t^L, i_t - y_t^L)}{var(i_t)} \\ & & .08 \quad + .79 \quad + .14 \end{array}$$

while in the data from 1945 to 2007 we find

$$\begin{array}{ccccc} 1 & = & \frac{var(y_t^L)}{var(i_t)} + \frac{var(i_t - y_t^L)}{var(i_t)} + 2\frac{cov(y_t^L, i_t - y_t^L)}{var(i_t)} \\ & & .85 \quad + .18 \quad - .03 \end{array}$$

This difference in pre- and post-war behavior of long and short rates is also evident in the data for of a large number of countries, including the United Kingdom (Figure 4B), France (Figure 4C), Germany (Figure 4D), and the Netherlands (Figure 4E).

Fact 1C. Post 1990, the relative volatility of long and short interest rates has declined.

In Figure 1 we can clearly see two episodes since 1990 in which the Fed has engineered a persistent dip in short rates without corresponding fluctuations in long rates. (The Fed may well be starting a third persistent dip in short rates starting in early 2008.) This apparent change in the relative volatility of long and short rates can be seen in our decomposition if we split the data in Figure 1 into two subsamples. From 1946:12 to 1989:12 our decomposition yields

$$\begin{array}{ccccc} 1 & = & \frac{var(y_t^L)}{var(i_t)} + \frac{var(i_t - y_t^L)}{var(i_t)} + 2\frac{cov(y_t^L, i_t - y_t^L)}{var(i_t)} \\ & & 0.98 \quad + 0.12 \quad - 0.11 \end{array}$$

while from 1990:1 to 2007:12 it yields

$$\begin{array}{ccccc} 1 & = & \frac{var(y_t^L)}{var(i_t)} + \frac{var(i_t - y_t^L)}{var(i_t)} + 2\frac{cov(y_t^L, i_t - y_t^L)}{var(i_t)} \\ & & 0.55 \quad + 0.68 \quad - 0.23. \end{array}$$

Here we see that relative variances of the long and the short has decreased from nearly 1 (.98) in the first part of the sample to only about 1/2 (.55) in the later part of the sample. While this change in the relative volatility of the long and short rates in the data post 1990 is not as obvious as it is in the pre World-War II data, it is suggestive that the behavior of short and long rates has changed in this more recent period.

We are not the only researchers who have noticed this recent reduction in the volatility of long rates. Indeed, Chairman Greenspan has referred to the stability of the long rate in the presence of a variable short rate during this period a conundrum. Moreover, Rudebusch, Swanson, and Wu (2006) also mention the reduction in the volatility of long rates in recent data.

We summarize these observations as

Fact 1. Pre World-War II, long rates are stable relative to short rates in many countries. Post World-War II, long rates are volatile relative to short rates in those same countries. After 1990, in the United States, long rates are stable again relative to short rates.

We now turn to our second fact.

Fact 2. The risk premium on long bonds moves more than one for one with the yield spread.

Our second fact concerns the comovement of interest rates and risk. There is a large empirical literature in finance that finds that movements in the spread between long and short yields forecasts movements in the returns earned by investors who buy a long bond and hold it for one year relative to the return that investors would earn at the risk free interest rate for that year. We call this difference in returns the *excess return* on long bonds and the expected value of this excess return the *risk premium* on long bonds. What we emphasize here is the empirical magnitude of this relationship between the yield spread and the risk premium on long term bonds — the risk premium on long bonds moves more than one for one with the yield spread.

We use the following notation to describe these empirical results more precisely. Let P_t^k denote the price in period t of a zero-coupon bond that pays off one dollar in period $t+k$ and let $p_t^k = \log P_t^k$. Then the return to holding this k period bond for one period is $r_{t+1}^k = p_{t+1}^{k-1} - p_t^k$ and the (log) excess return to holding this bond over the short bond is $r_{xt+1}^k = r_{t+1}^k - i_t$. The *risk premium* on long bonds is the expected excess return $E_t r_{xt+1}^k$. Many authors have run return forecasting regressions of excess returns against the yield spread similar to the regression

$$(2) \qquad r^k_{xt+1} = \alpha^k + \beta^k (y^L_t - i_t) + \varepsilon^k_{t+1}$$

Note that under the hypothesis that the risk premia on long bonds are constant over time, the slope coefficient β^k in this regression should be zero. In the data, however, these regressions yield estimates of β^k that are significantly different from zero with point estimates typically greater than 1 for moderate to large k.

We emphasize the magnitude of this slope coefficient here because these regression results thus imply that the risk premium on long bonds moves more than one for one with the yield spread. More precisely, note that a finding that the slope coefficient $\beta^k \geq 1$ implies that

$$(3) \qquad Cov(E_t r^k_{xt+1}, y^L_t - i_t) \geq Var(y^L_t - i_t)$$

which, using simple algebra, implies that the variance in the risk premium on long bonds is greater than the risk premium

$$(4) \qquad Var(E_t r^k_{xt+1}) \geq Var(y^L_t - i_t).$$

We illustrate this finding using the same data on bond yields from McCulloch and Kwon (1991) and Gurkaynak, Sack, and Wright (2001) cited in Figure 1. Following the literature, we set the holding period to be a year. One can run this return forecasting regression for long bonds of any maturity. For simplicity, for the left-hand side variable we use $\bar{r}_{xt+1} = \sum_{k=2}^{13} r^k_{xt+1}/12$, that is, the average one-year holding period excess returns on bonds of maturity 2 through 13, and for y^L_t we use the yield on the 13 year bond for i_t we use the yield on one-year zero coupon bonds. We estimate a slope coefficient $\beta = 1.97$ in this regression.

Regressions of the form of (2) have been run for 20 years, starting with the work of Fama and Bliss (1987). (See also, Campbell and Shiller 1991, and Cochrane and Piazzesi 2005.) During this period there has been extensive discussion of how to compute standard errors in this context. We choose not to report these errors here and refer the interested reader to Cochrane and Piazzesi (2005) for a discussion of the issues involved in evaluating the significance of these return forecasting regressions.

One reason that we find this second fact linking yield spread fluctuations and risk interesting for monetary policy analysis at business cycle frequencies is that the yield spread roughly captures the movements in the monetary policy instrument, namely the short rate, at those frequencies. In Figure 5 we show this by plotting the yield spread, as defined in Figure 1, together with the HP-filtered short rate.

2. Modeling Long and Short Rates

In this section we consider the implications of standard monetary models for the joint behavior of long and short term interest rates. Standard monetary models assume that all risk premia on long term bonds are constant. As a result of this assumption, they imply that all movements in the long rate correspond to movements in agents' expectations of the average of future short rates over the maturity of the long term bond. We review the finding that this class of models cannot reproduce the data on the volatility of long yields relative to short yields presented in Figures 1 and 2 if one maintains the assumption that the short rate is stationary and ergodic. We argue that to reproduce the post World War II data on the volatility on long rates relative to short rates in this class of models, one must assume that most of the movements in short rates are expected to be extremely persistent.

We then consider whether one might account for the volatility of long rates documented in Figure 1 as arising from large and persistent movements in term premia in long rates rather than from large and persistent movements in agents' expectations of the average of future short term rates over the maturity of the long term bond. We argue that the empirical results on bond risk premia in Cochrane and Piazzesi (2005) indicate this hypothesis is unlikely to hold. We conclude then that movements in long term yields in the post World War II period, at least at low frequencies, are most likely accounted for by large and persistent movements in agents' expectations of short term interest rates.

To define terms, we write the yield on a k period bond in two components

$$y_t^k = \left[\frac{1}{k} E_t \sum_{s=0}^{k-1} i_{t+s}\right] + \left[y_t^k - \frac{1}{k} E_t \sum_{s=0}^{k-1} i_{t+s}\right] \tag{5}$$

given by the two terms in brackets. The first term is the average of the short term interest rates expected over the next k periods. The second term is the difference between yield on the riskless strategy of investing in the k period bond and holding it to maturity and the expected yield on the risky strategy of rolling over a sequence of one period investments in the short term bond between period t and period $t+k-1$. We refer to the second component as the *term premium* on the long yield.

In order to develop a theory of the term premium, it is useful to recast it in terms of expectations of excess returns on a sequence of one-period investments in long-term bonds. Using the definition that $y_t^k = -p_t^k/k$, we have that the term premium can be written

$$\left[y_t^k - \frac{1}{k}E_t\sum_{s=0}^{k-1} i_{t+s}\right] = \frac{1}{k}E_t\left[p_{t+1}^{k-1} - p_t^k - i_t\right] + \frac{1}{k}E_t\left[p_{t+2}^{k-2} - p_{t+1}^{k-1} - i_{t+1}\right] + \ldots +$$

$$\frac{1}{k}E_t\left[p_{t+k-1}^1 - p_{t+k-2}^2 - i_{t+k-2}\right] + \frac{1}{k}E_t\left[0 - p_{t+k-1}^1 - i_{t+k-1}\right]$$

Note that each of the terms in square brackets on the right hand side of this equation the excess return on an investment in a long term bond. Having rewritten the term premium in this way, and noting that $r_{xt+1}^1 = 0$, we then have that long-term yields are equal to agents' expectation of the average of the short-term rate over the life of the long term bond plus their expectation of the excess returns to holding a sequence of long term bonds for one period

$$(6) \qquad y_t^k = \frac{1}{k}\left[E_t\sum_{s=0}^{k-1} i_{t+s} + E_t\sum_{s=0}^{k-1} r_{xt+s+1}^{k-s}\right].$$

Standard monetary models start with the assumption that all risk premia, and hence, all expected excess returns $E_t r_{xt+1}^k$, are constant. As a result, these models imply that term premia on long term bonds are constant and, thus, movements in long term yields correspond exactly to movements in agents' expectations of the average of the short rate over the life of the long-term bond.

The simplest versions of these models also assume that the short rate is stationary and ergodic. As we show next in Proposition 1, as a consequence, these simple models cannot account for the large volatility of long rates relative to that of short rates observed in Figure 1.

Proposition 1. If the short rate i_t is stationary and ergodic and and expected excess returns on long term bonds $E_t r_{xt+1}^k$ are constant, then the long term yield y_t^k converges to a constant $\bar{y}^\infty$ independent of the date t as their maturity k grows to infinity.

Proof. The yield on a k-period bond at t can be written as (6). From our assumption of ergodicity, as k grows to infinity, the long-run average of future short rates expected at date t converges to a constant independent of the date t. *Q.E.D.*

As this proposition makes clear, the prediction that long term yields should not be volatile follows from a basic assumption that underlies a great many macroeconomic models. Thus, the failure of standard monetary models to capture the volatility of long term yields does not lie in the modeling details.[2]

Our proposition is based on an argument about long yields as their maturity gets large. One might ask whether this result is relevant for the data on the long but finite yields that we do observe. A long literature argues that it is. See, for example, Shiller (1979), Singleton (1980), Leroy and Porter (1981), Backus and Zin (1993), and Fuhrer (1996) among many others.

The key tension highlighted in this literature is that our basic assumptions about the time series properties of the short rate chosen by the Fed are in conflict with the data on the volatility of long rates that we observe. Standard models that assume that the short rate is stationary and ergodic typically have the implication that averages of expectations of future short rates converge to a constant too fast to account for the observed variability of long term rates. This difficulty that standard models have in generating movements in long term yet finite yields is easy to illustrate by way of an example. Assume that the short rate is an AR1 of the form

$$i_{t+1} = (1-\rho)\bar{\imath} + \rho i_t + \varepsilon_{t+1}.$$

and that expected excess returns are constant over time. Using (6) we have that

$$y_t^k - \bar{y} = \frac{1}{k}\frac{(1-\rho^{k+1})}{1-\rho}(i_t - \bar{\imath}) \tag{7}$$

[2] Those readers immersed in the literature on bond pricing may note that almost all models of bond pricing, including those models that allow for a non-stationary short-term interest rate, have the implication that the yield on long-term bonds, in the limit, is constant. This implication follows from an argument that if the limiting yield, y_t^∞, exists, then it cannot fall or else there would be an opportunity for arbitrage. See Dybvig, Ingersoll, and Ross (1996) for details. Clearly, in the data, we have seen yields on 30 year bonds both rise and fall so the theoretical arguments about restrictions on the limiting yield y_t^∞ do not apply to the long-term yields that we do observe. Hence, the substantive empirical question regards the implications of our models for long yet finite yields.

This example is presented in Backus and Zin (1993). (Also note that Fama and Bliss (1987) assume that the short rate follows an AR1 in their empirical work.) Using monthly data on the 3-month TBill rate in Figure 1 we estimate ρ to be .986. For a 13 year zero coupon bond in monthly data, $k = 156$, and (7) reduces to

$$y_t^k - \bar{y} = .41(i_t - \bar{\imath}) \tag{8}$$

This model thus implies that the variance of the long rate relative to the short rate is

$$\frac{var(y_t^L)}{var(i_t)} = .17 \tag{9}$$

which is considerably smaller than the corresponding ratio of .98 that we reported earlier for as similar period (1946:12 to 1989:12).

The difficulty that this simple model has in accounting for the volatility of long rates grows as we consider longer term rates. For example, for a 23 year zero coupon bond in monthly data, $k = 276$ and (7) reduces to

$$y_t^k - \bar{y} = .25(i_t - \bar{\imath})$$

while for a 30 year zero coupon bond, $k = 360$ and (7) reduces to

$$y_t^k - \bar{y} = .20(i_t - \bar{\imath})$$

Because of the geometric mean reversion of the short rate, this simple model implies lower and lower volatility of long rates as the maturity grows. As a result, this simple example will also fail to reproduce the observation in Figure 2 that yields of 13, 23, and 30 year maturities are all quite similar.

In the finance literature, empirical researchers looking to model the behavior of both short and long-term yields in the data from the past 50 years have begun to consider models in which the short-term interest rate is non-stationary or has a very persistent, if not quite permanent, component. See, for example, Backus and Zin (1993), Kozicki and Tinsley (2001) and Fama (2006). If the short rate is non-stationary[3], then a variety of time series methods

[3] Technically, there are theoretical problems of assuming that there is literally a nontrivial martingale component to the short rate. (See, for example, Campbell, Lo, McKinley 1997 for details.) Here we are thinking of assuming that the short rate has a random walk component only as a way of approximating a very persistent short rate series.

provide a decomposition of the short rate into a random walk component (or more generally a martingale component) which we denote $\bar{\imath}_t$ and a stationary component $\hat{\imath}_t$, with $i_t = \bar{\imath}_t + \hat{\imath}_t$.

Given such a time series decomposition of the short rate, we then have the following proposition.

Proposition 2. Decompose the short rate i_t into a martingale component $\bar{\imath}_t$ and a stationary component $\hat{\imath}_t$. Assume that the stationary component $\hat{\imath}_t$ is ergodic and that expected excess returns on all long term bonds $E_t r^k_{xt+1}$ for $k \geq 1$ are constant. Then the yield on long term bonds converges to $\bar{\imath}_t$ plus a constant independent of t as their maturity k grows to infinity.

Proof. Given (6), the yield on a k-period bond at t can be written

$$(10) \quad y^k_t = \frac{1}{k}\left[\sum_{s=0}^{k-1} E_t\bar{\imath}_{t+s} + E_t\hat{\imath}_{t+s} + E_t\sum_{s=0}^{k-1} r^{k-s}_{xt+s+1}\right].$$

The result follows from the fact that $E_t\bar{\imath}_{t+s} = \bar{\imath}_t$ and the ergodicity of $\hat{\imath}_t$. *Q.E.D.*

Given this proposition, our data on long-term yields in figure 1 thus suggest that the martingale component of the short rate accounts for most of the total fluctuations in the short rate over the past 50 years.

The long historical data on the volatility of long term yields and the short term interest rate suggest that an improved understanding of how the public forms expectations of the long term course of monetary policy may also yield a resolution of Greenspan's conundrum regarding the stability of long term yields in the face of continued volatility of the short term interest rate. Specifically, if the public has come to perceive movements in short term interest rate over the past ten years or so as stationary and ergodic as opposed to non-stationary, then it would be quite natural for long term yields to be stable in the face of large swings in the short term rate. See the conclusion in Rudebusch, Swanson, and Wu (2006) for a discussion of this hypothesis.

Throughout this discussion we have assumed that the term premium on long term yields is constant and used that assumption to discuss the relation between long yields and expectations of averages of future short rates. We now consider the alternative hypothesis that the large and persistent movements in long yields that we observe in Figure 1 might be accounted for by large and persistent movements in term premia.

Empirical work in finance argues that the expected excess returns to holding long term bonds varies over time and, as a result, from (6) the term premium should also vary over time. To argue that there are large and persistent movements in the term premium, however, one must argue that there are large and persistent movements in agents' expectations of long averages of expected excess returns on long term bonds given in (6) by

$$\frac{1}{k}\sum_{s=0}^{k-1} E_t r_{xt+s+1}^{k-s}$$

We argue that such large and persistent movements in averages of expected excess returns are likely not there in the data.

Our argument is based on the work of Cochrane and Piazzesi (2005), who study the expected excess returns to holding long term bonds. Cochrane and Piazzesi find that annual excess returns on long term bonds of maturity k can all be forecast with a single linear combination of yields at t. Specifically, using data on annual holding period returns, they run regressions of the form

$$r_{xt+1}^{k} = \alpha^k + \beta_1^k y_t^1 + \beta_2^k y_t^2 + \ldots + \beta_5^k y_t^5$$

across maturities k and argue that these regressions are well summarized by a restricted set of regressions of the form

$$r_{xt+1}^{k} = \alpha_k + b_k x_t$$

where they call x_t the *return forecasting factor* defined by

$$x_t = \left[\gamma_1 y_t^1 + \gamma_2 y_t^2 + \ldots + \gamma_5 y_t^5.\right]$$

With this empirical summary of expected excess returns on long-term bonds, then the term premium on a bond of maturity k can be written in terms of this return forecasting factor as

$$\frac{1}{k}E_t\sum_{s=0}^{k-1} r_{xt+s+1}^{k-s} = \text{constant} + \frac{1}{k}E_t\sum_{s=0}^{k-1} b_{k-s}x_{t+s}$$

The question of whether fluctuations in the term premium can account for much of the fluctuations in the yields of long term bonds then comes down to an analysis of whether there are large and persistent movements in these weighted average of expected values of the returns forecasting factor x_t over time. Note that if the series x_t is stationary and ergodic, then we would not expect to see large and persistent swings in this measure of the term premium since these weighted averages of expectations of future values of x_t should converge to a constant independent of the date t.

Cochrane and Piazzesi (2005) plot the time series for the return forecasting factor x_t. One can see in that figure that this return forecasting factor appears to bounce around a relatively constant mean, suggesting the series x_t does not have persistent movements and that averages of the form

$$\frac{1}{k} E_t \sum_{s=0}^{k-1} b_{k-s} x_{t+s}$$

should converge to a constant as k grows.

We now examine this idea more precisely. We use the results in Cochrane and Piazzesi (2005) to construct an estimate of the term premium on the 13 year zero coupon yield as follows. We estimate the autocorrelation of their estimate of x_t in annual data to be $\rho = 0.26$. We use this autocorrelation to estimate the terms

$$(11) \qquad E_t x_{t+s} = \rho^s x_t$$

We use their estimates of the regression slopes b_{k-s} for returns on bonds with maturities $k = 2, 3, 4, 5$ and we use their linear method to extrapolate the regressions coefficients for $k = 6, \ldots, 13$. We then plug the following values of b_k to estimate the term premium on the 13 year bond as

$$\frac{1}{13} E_t \sum_{s=0}^{12} r_{xt+s+1}^{13-s} = \text{constant} + \frac{1}{13} \left[\sum_{s=0}^{11} b_{13-s} \rho^s \right] x_t$$

Regression Coefficients

b_2	b_3	b_4	b_5	b_6	b_7	b_8	b_9	b_{10}	b_{11}	b_{12}	b_{13}
0.47	0.87	1.24	1.43	1.82	2.15	2.47	2.80	3.13	3.45	3.78	4.11

We set the constant here so that our estimate of the average term premium in the sample equals the average difference between the 13 year yield and the one year yield in the sample.

Note that this procedure gives an estimate of the term premium whose fluctuations are simply a scalar times x_t and hence it also does not have persistent movements unless x_t does. In Figure 6 we show the 13 year yield and our estimate of agent's expectations of the average of the short term interest rate over the next 13 years given by subtracting this estimate of the term premium from the observed 13 year yield. While it is clear that the term premium may fluctuate in important ways at higher frequencies, it is also clear from this figure that fluctuations in the term premium do not account for the large and persistent movements in the 13 year yield that we have seen in the post World War II data.

Cochrane and Piazzesi (2008) are developing a much more sophisticated analysis of the dynamics of term premia and we look forward to learning about their improvements on our simple approximation based on (11).

3. Time-Varying Risk

A large body of empirical work in finance finds evidence time-varying risk premia in the expected excess returns on investments in stocks, long term bonds, and foreign currency denominated bonds. This evidence has had a large impact on work in finance as illustrated by the following quote.

> Overall, the new view of finance amounts to a profound change. We have to get used to the fact that most returns and price variation comes from variation in *risk premia*. (Cochrane 2001, p. 451)

Cochrane's observation directs our attention to a critical counterfactual assumption of the standard general equilibrium monetary model: constant risk premia. We now argue that this failure of standard monetary models to capture movements in risk premia is not a detail that can reasonably be abstracted from when analyzing the impact of monetary policy on the economy. Instead, it is a central problem in terms of using standard models to derive implications for the comovements of the short term interest rate and other macroeconomic aggregates. This is because the evidence in finance indicates that movements in risk premia are highly correlated with movements in the short rate at business cycle frequencies and are at least as large as these movements in the short rate itself.

Consider first the link between the short rate and macroeconomic aggregates built into standard monetary models. We begin with representative agent models. The short term nominal interest rate enters standard representative consumer models through an Euler equation of the form

$$(12) \quad \frac{1}{1+\mathrm{i}_t} \equiv \exp(-i_t) = \beta E_t \left[\frac{U_{ct+1}}{U_{ct}} \frac{1}{\pi_{t+1}} \right],$$

where i_t is the logarithm of the short term nominal interest rate $1+\mathrm{i}_t$, β and U_{ct} are the discount factor and the marginal utility of the representative consumer, and π_{t+1} is the inflation rate. Analysts then commonly assume that the data are well-approximated by a conditionally log-normal model so that this Euler equation can be written as

$$(13) \quad i_t = E_t \left[-\log \frac{U_{ct+1}}{U_{ct}} \frac{1}{\pi_{t+1}} \right] - \frac{1}{2} \mathrm{var}_t \left[\log \frac{U_{ct+1}}{U_{ct}} \frac{1}{\pi_{t+1}} \right].$$

The critical question in monetary policy analysis is what terms on the right hand side of (13) change when the monetary authority changes the interest rate i_t. The traditional assumption is that conditional variances are constant, so that the second term in (13) is constant. This leaves the familiar version of the Euler equation:

$$(14) \quad i_t = -E_t \log \frac{U_{ct+1}}{U_{ct}} + E_t \log \pi_{t+1} + \text{constant}.$$

Thus, by assumption, standard monetary models imply that movements in the short term nominal interest rate are associated one-for-one with the sum of the movements in the expected growth of the log of marginal utility for the representative consumer and expected inflation. The debate in the literature on the effects of monetary policy might be summarized roughly as a debate over how much of the movement in the short term interest rate is reflected in the expected growth of the log of marginal utility of consumption (representing a *real* effect of monetary policy) and how much of the movement is reflected in expected log inflation (representing a *nominal* effect of monetary policy). The answer to this question in the context of a specific model depends on the specification of the other equations of the model. However, virtually universally, the possibility that movements in the short term interest rate might be associated with changes in the conditional variances of these variables is ruled out by assumption.

We have described the standard approach in the context of a model with a representative consumer. Our discussion also applies to more general models which do not assume a representative consumer. To see this note that we can write equations (12)-(14) more abstractly in terms of a nominal *pricing kernel* (or stochastic discount factor) m_{t+1} as

$$(15) \quad \exp(-i_t) = E_t m_{t+1}.$$

In a model with a representative agent this pricing kernel is given by $m_{t+1} = \beta U_{ct+1}/(U_{ct}\pi_{t+1})$ and (15) is the representative agent's first order condition for optimal bond holdings. In some segmented market models (15) is first order condition for the subset of agents who actually participate in the bond market while in others (15) is no single agent's first order condition. In general equation (15) is implied by lack of arbitrage possibilities in financial market.

Using conditional log-normality (15) implies

$$(16) \quad i_t = -E_t\left[\log m_{t+1}\right] - \frac{1}{2}var_t\left[\log m_{t+1}\right]$$

and with constant conditional variances we have

$$(17) \quad i_t = -E_t \log m_{t+1} + \text{constant}.$$

Thus the more general assumption made in the literature is that movements in the short term interest rate are associated with movements in the conditional mean of the log of the pricing kernel and not with movements in its conditional variance.

It is clear that standard monetary models with constant conditional variances are inconsistent with the evidence from finance of time-varying risk premia. Is this a serious problem if we want to use these models to understand what in the macroeconomy moves when the short rate moves? We argue that the answer to this question is yes.

To begin to address this question, consider first what aspects of the comovements of the short rate and macroeconomic aggregates that we miss in the Euler equation of standard monetary models. The basic problem with the simplest standard monetary models is that the terms

$$-E_t \log \frac{U_{ct+1}}{U_{ct}} + E_t \log \pi_{t+1}$$

are too smooth relative to the short rate at business cycle frequencies so they account for virtually none of the fluctuations in the policy variable, the short rate, at these frequencies. To illustrate this point, we[4] have estimated a version of the Smets Wouters (2007) model with

[4]Actually, we asked Ellen McGrattan to reestimate the model using codes kindly provided by Smets and Wouters and she kindly obliged. A similar remark applies later to the computations underlying Figures 10 and 11.

standard CRRA preferences and computed the errors in the consumption Euler equation, where the error is computed as

$$\text{error} = i_t - \left[-E_t \log \frac{U_{ct+1}}{U_{ct}} + E_t \log \pi_{t+1} \right].$$

In Figure 7, we plot the HP filtered short term interest rate from the model (the Fed Funds Rate) and the HP filtered error in the Euler equation. We find this figure striking. As we have explained, in theory the standard monetary models imply that movements in the short rate are associated one-for-one with the sum of the movements in the expected growth of the log of marginal utility for the representative consumer and expected inflation. Figure 7 shows that, in practice in a standard monetary model, movements in the short rate are associated almost one-for-one with Euler equation error and the model captures essentially none of the link between the short rate and the macroeconomy. Since this Euler equation is the fundamental link between monetary policy and the macroeconomy, the standard model can hardly be said to be useful for analyzing monetary policy at business cycle frequencies if the observed movements in the monetary policy instrument at these frequencies correspond simply to the unexplained error in this equation.

What is this Euler equation error that moves with the short rate at business cycle frequencies? In previous work (Alvarez, Atkeson, and Kehoe 2007), we used data on US and foreign currency denominated interest rates and exchange rates to argue that these errors correspond to movements in risk premia. More specifically, we argued that what the standard monetary model misses is large fluctuations in the conditional variance of the pricing kernel.

We sketch our previous argument here. Let

$$r^*_{xt+1} = i^*_t + e_{t+1} - e_t - i_t$$

denote the (log) excess return on a foreign short bond with rate i^*_t where e_t is the log of the exchange rate. Regressions of the excess return on an investment in the foreign currency bond on the interest differential of the form

$$r^*_{xt+1} = a + b(i^*_t - i_t) + u_{t+1}. \tag{18}$$

yield estimates of b greater than one. Defining the risk premium on foreign bonds as $E_t r^*_{xt+1}$, note that if risk premia were constant, then this regression should yield a slope coefficient estimate of zero. The finding that these regressions have a slope coefficient greater than one imply the risk premium moves more than one for one with the interest differential in the sense that

$$cov\left(E_t r^*_{t+1}, i^*_t - i_t\right) \geq var\left(i^*_t - i_t\right)$$

which in turn implies that

$$var\left(E_t r^*_{t+1}\right) \geq var\left(i^*_t - i_t\right).$$

This finding has immediate implications for the conditional means and the conditional variances of pricing kernels. To see this observe that lack of arbitrage in complete financial markets implies that

$$(19) \quad e_{t+1} - e_t = \log m^*_{t+1} - \log m_{t+1}.$$

Combining (19) with (16) and its analog for i^*_t gives

$$(20) \quad E_t r^*_{xt+1} = \frac{1}{2}\left[var_t \log m^*_{t+1} - var_t \log m_{t+1}\right].$$

Thus the regression evidence implies that this difference of conditional variances moves more than one for one with the interest differential in the sense that

$$(21) \quad cov\left(\frac{1}{2}\left[var_t(\log m^*_{t+1}) - var_t(\log m_{t+1})\right], i^*_t - i_t\right) \geq var\left(i^*_t - i_t\right).$$

which in turn implies that

$$(22) \quad var\left(\frac{1}{2}\left[var_t(\log m^*_{t+1}) - var_t(\log m_{t+1})\right]\right) \geq var\left(i^*_t - i_t\right).$$

Equations (21) and (22) make precise the sense in which risk premia at least as much as interest differentials when these interest differentials move.

We find the data on risk premia on foreign currency bonds particularly informative about the comovements of the short rate and the pricing kernel because the relation (19) directly links pricing kernels to exchange rates realization by realization. This tight link allows us to identify movements in foreign currency risk premia with movements in conditional variances in the pricing kernel with minimal assumptions.

The data on time-variation in risk premia on long term bonds in Fact 2 also have implications for movements in the conditional variances of the pricing kernel. While this link is suggestive, it is less direct because we have no analog of (19) that links observables to movements in the pricing kernel realization by realization. We explore this link here.

To do so, consider a time series process for the level of the marginal utility of a dollar denoted by $\{Q_t\}$. In a standard representative agent monetary model, this marginal utility corresponds to

$$Q_t = \lambda_0 \beta^t \frac{U_{ct}}{P_t}.$$

where λ_0 is the multiplier on the period 0 budget constraint. This marginal utility of a dollar is directly proportional to the date zero price of a dollar delivered at date t in state s^t (with notation for the state suppressed) divided by the unconditional probability of that state being realized at that date. Note that the pricing kernel for assets that pay off at $t+1$ is

$$m_{t+1} = \frac{Q_{t+1}}{Q_t}.$$

Standard asset pricing formulas then give the price of a k period zero coupon bond at date t as

$$(23) \quad p_t^k = \log E_t Q_{t+k} - \log Q_t.$$

We now examine the relationship between excess holding period returns for bonds of different maturities and the conditional moments of the pricing kernel. While we assume that Q_{t+1} is conditionally lognormal, we do not assume that Q_{t+k} is conditionally log-normal for $k > 1$ conditional on information in period t since we allow for random variances. Instead we assume that the conditional expectation $E_{t+1}Q_{t+k}$ is itself conditionally lognormally distributed given information available in period t (as would be the case in an affine model of the pricing kernel). We prove the following proposition in the appendix.

Proposition 3: The expected log excess holding period return on a k period bond is given by

$$(24) \quad E_t r_{xt+1}^k = \frac{1}{2} var_t \left(\log Q_{t+1} \right) - \frac{1}{2} var_t \left(\log E_{t+1} Q_{t+k} \right).$$

Note that this expression for the risk premium on long term bonds links excess returns to the conditional variance of the pricing kernel in a manner somewhat analogous to that in expression (20)

Expression (24) implies that expected excess returns on long bonds depends on two types of uncertainty. The first, that we refer to as *short run risk*, is uncertainty at time t about how news at time $t+1$ will affect the price of a dollar at $t+1$, namely Q_{t+1}. The second, that we refer to as *long run risk*, is uncertainty at time t of how news at time $t+1$ affects the expected price of a dollar at time $t+k$, namely $E_{t+1}Q_{t+k}$. Intuitively, an increase in short run risk raises the risk premium on long term bonds because this increase raises the uncertainty faced by an investor planning on selling a long term bond at $t+1$. Correspondingly, an increase in long run risk lowers the risk premium on long bonds because these bonds are a hedge against long run risk, so investors require a lower expected return to hold them.

If we assume that short run and long run risk both increase when the yield spread increases, in the sense that

$$cov(\frac{1}{2}var_t(\log Q_{t+1}), y_t^L - i_t) \geq 0$$

and

$$cov(\frac{1}{2}var_t(\log E_t Q_{t+k}), y_t^L - i_t) \geq 0,$$

then we can show that the conditional variance of the pricing kernel moves more than one-for-one with the yield spread. We demonstrate this result in the following proposition.

Proposition 4: Assume that short run and long run risk both increase when the yield spread increases, then the regression evidence from Fact 2 implies that

$$(25) \quad cov(\frac{1}{2}var_t(\log Q_{t+1}), y_t^L - i_t) \geq var(y_t^L - i_t).$$

which in turn implies

$$var_t(\log Q_{t+1}) \geq var(y_t^L - i_t).$$

Proof: To see this result use (24) to write

$$(26) \quad cov(E_t r_{xt+1}^k, y_t^L - i_t) = cov(\frac{1}{2}var_t(\log Q_{t+1}), y_t^L - i_t) - cov(\frac{1}{2}var_t(\log E_t Q_{t+k}), y_t^L - i_t).$$

Under the assumption that short run and long run risk both increase when the yield spread increases both covariances on the right side of (26) are positive. Hence, (3) implies (25). *Q.E.D*

In the appendix we present an example of an affine pricing kernel that satisfies the assumptions of proposition 4. This example also captures the main mechanisms generating time varying risk premia in the segmented markets model of Alvarez, Atkeson, and Kehoe (2007).

To summarize, on the basis of this evidence of time varying risk premia on foreign currency and long term bonds, we argue that the movements in the short rate that are missed by standard monetary models, that is, the Euler equation errors, shown in Figure 7 are likely movements in the conditional variance of the log of the pricing kernel.

4. Implications for Monetary Economics

We have presented two stylized facts — one about the low frequency behavior of interest rates and one about their behavior at business cycle frequencies.

At low frequencies we see two regimes, a *stable regime* in which agents have stable long term expectations of future short rates and a *volatile regime* in which agents have volatile long term expectations of future short rates. The stable regime was in place in many countries for over 100 years prior to World War II. After World War II these countries switched to the volatile regime as early as the 1950s. Finally, data post 1990 suggests that at least the United States has switched back to a stable regime.

At business cycle frequencies in postwar data we see that it is mainly risk that moves when the short rate moves. We argued using both data on international interest differentials and on yield spreads that this movement in risk corresponds to movements in the conditional variance of pricing kernels.

New approaches to analyzing monetary policy are needed to confront these two stylized facts.

We begin with a discussion of the central questions that should shape the long term research agenda in monetary economics. We then discussion some immediate steps researchers should take with their current frameworks.

A. A Long Term Research Agenda

Consider first the low frequency evidence on the changing volatility of agents' long run expectations of future short rates. If we step back from day-to-day analysis of monetary policy, we see that low frequency data raise a fundamental question: After 100 years of stability why did so many countries enter a volatile regime after World War II? The existing literature has not answered this question.

The literature has offered two basic approaches to modeling the volatile regime in postwar U.S. data. The first approach is mechanically describes the aspects of Fed policy over this period that may have made the regime volatile. The second approach looks to explicitly model the Fed's objectives and information that led to its volatile behavior.

We begin by discussing several prominent examples of this first approach. Clarida, Gali, and Gertler (2000) QJE estimate Taylor rules using postwar data and argue that, for at least a portion of this period, the Fed's reaction function led to highly variable inflation. Given that the short rate is stationary and ergodic under the policy regime that leads to a unique equilibrium in this paper, we strongly suspect that this model cannot generate the observed movements in the long rate observed in this period. It is not clear how to compute this model's implications for the long rate under the policy regime that allows for multiple equilibria. Hence it is not evident that this model can confront the evidence of volatility in agents' long term expectations.

Next, Sims and Zha (2005) also estimate a Taylor rule allowing for multiple regimes both for the coefficients in the rule and the processes on the shocks. In Figure 8 we plot their model's implications for the expectations of the average of the short rate over a 13 year and 30 year horizon. (This figure is not yet available).

Finally, a branch of literature in this approach explicitly uses data on long rates to help estimate the Fed's choice for short rates. (See, for example, Ang, Dong, and Piazzesi 2007, Bekaert, Cho, and Moreno 2006, Rudebusch and Wu 2005, and Gallemeyer, Hollifield, and Zin 2005) This work is clearly useful for many questions and represents a step forward in integrating important information in long yields into the study of monetary policy.

None of these examples of the first approach, however, address the fundamental question of why a central bank would choose policy in such a way that agents would have such volatile expectations of its average choice of its short rate over the long run.

We turn now to several prominent examples of the second approach. Orphanides (2002) argues that the Fed's difficulties in interpreting real time economic data in the 1970s played a key role in shaping the Fed's choice of the short rate during that time. It is unclear, however, what mechanism in this framework would lead to persistent movements in agents' expectations of future policy. Thus, we do not see how an explanation of this sort would be able to account for the volatility in agents expectations of the Fed's average choice of short rates in the long run.

Sargent, Williams, and Zha (2005) and Primiceri (2006) represent the most ambitious attempts to reconcile the observed low frequency movements in Fed policy with optimizing behavior by the Fed. In these papers the Fed uses a misspecified model to choose policy and continually revises that model in light of the data. This approach is clearly aimed at fundamental questions in analysis of monetary policy in the post World War II period. Unfortunately, however, data on long rates pose a formidable challenge to models of this type. The basic problem is these models have a very difficult time generating volatile long run expectations simply from learning dynamics. To illustrate this point we graph in Figure 9 the time series for long run averages of expected inflation over horizons of 20 and 30 years from the model of Sargent, Williams, and Zha (2005) together with the data on the 13 year and the 23 year zero coupon yield. (Tao Zha kindly provided us with these long run expectations of inflation from the Sargent, Williams, and Zha model.)

Now consider our second stylized fact. If we are interested in the analysis of monetary policy at business cycle frequencies we need to deal with a fundamental issue: Since most of the movements in the short rate at business cycle are movements in risk, how do we introduce time-varying risk into our monetary models?

We see two basic approaches one might take to introduce such risk. One approach, termed the *exogenous risk* approach, models time-varying risk as an exogenous feature of the real economy that the Fed passively accommodates in setting policy. McCallum (1994) is a prominent early proponent of such an approach. The second approach, termed the *endogenous risk* approach, models the Fed as an active player in generating this time-varying risk. Alvarez, Atkeson, and Kehoe (2002 and 2007) propose such an approach. Clearly, before progress can be made in modelling it is essential to sort out which way the causality runs: from risk to the Fed or from the Fed to risk.

The exogenous risk approach suggests a new view of monetary policy. Under it the Fed must continually adjust the short term nominal interest rate in response to time variation in risk even if the sole objective of the central bank is to maintain a constant level of expected inflation. To illustrate this view, consider a simple, constant-velocity cash-in-advance model in which aggregate consumption follows an exogenously given stochastic process so that monetary policy affects only nominal variables. As in Campbell and Cochrane (1999) or Bansal and Yaron (2004), assume that, corresponding to this exogenously given stochastic process for consumption, the log of the marginal utility of consumption of the representative consumer follows an exogenously given stochastic process with a time-varying conditional mean and variance. In such a model the Euler equation (13) clearly implies that the Fed must adjust the short rate i_t to accommodate movements in conditional variances if it is to keep the conditional means of inflation and the growth of the marginal utility of consumption constant.

Under this interpretation the Euler equation errors shown in Figure 7 are simply the result of the Fed adjusting the short rate to accommodate exogenous shocks to risk that would otherwise cause fluctuations in either consumption growth or inflation. Put simply, the Euler equation implies that the volatility in risk has to show up somewhere: either in the short rate or in the expected growth of consumption and inflation. Under this view, the outcome that we observe is one in which the Fed soaks up this volatility in risk in the short rate so as to avoid having it show up in consumption and inflation.

The endogenous risk approach also suggests a new view of monetary policy. Under it we see a weak link at business cycle frequencies between fluctuations in the short rate and expected consumption growth (of aggregate consumption) and inflation. That model has heterogeneous agents with some in and some out of the asset market. Hence, in that model there is no Euler equation of the form (13) linking aggregate consumption and inflation to interest rates. (See Alvarez, Atkeson, and Kehoe 2002 and 2007 for details.) Instead, the pricing kernel is formed from the marginal utilities of the marginal investor at each date and state of nature. Thus, errors in a representative agent version of the Euler equation correspond to gaps between aggregate consumption and the marginal utility of consumption of the marginal investor.

Either of these approaches will lead to a quantitatively nontrivial reassessment of Taylor rules as descriptions of policy. (See, for example, Ang, Dong, and Piazzesi 2007.)

Since McCallum (1994) a large literature have studied pure exchange economies with exogenous time-varying risk. Examples include Wachter (2006) who uses habit persistence, Bansal and Shaliastovich (2007), Gallmeyer, Hollifield, Palomino, and Zin (2007) who use Epstein-Zin preferences with time-varying variances of consumption. Alvarez, Atkeson, and Kehoe (2002 and 2007) study a pure exchange economy with endogenous time-varying risk arising from endogenously segmented asset markets. Incorporating production in these models has proved challenging, if not daunting. (See, for example, Boldrin, Christiano, and Fisher 2001 and Uhlig 2007.)

B. Immediate Recommendations

Developing the agenda above is a long term project because it will require developing new models. Here we discuss two shorter term steps that are within the current frontier that represent an improvement over current practice.

First, researchers should explicitly incorporate long term interest rates into their monetary business cycle models to ensure that they capture changes in agents's expectations about the long run course of policy. Second, researchers should not simply stick exotic preferences from finance into an otherwise standard model in the hope that this will capture time-varying risk.

Cogley and Sbordone (2005) and Ireland (2007) nicely illustrate why it is critical to capture changes in agents' expectations about the long run course of policy in applied settings. As is well-known inflation is very persistent in the U.S. data. As we have shown agents expect the short rate to be very persistent as well. Standard models fail to capture the persistence of policy and hence regard the persistence of inflation as a puzzle. A popular way to make a model with nonpersistent generate persistent inflation is to mechanically assume that prices are backward indexed. (See, for example, Christiano, Eichenbaum, and Evans 2005 and Smets and Wouters 2007.) Cogley and Sbordone (2005) and Ireland (2007) find that once the persistence of policy is properly accounted for the model matches the persistence of inflation even when there is no backward indexation and it fit deteriorates if backward indexation is included. Distinguishing between the two ways of generating persistence is critical because the costs of disinflation are large if the persistence is coming from backward indexation and the costs are trivial if the persistence is coming from policy. Indeed, as Chari, Kehoe, and McGrattan (2008) stress if the persistence is actually coming from policy and not from backward indexation then the policy advice from a model in which it is mechanically assumed to be coming from backward indexation is not useful.

For some interesting preliminary work in explicitly modeling the term structure in a New Keynesian model see Bekaert, Cho, and Moreno (2006).

Next, we illustrate the difficulties that arise from simply sticking exotic preferences popular in finance into otherwise standard macroeconomics models with the hope that doing so will improve results when the model is estimated on U.S. data. We are very skeptical that this approach will lead to a step forward in our modeling of monetary policy and the economy.

To understand our skepticism consider the model of Smets and Wouters (2007) that is widely regarded as the state-of-the-art monetary model. This model incorporates external habit persistence in consumption in the representative agent's utility function. In evaluating the role of habit persistence in this model we find it useful to the log-linearized Euler equation of the form (14) in which the constant term has been replaced by the error in the Euler equation ε_{bt} as in

$$(27) \quad i_t = -E_t \log \frac{U_{ct+1}}{U_{ct}} + E_t \log \pi_{t+1} + \varepsilon_{bt}.$$

The term i_t is given in data by the Fed funds rate. A time series for the conditional means of marginal utility and inflation

$$(28) \quad -E_t \log \frac{U_{ct+1}}{U_{ct}} + E_t \log \pi_{t+1}$$

on the right side of (27) can be computed from the data using the model to compute the conditional expectations at each date. In a slight abuse of terminology we refer to the terms in (28) as the *predicted short rate*.

In Figure 10 we plot the Smets and Wouters (2007) data on the Fed funds rate together with their model's predicted short rate using their estimated level of habit persistence. As is clear from the figure the model's prediction for the short rate is simply wild: it is extremely volatile and has little or no relation to the actual short rate.

We now show that the model's extreme predictions for the short rate come entirely from the inclusion of habit in consumer preferences. To do so we reestimated the model imposing that the habit parameter is zero. In Figure 11 we plot the same data on the Fed funds rate along with the model's predicted short rate under this specification. Here the predicted short rate does a much better job of tracking the actual short rate. In this sense, including at least this form of exotic preferences clearly does not represent progress.

5. Conclusion

What have put forward two stylized facts about the behavior of interest rates and argued that these facts call for a new approach to analyzing monetary policy.

We have argued that the low frequency data on long and short rates indicates that there has been a remarkable shift in the volatility of agents' expectations of long run averages of the future short rate. In the pre World War II data, it appears that agents had very stable expectations about the path of the short rate on average over the long term. After World War II, it appears that agents' long run expectations became unanchored — it appears that agents interpreted most of the high frequency movements in the short rate as nearly permanent changes in policy. More recently, in the data after 1990, it appears that agents' long run expectations may be finding an anchor again.

A central question in the analysis of monetary policy at low frequencies then is what institutional changes led to this pattern. To answer this question at a mechanical level, we see that the Gold Standard was the main institution governing monetary policy in the pre-war era and that after the war most countries switched to a fiat standard governed for part of the time by the Bretton Woods agreement. But this answer is, at best, a superficial one. In the pre war era, countries chose to be on the Gold Standard for the majority of the time and chose to leave it when it suited their purposes. Thus, the relevant question is what forces at a deeper level led agents to have confidence that their governments would choose stable policy over the long term, what forces led them to lose this confidence after World War II, and finally what forces led them to gain it again in the 1990s. Only if we can quantitatively account for this history can we give advice on how to avoid another Great Inflation.

This question is particularly relevant given the recent conflict at the Fed between fighting rising inflation and stimulating a potentially stagnating economy. In our hurry to moderate the current economic dip are we unhinging long-term expectations from their long-term anchor?

We have argued that at business cycle frequencies, most of the movements in the short rate correspond to movements in risk, rather than to movements in the expected growth of marginal utility and inflation as assumed in the standard models. This fact gives us a new perspective as to what moves in the macro-economy when the interest rate moves.

A central question in the analysis of monetary policy at business cycle frequencies is what is the direction of causation? Is the Fed reacting passively to real risks in the economy or is the Fed itself causing some of this risk? Until we answer this question we can't answer the key counterfactual questions like what would happen if the Fed changed policy.

Appendix A

Proof of Proposition 3

To see this result, observe that

$$E_t r_{t+1}^k = E_t p_{t+1}^{k-1} - p_t^k = E_t \log E_{t+1} Q_{t+k} - E_t \log Q_{t+1} - (\log E_t Q_{t+k} - \log Q_t).$$

where the second equality follows from (23). Using conditional log-normality one-step ahead, the short-term interest rate is given by

$$i_t = -E_t (\log Q_{t+1} - \log Q_t) - \frac{1}{2} var_t (\log Q_{t+1} - \log Q_t). \tag{29}$$

Substituting this expression for the short rate into the expressions for the expected return above gives

$$E_t r_{xt+1}^k = \frac{1}{2} var_t (\log Q_{t+1} - \log Q_t) + (E_t \log E_{t+1} Q_{t+k} - \log E_t Q_{t+k}). \tag{30}$$

Consider next the term

$$E_t \log E_{t+1} Q_{t+k} - \log E_t Q_{t+k}$$

in (30). Note that $E_{t+1} Q_{t+k}$ is random as of period t and, by the Law of Iterated Expectations, $E_t Q_{t+k}$ is the conditional expectation of this random variable given information at time t. Thus, since the function *log* is strictly concave, we have in general

$$E_t \log E_{t+1} Q_{t+k} < \log E_t Q_{t+k}.$$

The magnitude of the difference between these two quantities is typically increasing in the variance of $E_{t+1} Q_{t+k}$. Thus, all else equal, the more uncertainty there is about the conditional expectation $E_{t+1} Q_{t+k}$, the lower is the expected excess holding period return to a long term bond. With our assumption that the conditional expectation $E_{t+1} Q_{t+k}$ is itself conditionally lognormally distributed given information available in period t, we have

$$E_t \log E_{t+1} Q_{t+k} = \log E_t Q_{t+k} - \frac{1}{2} Var_t \log E_{t+1} Q_{t+k}$$

This gives the result. *Q.E.D.*

Appendix B

Affine pricing kernel example that satisfies assumptions of proposition 4 Here we present the “negative" Cox-Ingersoll-Ross from Backus, Foresi, Mozumdar, and Wu (2001). In affine models of bond prices, bond prices p_{t+1}^{k-1} are linear functions of a state vector z_{t+1} that is conditionally log normally distributed with possibly heteroskedastic innovations and the pricing kernel $\log m_{t+1}$ is hit by the same innovations.

In our example, we have a process for the single state variable

$$z_{t+1} = (1-\varphi)\theta + \varphi z_t + \sigma z_t^{1/2}\epsilon_{t+1}$$

and the pricing kernel given by

$$-\log m_{t+1} = \delta + (-1+\lambda^2/2)z_t + \lambda z_t^{1/2}\epsilon_{t+1}$$

with $0 < \varphi < 1$, $(1-\varphi)\theta > \sigma^2/2$ and ϵ_{t+1} a standard normal innovation. The conditional variance of innovations to both the state z_{t+1} and the pricing kernel $\log m_{t+1}$ varies with the state z_t. In particular

$$Var_t(z_{t+1}) = \sigma^2 z_t$$

and

$$Var_t(\log m_{t+1}) = \lambda^2 z_t$$

In this model, bond prices are linear functions of the state given by

$$p_t^k = -A_k - B_k z_t$$

where A_k and B_k are defined by the recursion

$$A_{k+1} = \delta + A_k + B_k(1-\varphi)\theta$$

$$B_{k+1} = -1 + \lambda^2/2 + B_k\varphi - (\lambda + B_k\sigma)^2/2$$

starting with $A_0 = B_0 = 0$. Note that since $A_1 = \delta$ and $B_1 = -1$, the short rate in this model is given by $i_t = \delta - z_t$.

We can now compute the expected excess returns in this model. The term

$$var_t(\log m_{t+1}) = \lambda^2 z_t$$

Since we have

$$p_{t+1}^{k-1} = \log E_{t+1} Q_{t+k} - \log Q_{t+1}$$

and, in standard notation

$$\log m_{t+1} = \log Q_{t+1} - \log Q_t$$

we have

$$Var_t(\log E_{t+1} Q_{t+k} - \log E_t Q_{t+k}) = Var_t\left(p_{t+1}^{k-1} + \log m_{t+1}\right)$$

In this model,

$$p_{t+1}^{k-1} + \log m_{t+1} =$$

$$-A_{k-1} - B_{k-1}\left[(1-\varphi)\theta + \varphi z_t + \sigma z_t^{1/2} \epsilon_{t+1}\right] - \delta - (-1 + \lambda^2/2) z_t - \lambda z_t^{1/2} \epsilon_{t+1}$$

Gathering terms in ϵ_{t+1} gives

$$Var_t(\log E_{t+1} Q_{t+k} - \log E_t Q_{t+k}) = Var_t\left(p_{t+1}^{k-1} + \log m_{t+1}\right) = (B_{k-1}\sigma + \lambda)^2 z_t$$

Hence, plugging these results into (24) implies that excess returns are given by

$$E_t r_{xt+1}^k = \frac{1}{2}\left[\lambda^2 - (B_{k-1}\sigma + \lambda)^2\right] z_t$$

The slope of the regression coefficient of expected excess returns for a bond of maturity k on the spread between the yield on a bond of maturity L and the short rate is given by

$$\text{slope} = \frac{1}{2} \frac{\left[\lambda^2 - (B_{k-1}\sigma + \lambda)^2\right]}{-B_L/L + 1}$$

where the expression in the denominator is the yield spread $y_t^L - i_t$ as a function of z_t. If we set the persistence of the state $\varphi = .99$, the constant $\delta = .05/12$ (5% on a monthly basis), $\lambda = \sqrt{2}$, and $\sigma = 0.0088$, and use the 13 year zero coupon yield for y_t^L, we get regression slopes greater than one for long yields and both short and long term risk rise when the yield spread rises. This is verified because $B_k\sigma$ converges to $-.43$ for large k so that $B_k\sigma + \lambda > 0$.

Note that when $\lambda = \sqrt{2}$, then the conditional mean of the pricing kernel is constant and all movements in the term spread are movements in the conditional variance just as discussed in Alvarez, Atkeson, and Kehoe (2007).

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Figure 1:
3 month T-Bill yield and 13 year zero coupon yield

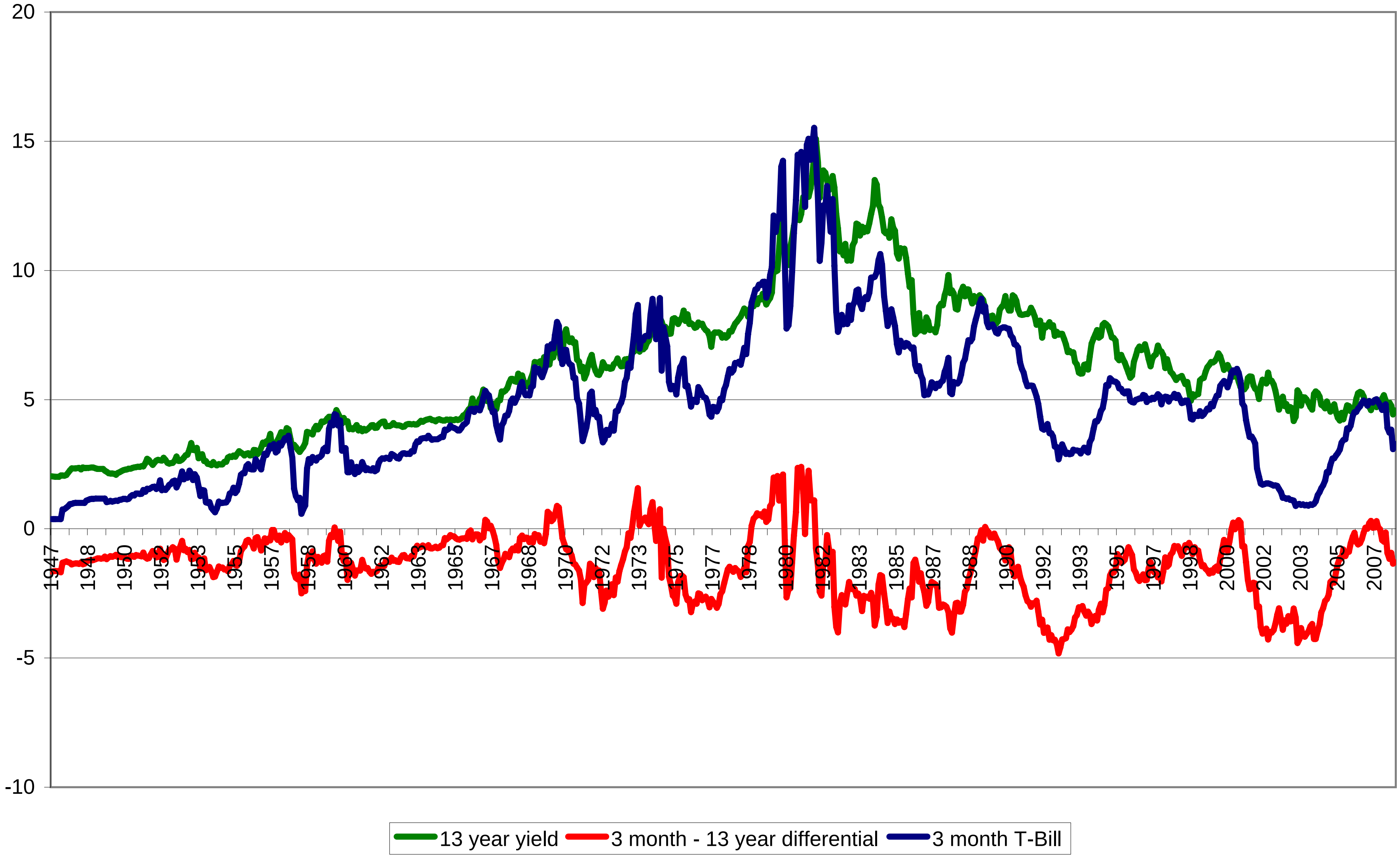

Figure 2:
13 year, 23 year and 30 year zero coupon yields

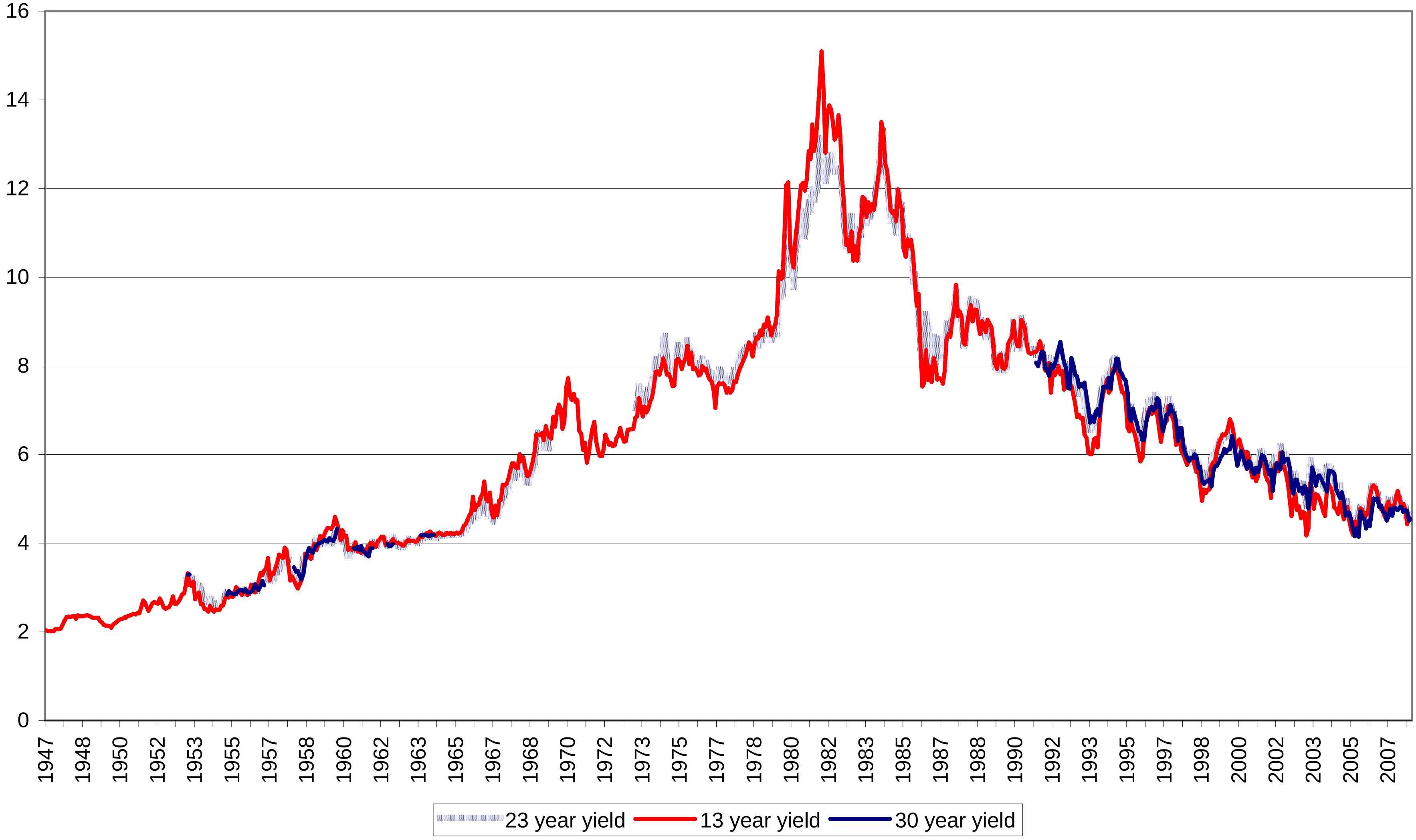

Figure 3:
Federal Funds Target Rate and 3 month T-Bill

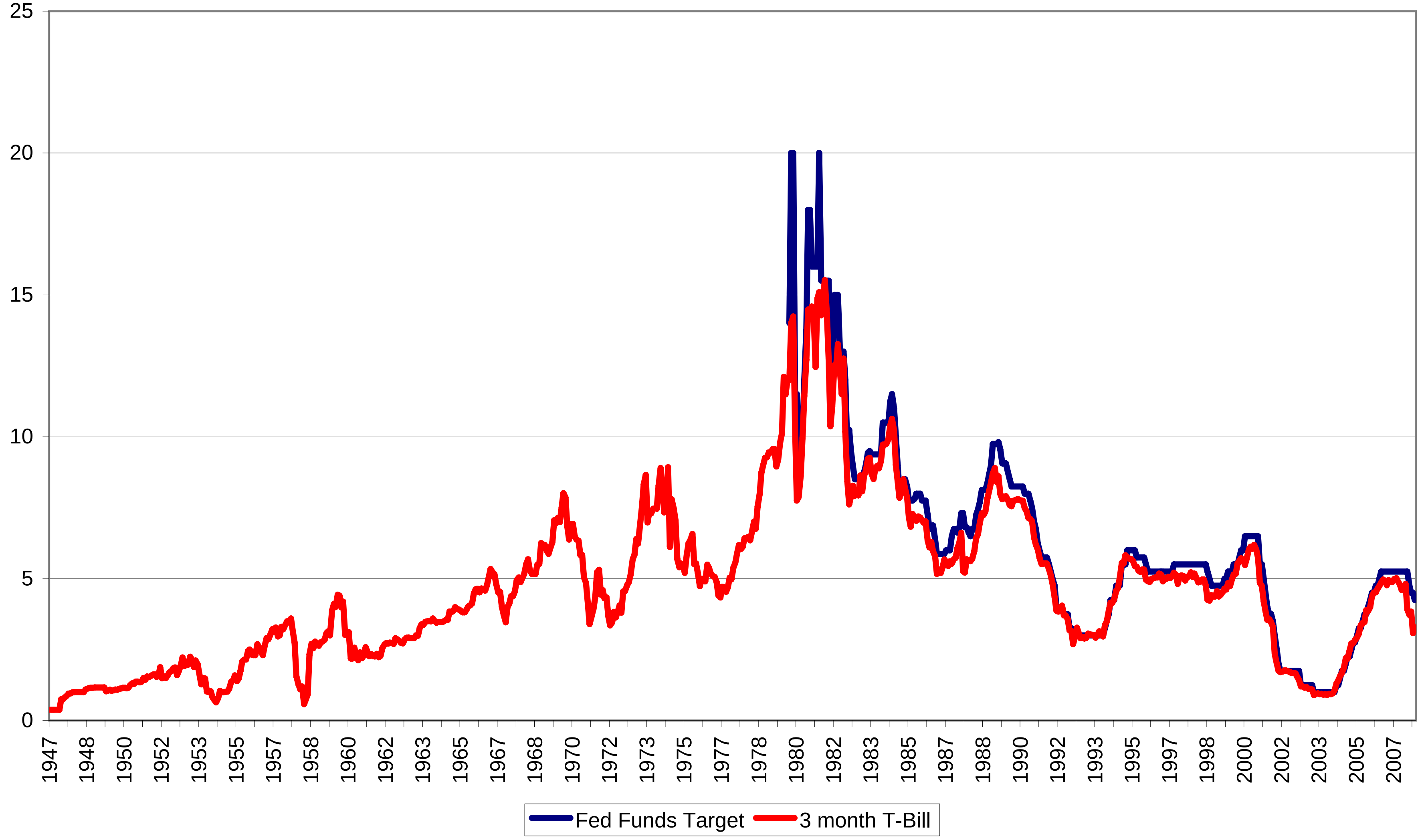

Figure 4A:
Long and Short Rates in the United States

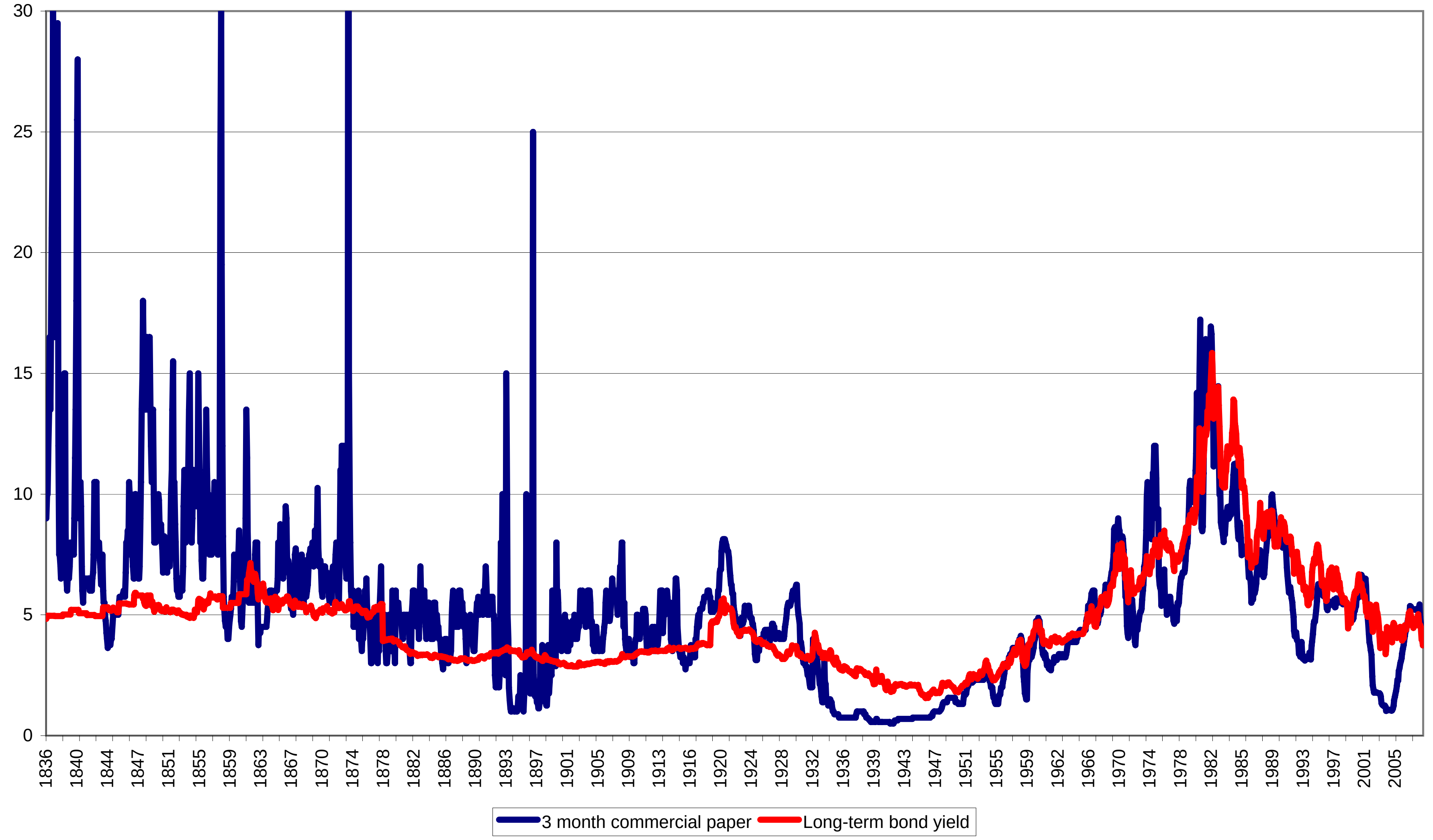

Figure 4B:
Long and Short Rates in the United Kingdom

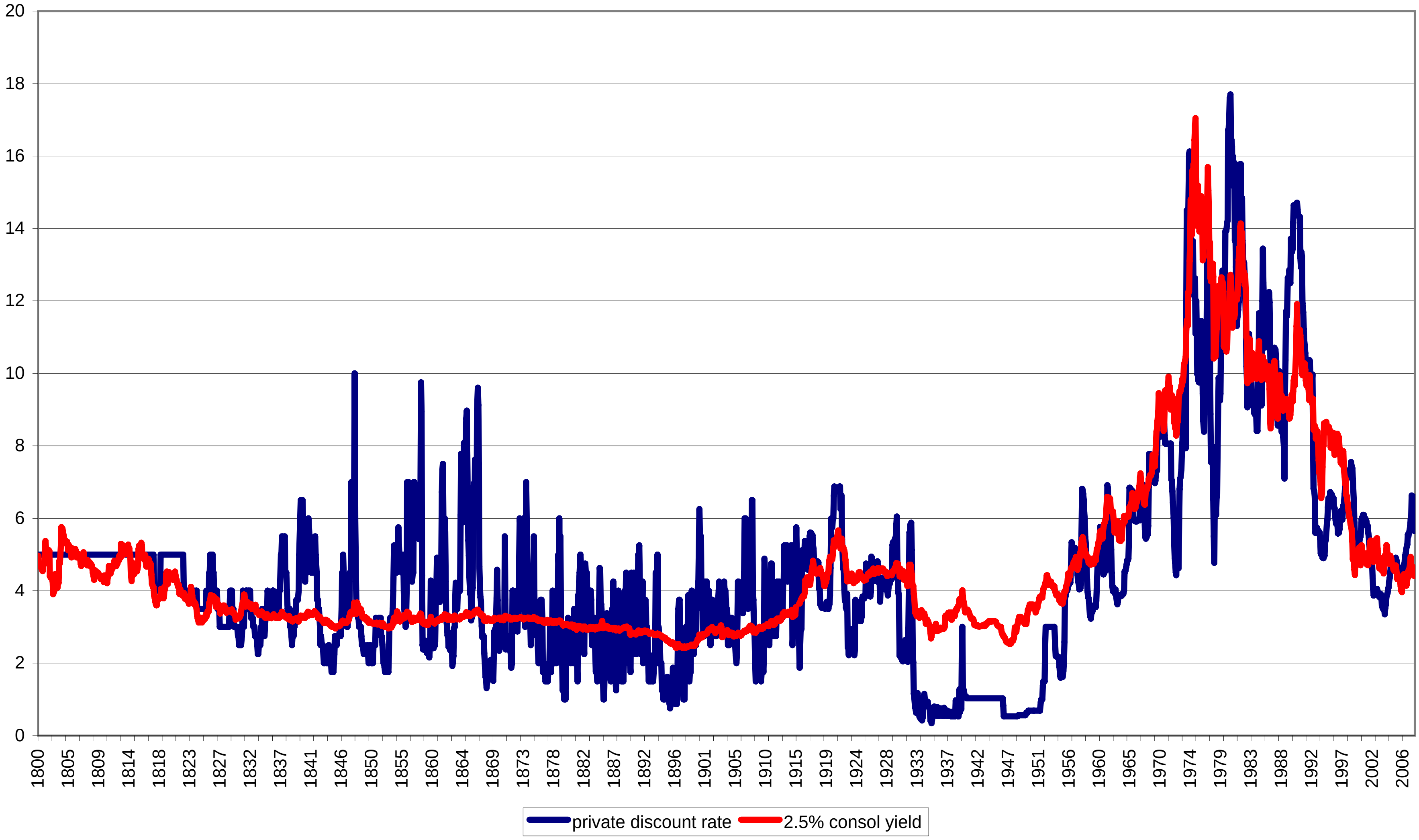

Figure 4C:
Long and Short Rates in France

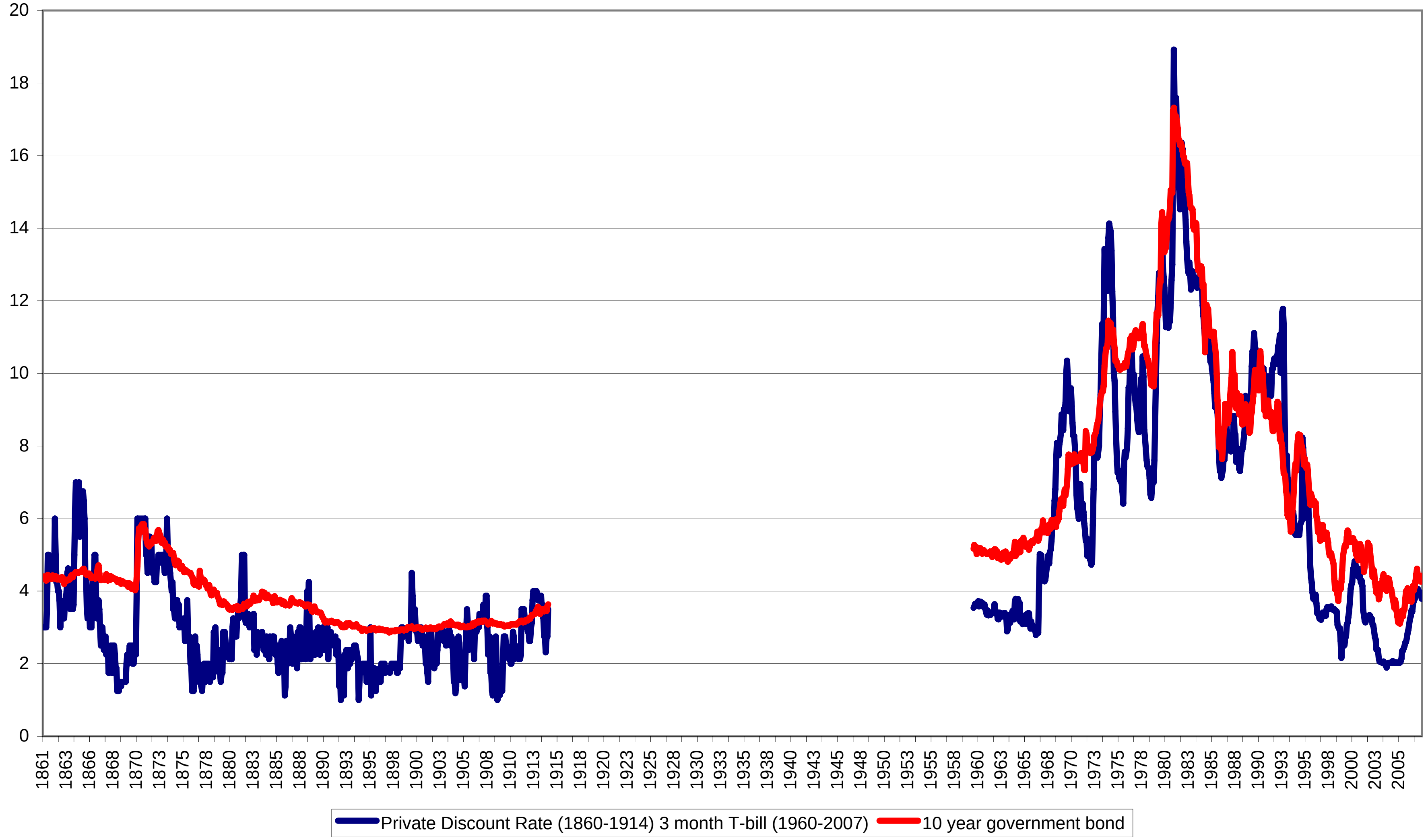

Figure 4D:
Long and Short Rates in Germany

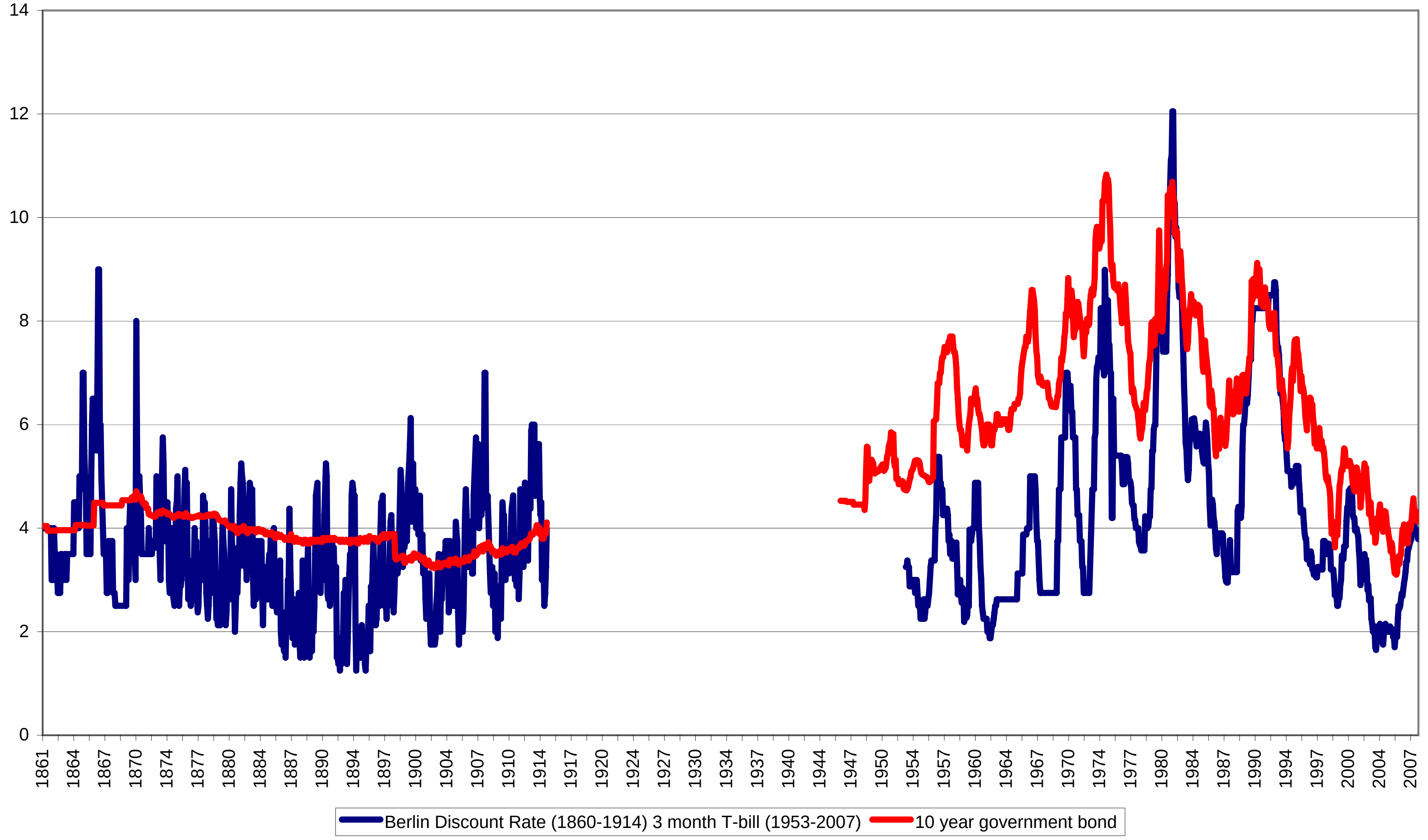

Figure 4E:
Long and Short Rates in Netherlands

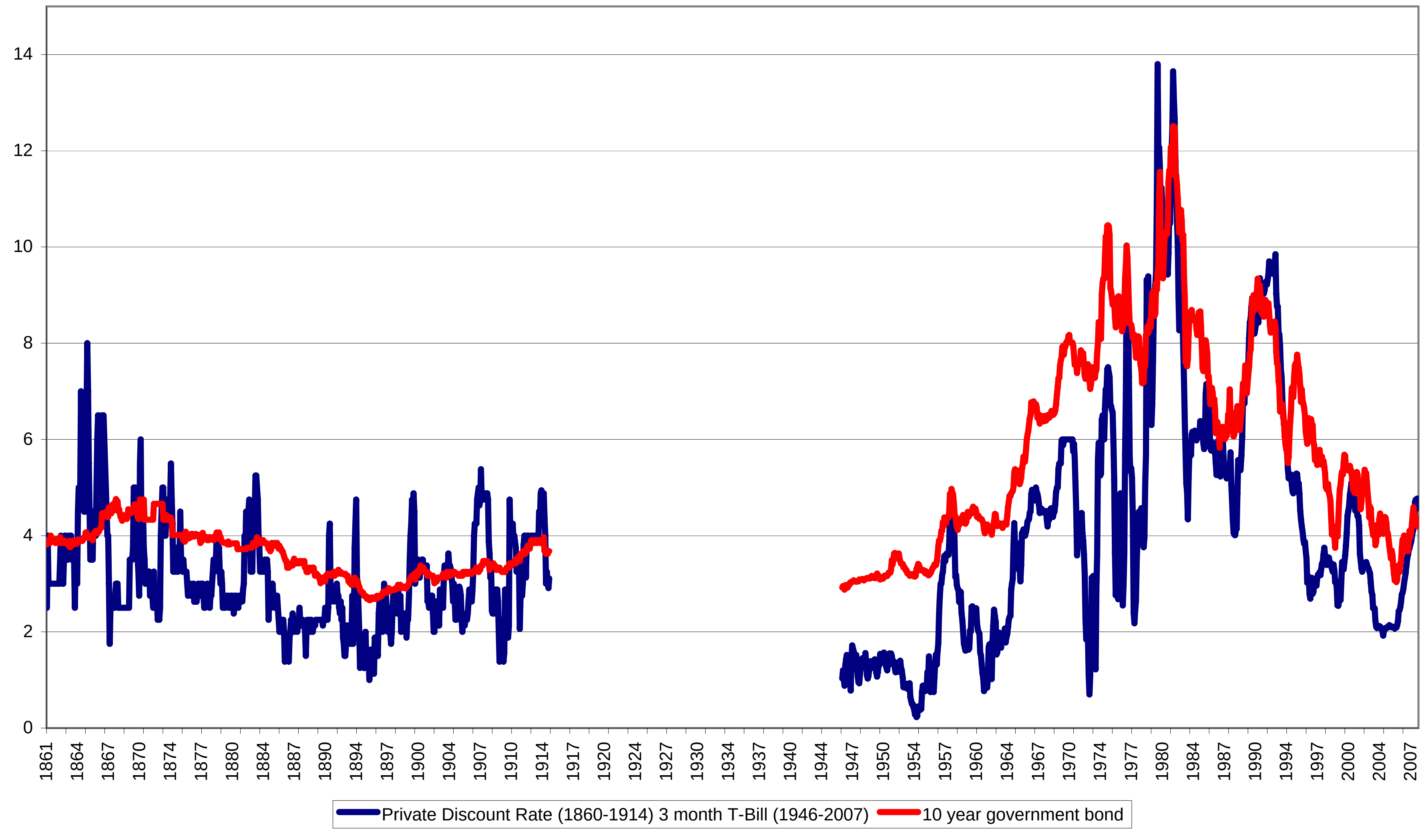

Figure 5:
Yield Spread and HP Filtered Short Rate

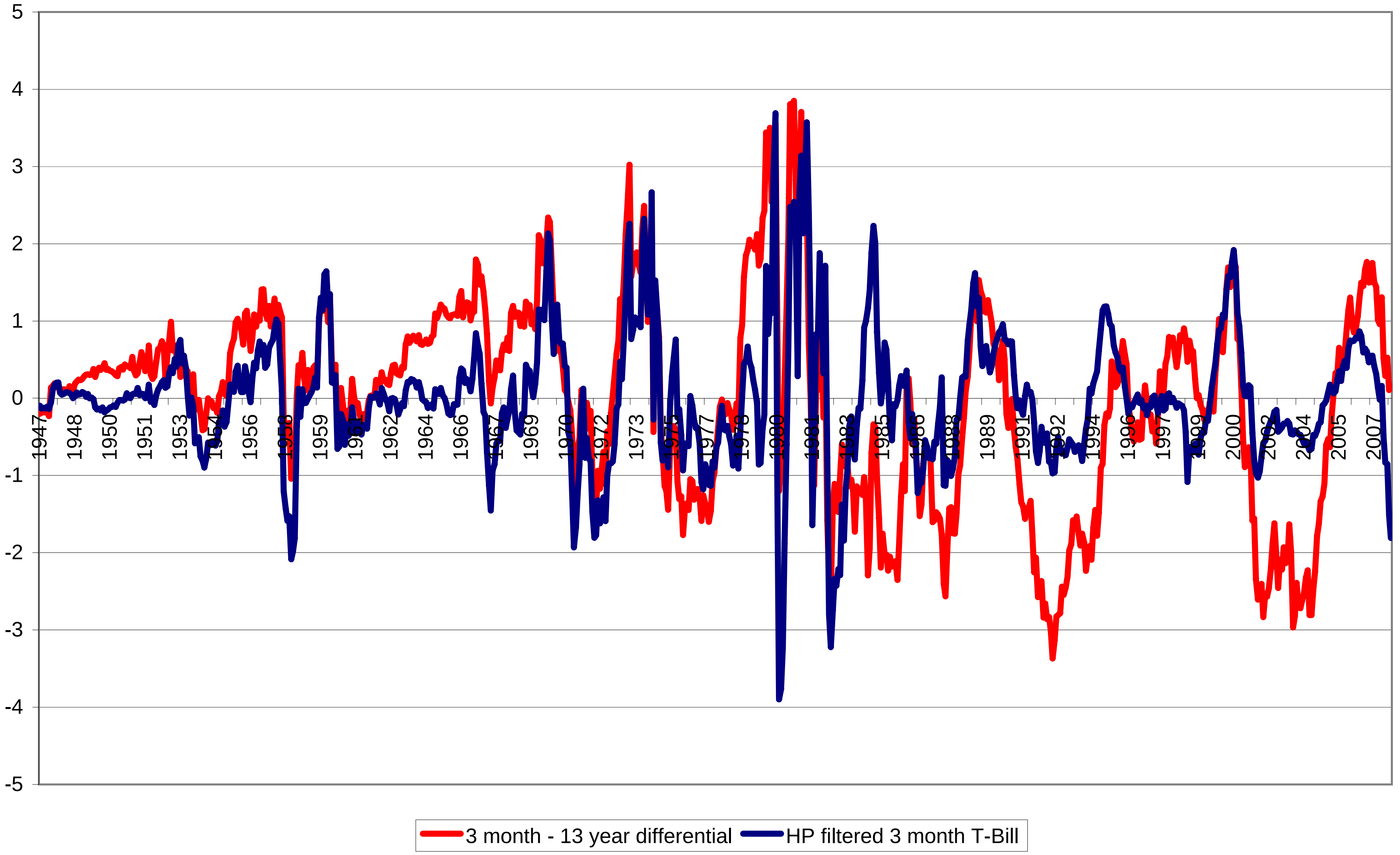

Figure 6: Movements of 13 year yield and our Cochrane-Piazzesi based expectations of 13 year averages of expected short rate

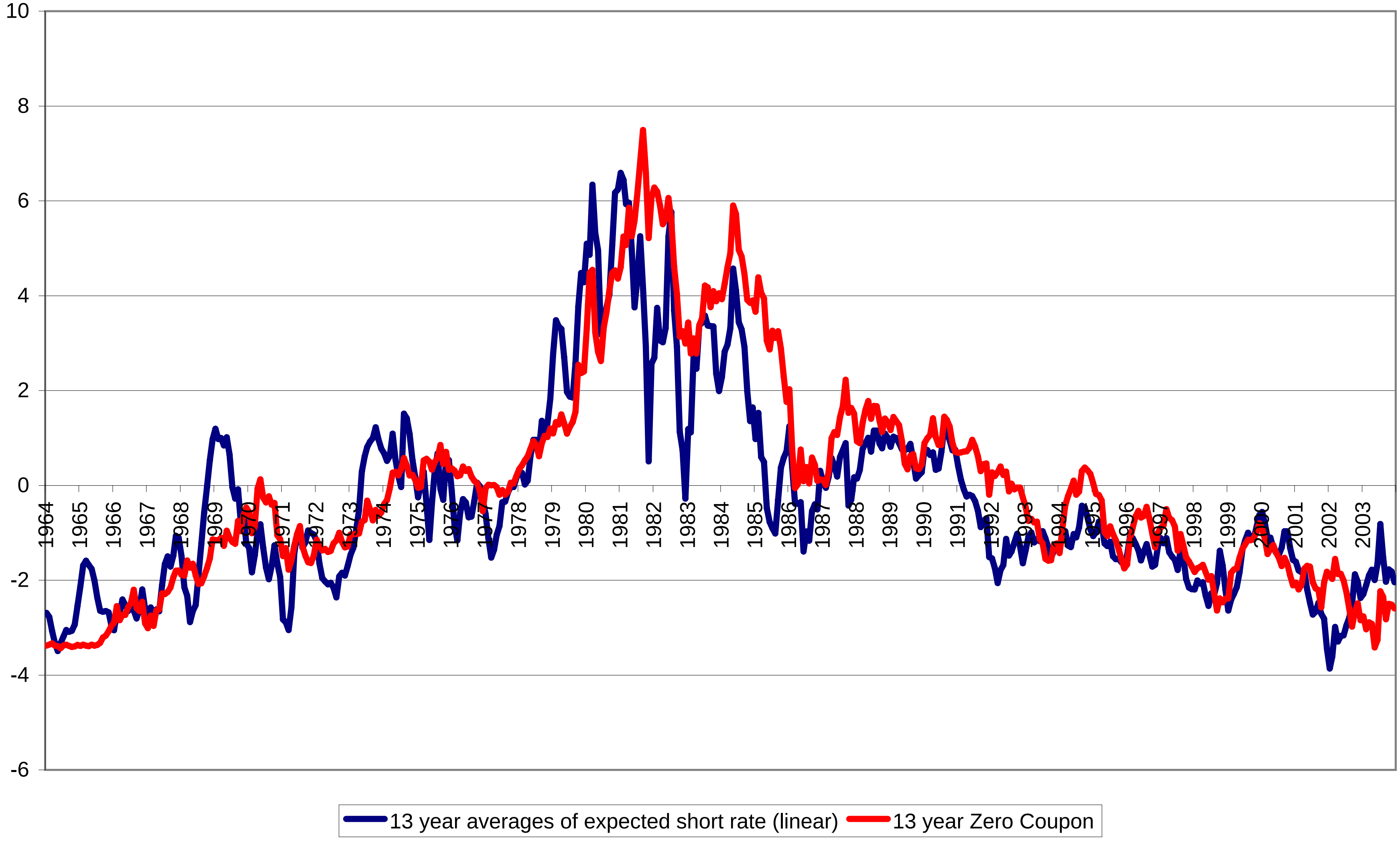

Figure 7 - HP Filtered Fed Funds Rate And HP Filtered Euler Equation Error Without Habit

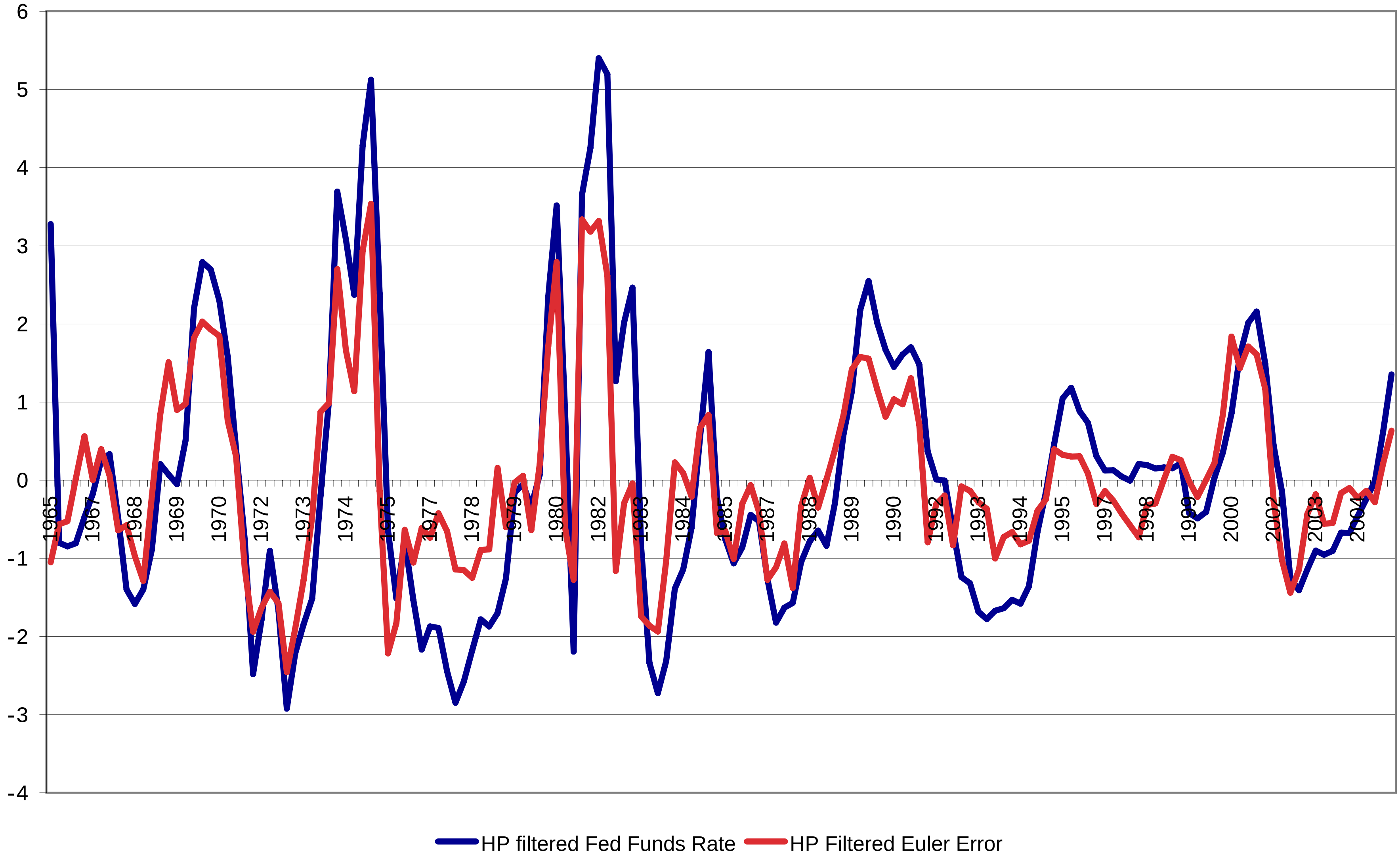

Figure 9: Sargent-Williams-Zha expectations of 20 and 30 year average inflation and 13 year yield

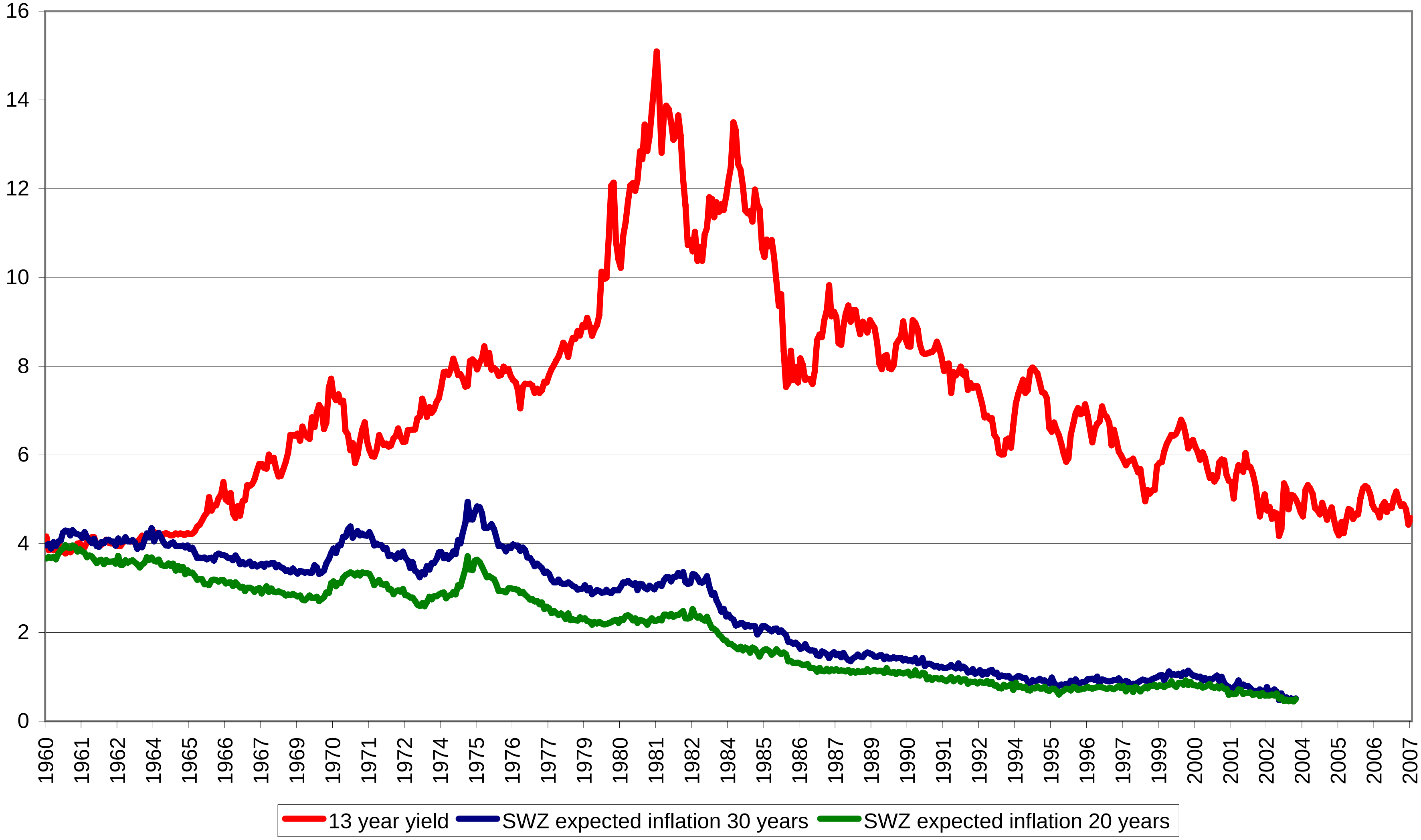

Figure 10: Federal Funds Rate and Predicted Interest Rate with Habit

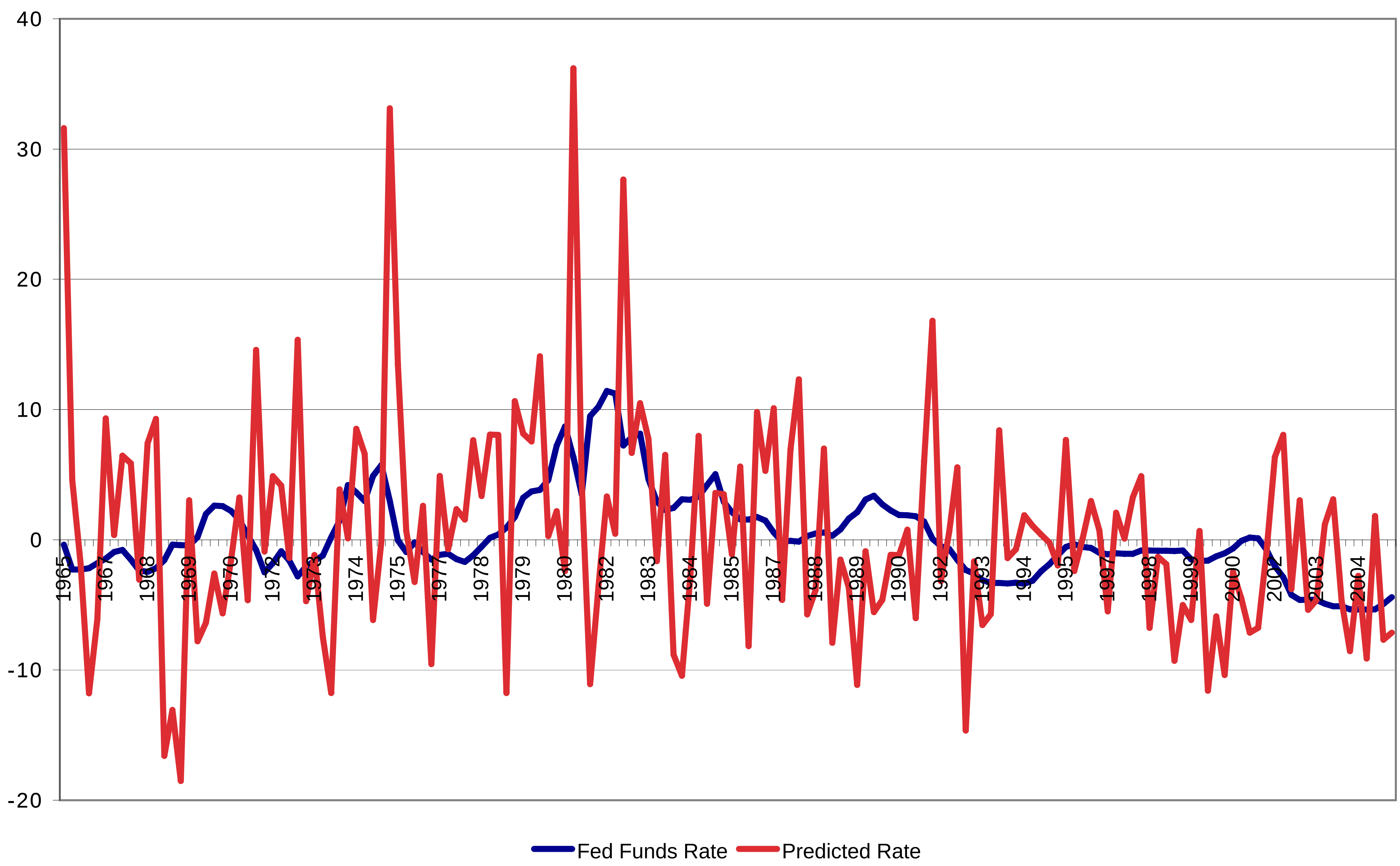

Figure 11: Federal Funds Rate and Predicted Interest Rate without Habit