Inflation Risk Premia and Survey Evidence on Macroeconomic Uncertainty

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Abstract

Nominal and real U.S. interest rates (1997–2007) are combined with inflation expectations from the Survey of Professional Forecasters to calculate time series of risk premia. It is shown that survey data on inflation and output growth uncertainty, as well as a proxy for liquidity premia can explain a large amount of the variation in these risk premia.

Keywords: break-even inflation, liquidity premium, Survey of Professional Forecasters.

JEL Classification: E27, E47.

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1 Introduction

The idea of this paper is very simple. Data on nominal and real (inflation indexed) interest rates is combined with survey data on inflation expectations to construct quarterly time series of “inflation risk premia” for the sample 1997–2007. These time series are then regressed on proxies for liquidity premia and measures of inflation and output uncertainty (from survey data on probability distributions). The results indicate that the regressors are significant and explain a considerable fraction of the movements of the inflation risk premia.

Several studies have made use of real interest rates, liquidity premia or survey data, but very few (if any) have combined them to have data on all the key ingredients of a modern Fisher equation.

Reschreither (2004), Chen, Liu, and Cheng (2005) and D’Amico, Kim, and Wei (2008) use nominal and real interest rates to estimate no-arbitrage yield curve models with latent (unobservable) factors. This paper instead estimates a (not fully specified) yield curve model using only observable factors that measure beliefs of economic agents: survey measures of inflation and output uncertainty and disagreement. It is potentially very important to have direct measures of subjective market beliefs—since it allows us to separate the issue of the economic theory (about the inflation risk premium) from that of rational expectations/small sample problems. A particular strength of the survey data used in the current paper is that it allows us to calculate explicit measures of both uncertainty and disagreement.

Lahiri, Teigland, and Zaporowski (1988) and García and Manzanares (2007) also use survey data to estimate inflation uncertainty (and even skewness/kurtosis). The former paper does not find any effect of uncertainty on nominal interest rates (on the sample 1969–1986), but García and Manzanares (2007) find that it is important for understanding the inflation scares during the Volcker period in the early 1980s. (Both papers find that higher moments may be important). The current paper differs from these earlier contributions in many details (in particular, on how to estimate uncertainty from survey data), but more importantly by focusing on the period when real (inflation indexed) interest rates are available (1997–). This allows for controlling for one potentially important driving force of nominal rates.

Carlstrom and Fuerst (2004) and Shen (2006) compare the “break-even inflation rate
(the difference between the nominal and real yields) to survey measures of inflation expectations. They argue that the break-even inflation rates are too low—due to a considerable liquidity premium in the observed real rates. The current paper incorporates those results—and goes on to study if other factors (inflation and output uncertainty/disagreement) add explanatory value.

The outline of the paper is as follows. Section 2 derives a modern Fisher equation; section 3 discusses the data and the estimation approach; section 4 presents the empirical results and section 5 concludes.

2 The Fisher Equation

This section presents the theoretical foundation of my empirical approach. A CRRA utility function gives the following first order condition for an investor’s investment in a nominal bond with interest rate $i$

$$1 = \beta E \exp (i - \pi - \gamma \Delta c),$$

(1)

where $\beta$ is the time discount rate, $E$ is the (conditional) expectation, $\pi$ is the inflation rate, $\gamma$ is the risk aversion and $\Delta c$ is consumption growth. Notice that $i - \pi$ is the real return on the bond. The first order condition for a real bond (without liquidity risk) is similar: substitute the real interest rate on a real bond without liquidity risk ($\tilde{r}$) for $i - \pi$.

If $i - \pi - \gamma \Delta c$ has a normal distribution, then the inflation risk premium satisfies

$$i - E \pi - \tilde{r} = -\gamma \text{Cov}(\pi, \Delta c) - \text{Var}(\pi) / 2,$$

(2)

where $E \pi$ is expected inflation. This expression shows that the inflation risk premium on a nominal bond equals a covariance term plus a Jensen’s inequality term. A negative covariance between inflation and consumption makes a nominal bond risky (since it has a low real return when consumption is low), so investors ask for a risk premium. The classical Fisher relation says that the risk premium is zero (or at least constant).

The real (inflation indexed) bonds have occasionally been somewhat illiquid assets and it is often argued that they carry a liquidity premium (for instance, see Shen (2006)). Let $r$ indicate the real interest rate observed in data and let $LPR$ indicate the real liquidity premium. Clearly, $r = \tilde{r} + LPR$, so the Fisher equation can (after dropping the Jensen’s
inequality term) be written

\[ i - E \pi - r = -\gamma \text{Cov} (\pi, \Delta c) - LPR. \quad (3) \]

Central banks often calculate a “break-even inflation” as the difference between nominal and real yields \((i - r)\). If the inflation risk premium and the liquidity premium in the Fisher equation (3) are constant, then the break-even inflation differs from inflation expectations by a constant. Changes in the break-even inflation are then unbiased measures of the changes in inflation expectations. Otherwise, an increased inflation risk premium increases the break-even inflation rate, while an increased real liquidity premium decreases it—and may hurt the informational value of the break-even inflation.

3 Data and Relation to the Fisher Equation

This section describes the data and how it is used to construct proxies for macroeconomic uncertainty and liquidity premia.

I use McCulloch’s (2008) end of month estimates of nominal and real zero coupon rates (available since 1997 when TIPS, Treasury Inflation-Protected Securities, were introduced), quarterly data on subjective beliefs from the Survey of Professional Forecasters (SPF) at the Federal Reserve Bank of Philadelphia (2008) and the daily proxy for the liquidity premium from the Federal Reserve Bank of Cleveland (2008). The effective sample is quarterly and covers the period 1997–2007.

The SPF is a quarterly survey of forecasters’ views on key economic variables. The respondents, who supply anonymous answers, are professional forecasters from the business and financial community. The survey is (since 1991) administered by the Federal Reserve Bank of Philadelphia (see Croushore (1993) for details). It asks for point forecasts of many macroeconomic and financial variables, but also for probability distributions (histograms) of real output growth and GDP deflator inflation. The deadline for the survey is approximately in the middle of the quarter. I therefore combine the survey data with interest rates and liquidity premia for the end of the first month of the quarter (using data for the end of the second month gives very similar results).

The survey data from the SPF is very interesting, but far from perfect, for investigating the empirical relevance of the (modern) Fisher equation. It is therefore necessary to discuss the compromises that need to be made.
The Fisher equation (2) applies to an individual investor: it is his/her first order condition. However, the survey asks about macroeconomic aggregates, not the respondent’s own consumption—and the forecasters in the SPF do not agree (especially not about the point forecasts). A solution to this dilemma is found in the asset pricing literature on heterogenous beliefs: it suggests that the difference between (planned) individual and aggregate consumption is a function of the disagreement among investors.\footnote{See, for instance, Anderson, Ghysel, and Juergens (2005), David (2008) and Giordani and Söderlind (2005).} For this reason, measures of disagreement about the point forecasts of inflation and real growth will be included in the empirical analysis. In addition, disagreement is often thought of as an alternative proxy for uncertainty, which gives another reason for including it.

The survey data has a number code for each respondent, so it should be possible to estimate the Fisher equation on individual data.\footnote{This gives a dynamic panel regression since the participants enter and leave the survey at different times.} However, there would be several problems with such an approach—in particular with how to handle outliers and other strange data points (discussed below). Instead, I estimate a Fisher equation on cross-sectional averages (trimmed means or medians) of all the terms. This is a valid way of doing it, since the first order condition (using disagreement to proxy for the difference between individual and aggregate consumption) should hold for each investor—and therefore for an average. It might sacrifice a bit of econometric efficiency, but makes it more straightforward to handle outliers.

The SPF contains only univariate distributions, and hence no information about the covariance of inflation and consumption in the theoretical asset pricing equation (3). I therefore approximate the covariance by a linear combination of the two standard deviations. One possible interpretation is that this is a Taylor series expansion.

To analyse long-term interest rates, we ought to have data on long-run beliefs. Unfortunately, the SPF asks mostly for short-run beliefs (except very recently). I am therefore left with using the short-run information—hoping that the “term spreads” of beliefs are linear functions of the short-term beliefs.

The Federal Reserve Bank of Cleveland (2008) calculates an adjusted break-even inflation by adding a real liquidity premium ($LPR$). Since it is difficult to construct a direct measure of the real liquidity premium, it is proxied by a function in terms of the yield difference between less and more liquid (“off the run” and “on the run”) 10-year nominal...
Figure 1: **Liquidity premium according to the Federal Reserve Bank of Cleveland (2008).** This figure shows a 4-quarter moving average of the liquidity premium on relatively less liquid nominal bonds, the functional form of equation (4) and the resulting liquidity premium on real bonds.

It is argued that liquidity premia on the nominal and real treasuries ($LPN$) are strongly correlated. In practice, the following function is used

$$LPR = -0.948 + 12.71 LPN - 20.9 (LPN)^2.$$  \hspace{1cm} (4)

*Figure 1* shows the time series of $LPN$, the functional form of equation (4), with markers for the quartiles of the sample, and the resulting time series of $LPR$. The liquidity premium peaks around 1999–2000 and 2002–2003 and has been markedly low since around 2005. The functional form of the $LPR$ suggests that it is virtually a linear function of $LPN$, except for some very rare cases (the peak in 1999), with a slope of 3.5 or so. The constant in (4) is meant to capture a constant inflation risk premium on the nominal bond.
I use LPN as a regressor together with my proxies for macroeconomic uncertainty. In practice, this means that I try to replace the constant inflation risk premium used by the Federal Reserve Bank of Cleveland (2008) with time-varying proxies of inflation risk. All in all, the main empirical specification is to regress the risk premium as

\[ i - \text{E} \pi - r \text{ on } \left[ \begin{array}{c} \text{individual uncertainty } \\ \text{disagreement} \end{array} \right] \text{ about } \left[ \begin{array}{c} \text{inflation} \\ \text{real growth} \end{array} \right] \text{ and } LPN. \]  

Notice that if the nominal interest rate and inflation expectations are non-stationary, but cointegrated (with a coefficient of unity), then (5) is expressed in terms of stationary variables (assuming, of course, that the real interest rate is stationary).

To summarize, an asset pricing model (with potentially heterogeneous beliefs) suggests that the nominal interest rate should be driven by the terms in (5). The coefficients are essentially an empirical issue as they depend on several unknown parameters: the risk aversion coefficient, the importance of disagreement and the properties of the approximations. The rest of this section discusses how the variables in the regression are constructed.

From the point forecasts, I use the median (across forecasters) CPI inflation and real consumption growth forecasts for several forecasting horizons: the next 4 quarters, next calendar year (with an effective horizon of 5–8 quarters, depending on the quarter of the survey) and for the CPI also the next 10 years. (Using the mean or some robust estimator of the central location gives very similar results.) Numbers for intermediate horizons are calculated by linear interpolation.

The nominal and real interest rates as well as the (median) inflation expectations are shown in Figure 2. I choose to focus on the horizons of 3 and 10 years: the real rates for shorter maturities are less reliable—and it can be shown that the regression results for maturities between 3 and 10 years are well approximated by straightforward interpolation.

Figure 3 shows the break-even inflation expectation \( (i - \pi - r) \) and the “inflation risk premium” \( (i - \text{E} \pi - r) \). There are marked cyclical swings in the inflation risk premia and they are strongly correlated across maturities: they peak in 1997, 2000, 2004–7 and are very low in 1999 and 2002–3. The aim of this empirical study is to investigate if these movements can be explained in terms of macroeconomic uncertainty and a liquidity premium.

Disagreement (about inflation and real growth) is measured by the inter-quartile range
Figure 2: **Interest rates (nominal and real) and inflation expectations.** This figure shows estimated nominal and real zero coupon rates from McCulloch’s homepage, as well as inflation expectations from the SPF.

(divided by 1.35) of the individual point forecasts. This is a robust estimator of the cross-sectional dispersion—and would coincide with the standard deviation if the cross-sectional distribution were Gaussian. (Using a cross-sectional standard error produces a very erratic time series, due to occasional strange forecasts.)

The individual uncertainty of deflator inflation and output growth is estimated by fitting distributions to the individual histograms in the SPF. The estimation procedure is as follows. If only one bin is used (that is, the respondent puts 100% of the probability on one of the prespecified bins), then I assume a triangular distribution within that bin. If two bins or more are used, then a normal distribution is fitted (a mean and a variance). The fitting is done by minimizing the sum of the squared deviations of the theoretical from the observed probabilities. (See Giordani and Söderlind (2003) for an early application and García and Manzanares (2007) for a critique of the least squares criterion.)
There is a lot of cross-sectional (across forecasters) dispersion in the fitted values from the individual histograms—some of which appears to be caused by typos and other data errors. To get robust estimates of the average uncertainty, I use the cross-sectional trimmed mean (20% trimming from both bottom and top) of the individual standard deviations. (Using the median gives very similar results and using the mean produces fairly similar, but more erratic, results.)

The probability distributions (histograms) of real output growth and inflation are for year-on-year data (the value in the current calendar year divided by the value in the previous calendar year, minus one). This means that the forecasting horizon varies: an almost four-quarter forecast is made in Q1 of year $t$, the three-quarter forecast is made in Q2 and so forth. The estimated uncertainty therefore has clear seasonality—as the effective fore-
casting horizon decreases over the calendar year. I therefore report results where all series (not just the uncertainty) are 4-quarter moving averages. It can be shown that regressions on seasonally adjusted (X12-ARIMA) uncertainty/disagreement data give similar results, but somewhat lower $R^2$ values. However, the moving average may induce serial correlation in the regression residuals below, so a Newey-West estimator of the covariance matrix is used.

*Figure 4* shows uncertainty and disagreement about inflation and real growth for the current year (4-quarter point forecasts, and uncertainty about the current calendar year) and next calendar year. There are significant movements in these data series, with some
interesting patterns. First, the results for year 1 and year 2 convey virtually the same information. The uncertainty about the next calendar year is higher for both variables, but the correlations are very strong (around 0.7). Inflation disagreement is almost the same for both years and the two series for real output growth disagreement are also strongly correlated (around 0.8). Second, inflation uncertainty and disagreement are strongly correlated (with a correlation coefficient of 0.7). Third, real growth uncertainty and disagreement are less correlated (0.34) and the disagreement series looks very jumpy. Fourth, inflation and real growth uncertainty are only mildly correlated (0.3).

4 Empirical Results

This section reports the empirical results for quarterly data from 1997 to 2007 for the 3- and 10-year maturities. It can be shown that results for intermediate maturities are well approximated by linear interpolation. All data is in the form of 4-quarter moving averages—as a simple way to handle the strong seasonality in the survey data, due to the time variation in the effective forecasting horizon. (Regressions on seasonally adjusted uncertainty/disagreement data give very similar results, but somewhat lower $R^2$ values.)

Figure 5 shows the risk premium $(i - E \pi - r)$ and some of the key regressors. Only regressor values for year 1 are shown and used in the subsequent analysis, since the data for year 2 contributes very little extra information (compare Figure 4) and since the data for year 1 appears to be somewhat more reliable (in particular, it has fewer suspicious histograms).

4.1 Explaining the Observed Risk Premium with the Nominal Liquidity Premium

The potential fit of the risk premium $(i - E \pi - r)$ and the nominal liquidity premium is obvious from Figure 5. This is also seen in the first column of Table 1 which reports results from an OLS regression of the 10-year risk premium on the nominal liquidity premium (and a constant). The t-statistics (in parentheses) are based on a Newey-West estimator of the covariance matrix, using 4 lags. The slope coefficient is significantly negative, with a value that fits well with the slope of the functional form of (4)—previously illustrated in Figure 1—and the $R^2$ is high. (Based on the earlier findings, I use only a linear term.) This confirms the usefulness of the nominal liquidity premium, as argued by the Federal Reserve Bank of Cleveland (2008).
Figure 5: **Risk premia and the key regressors.** This figure shows the risk premium \((i - E\pi - r)\), liquidity premium, inflation uncertainty and real growth uncertainty. All data is in the form of 4-quarter moving averages.

It can be shown that the nominal liquidity premium is correlated with several of the other potential regressors. In particular, it is strongly correlated with real growth uncertainty and disagreement (correlation coefficients of 0.55–0.75). Moreover, the nominal liquidity premium is also strongly correlated with CBOE’s volatility index VIX (0.80), which is an average of implied volatilities from stock options—and should therefore provide a summary measure of the market’s belief about future stock market volatility. These correlations are not surprising since it is often argued that liquidity premia are related to both macroeconomic uncertainty and events that are specific to financial markets. In contrast, the nominal liquidity premium is only weakly (and negatively) correlated with inflation uncertainty and disagreement.

Columns 2–4 of Table 1 show that both real growth uncertainty and disagreement could potentially capture the same effects as the nominal liquidity premium: the coef-
Table 1: **Regression results, risk premium for the 10-year maturity.** The table shows regression coefficients and t-statistics (in parentheses) for quarterly data 1997–2007. The dependent variable is the nominal interest rate minus the real interest rate and minus expected inflation (10-year horizon). The t-statistics are based on a Newey-West estimator with 4 lags.

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The coefficients are significantly negative. However, the $R^2$ values are clearly lower than when using the nominal liquidity premium. Interestingly, using the VIX gives a strongly significant coefficient and actually a higher $R^2$ than with the nominal liquidity premium. I will still keep the liquidity premium in the subsequent regressions, since it is by now well established by the Federal Reserve Bank of Cleveland (2008) and since it is focused on the liquidity on the treasury market, rather than general financial uncertainty.

### 4.2 Adding Proxies for Macroeconomic Uncertainty (and Disagreement)

**Table 2** reports results from a number of regressions of the 10-year risk premium. The first column shows that both inflation uncertainty and disagreement have significantly positive effects on the risk premium, output growth uncertainty has a significantly negative effect while real growth disagreement has a small and insignificant coefficient. The latter variable is dropped in the second column and all further regressions (it is almost invariably insignificant in these and other regressions). The liquidity premium still has a negative effect, although it is now only borderline significant. The $R^2$ is 0.78, which is significantly higher than the 0.64 obtained from using only the liquidity premium (see Table 1).
The remaining columns in Table 2 report results when some further variables are dropped. In short, the results are that inflation uncertainty and disagreement are close substitutes (columns 3 and 4), that dropping output uncertainty lowers the $R^2$ quite a bit (column 5), but that dropping the liquidity premium does not (column 6).

One interpretation of the results in Table 2 is that (i) uncertainty/disagreement about inflation seems to induce an “inflation risk premium” and (ii) output growth uncertainty and the nominal liquidity premium together capture the liquidity premium on real bonds.

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Table 2: **Regression results, risk premium for the 10-year maturity.** The table shows regression coefficients and t-statistics (in parentheses) for quarterly data 1997–2007. The dependent variable is the nominal interest rate minus the real interest rate and minus expected inflation (10-year horizon). The t-statistics are based on a Newey-West estimator with 4 lags.

The results for the 3-year maturity in Table 3 are similar, with a few exceptions. First, inflation uncertainty is never significant if inflation disagreement is included in the regression. Second, the liquidity premium is even less significant than before. Third, the $R^2$ values are generally lower.

*Figure 6* emphasises the economic importance of the different regressors by illustrating the regression results for the 10-year maturity in Table 2. To keep the figure simple, we focus on a regression with only inflation uncertainty, output uncertainty and the liq-
Table 3: Regression results, risk premium for the 3-year maturity. The table shows regression coefficients and t-statistics (in parentheses) for quarterly data 1997–2007. The dependent variable is the nominal interest rate minus the real interest rate and minus expected inflation (3-year horizon). The t-statistics are based on a Newey-West estimator with 4 lags.

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<td>-3.33</td>
<td>-5.62</td>
<td></td>
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<tr>
<td></td>
<td>(-1.25)</td>
<td>(-1.04)</td>
<td>(-1.04)</td>
<td>(-1.88)</td>
<td>(-3.40)</td>
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<td>constant</td>
<td>1.96</td>
<td>2.42</td>
<td>2.39</td>
<td>1.33</td>
<td>0.33</td>
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<tr>
<td></td>
<td>(1.10)</td>
<td>(1.72)</td>
<td>(2.06)</td>
<td>(0.93)</td>
<td>(0.33)</td>
<td>(1.95)</td>
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<td>R2</td>
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<td>0.64</td>
<td>0.56</td>
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<tr>
<td>obs</td>
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The figure tells the following story for the peaks and troughs of the risk premium, $i - E\pi - r$. The local peak of the risk premium in late 1997 and early 1998 is mostly due to low output growth uncertainty and low liquidity premium. This gives a high risk premium since both these regressors have negative coefficients. The trough in 1999 is due to all three variables: low inflation uncertainty, high output uncertainty and high liquidity premium. The peak in 2000 is mostly driven by a sharp increase in the liquidity premium.
Figure 6: **Risk premia and the effect of key regressors (demeaned), 10-year maturity.** This figure shows demeaned risk premia ($i - E\pi - r$) as well as demeaned regressors times the respective slope coefficients (from column 4 in Table 2). All data is in the form of demeaned 4-quarter moving averages.

(although the other variables both move in the “right” direction). The trough in 2002–2003 is driven by high output uncertainty and liquidity premia—in spite of an increase in inflation uncertainty. The sustained high level from mid 2004–2007 is a mix of several forces: the liquidity premium has fallen throughout most of this period (it picks up again in late 2007), output growth uncertainty fell 2005–2006 but picked up again in 2007, while inflation uncertainty increased rapidly in 2004, fell back in 2005, but has increased thereafter.

Towards the end of the sample, all three regressors work in the same direction: making the risk premium higher than the historical average. For this period, adjusting the break-even inflation rate with just a historical average—or by just the liquidity premium—is therefore likely to exaggerate the inflation expectations. It would actually be better to
5 Concluding Remarks

This paper tries to explain observed “inflation risk premia” (nominal minus real interest rates minus survey investor on inflation expectations) by proxies for liquidity premia as well as inflation and output uncertainty. It is shown that these variables are both statistically and economically significant.

There are several caveats, however. The sample is short, has only a quarterly frequency and must be seasonally adjusted. The reasons are that investor on real interest rates starts in 1997, the survey investor is quarterly and it has a forecasting horizon that varies over the calendar year.

In spite of these limitations, it is interesting to study a modern Fisher equation where we have at least proxies for the terms that asset pricing theory suggests. The empirical results are encouraging since they do suggest plausible effects of macroeconomic uncertainty on nominal interest rates.

Evidence from quarterly investor is perhaps of limited help for central bank staff or others who need to adjust the latest figures on the “break-even inflation.” I still hope that this paper contributes to the understanding of the economic factors behind the inflation risk premium. Further research might be able to find high-frequency proxies for inflation and output uncertainty/disagreement and thereby improve current practices.
References


