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**Abstract**

This paper examines the relation between variations in perceived inflation uncertainty and bond premia. Using the subjective probability distributions available in the Survey of Professional Forecasters we construct a quarterly time series of average individual uncertainty about inflation forecasts since 1968. We show that this ex-ante measure of inflation uncertainty differs importantly from measures of disagreement regarding inflation forecasts and other proxies, such as model-based ex-post measures of macroeconomic risk. Inflation uncertainty is an important driver of bond premia, but the relation varies across inflation regimes. It is most important in the high-inflation regime early in the sample and the low-inflation regime over the last 15 years. Once the role of inflation uncertainty is accounted for, disagreement regarding inflation forecasts appears a much less important driver of bond premia.

**Keywords**: Survey expectations, probabilistic forecasts, heterogeneity, inflation uncertainty, bond risk premia.

**JEL** Classifications: E37, E44, E47, C53, G12

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# 1 Introduction

Economic theory suggests that investors must be compensated for risks associated with future macroeconomic developments. Uncertainty regarding future inflation and economic growth, for example, would be expected to be important drivers of risk premia.[1] Yet, there is little evidence of a direct link between variations in perceptions of macroeconomic uncertainty and risk premia in bond markets. Available evidence is indirect or relies on proxies of macroeconomic uncertainty.[2] A key reason for the lack of direct evidence is the difficulty in measuring the pertinent uncertainty. As any asset price, bond risk premia reflect information regarding current and future events that is available in real time. Therefore, to explain their variations over time, ideally we need a real-time measure of the uncertainty perceived by investors about future macroeconomic developments. Specifically, with regard to inflation, we need an ex-ante measure of inflation forecast uncertainty over a sample period long enough to include meaningful economic fluctuations.

In this paper, we provide direct evidence of the link between bond risk premia and ex-ante inflation uncertainty by focusing on a measure of individual perceptions of inflation uncertainty derived from the subjective inflation probability distributions in the Survey of Professional Forecasters (SPF). Building on the methodology developed in D'Amico and Orphanides (2008), we use the survey data from 1968 to 2013 to construct a quarterly time series of the average individual uncertainty about inflation forecasts and use this series to examine the relation between inflation uncertainty and bond risk premia.

The availability of a direct measure of inflation uncertainty allows us to draw comparisons with ex-post measures of uncertainty and other proxies such as interpersonal disagreement about inflation forecasts. For example, Wright (2011) shows that disagreement regarding inflation forecasts is significantly positively related to Treasury term premium estimates across many countries. Similarly, Buraschi and Whelan (2012) find that disagreement about real economic activity is important in explaining time variation in bond excess returns.[3]

Our ex-ante measure inflation uncertainty exhibits large variation over time, rising sharply just before or during most recessions, with the last recession being a notable exception. As we document, it does not correlate very highly with commonly-used proxies of inflation risk such as inflation disagreement and forecast accuracy, and exhibits only a weak relation to broader proxies of macroeconomic risk such as the VXO index of Bloom (2009), and the broad uncertainty factor of Jurado, Ludvigson, and Ng (2013).

Over the 1968-2013 sample period, inflation uncertainty is significantly positively related to bond risk premia. The correlation increases with the bond maturity. Disagreement regarding inflation forecasts loses its marginal explanatory power once we control for the

[1] For example, Buraschi and Jiltsov (2005), Piazzesi and Schneider (2006), Campbell, Sunderam, and Viceira (2007), Rudebusch and Swanson (2008), and others argue that nominal bond risk premia mainly reflect uncertainty about future inflation, that is, inflation risk.

[2] For example, some studies that focus on financial indicators show that risk premia in bond returns can be explained by a combination of bond yield spreads or forward rates (e.g., Campbell and Shiller 1991, Cochrane and Piazzesi 2005).

[3] Other relevant work includes de Rezende (2013) who uses survey consensus forecasts to generate a proxy of macro risks, and David and Veronesi (2013) who rely on the aggregate probabilistic distribution from the SPF to proxy inflation uncertainty. A difficulty with using the moments of such aggregate distributions, as discussed in Wallis (2005), is that they combine both forecast uncertainty and disagreement about point forecasts, so separating the two is necessary to focus on the role of uncertainty.

mean inflation forecast and uncertainty. These results are robust to the inclusion of a number of controls that prior works found to have strong predictive power for excess bond returns, including the macroeconomic factors of Ludvigson and Ng (2009) and proxies for macroeconomic disagreement similar to Buraschi and Whelan (2012).

Overall, our results are consistent with the predictions of standard theoretical models: Risk premia fluctuations relate to real-time, forward-looking information about second moments. In particular, individual perceptions of uncertainty about possible future inflation realizations are an important driver of bond risk premia. However, these correlations are changing over time, suggesting important differences across regimes. Our sample includes the period of high and volatile inflation expectations of the 1970s and very early 1980s, the long process of declining inflation expectations that followed, and the most recent period of low and stable inflation beliefs. In line with the evidence provided by David and Veronesi (2013) on the importance of changing beliefs about composite regimes for inflation uncertainty and its implications for bond risk properties, we find that inflation uncertainty matters most for bond risk premia during periods characterized by very high or very low inflation expectations accompanied by low-growth expectations.[4]

The ability to examine separately the impact of average individual uncertainty and interpersonal disagreement about inflation on bond risk premia has important implications for the assessment of alternative asset pricing models, for example, models with rational investors learning about composite regimes, versus those with heterogeneity either in the agent's reference models (e.g., Ulrich 2013) or in beliefs across agents (e.g., Buraschi and Whelan 2012) where disagreement matters for bond risk premia.[5] Our results also suggest that incorporating information from our measure of inflation uncertainty in the estimation of dynamic term-structure models may improve the identification of the inflation risk premium. For example, standard no-arbitrage affine term-structure models could be augmented in the spirit of Kim and Orphanides (2012), who incorporate information about first moments from surveys, to also include information from second moments.

Another appealing aspect of using survey data in the context of our study is that allows to obtain model-free measures of inflation uncertainty that are not derived from financial quotes as in Christensen, Lopez, and Rudebush (2012) and Kitsul and Wright (2013), and therefore are not affected by a financial risk premium adjustment. This adjustment would pollute our estimates of the link between inflation risk and bond risk premia as it is conceivable that it would comove with the risk premium contained in bonds, even if the underlying fundamental inflation risk priced by investors did not change.

The paper is organized in 6 sections. Following the introduction, sections 2 and 3 develop our ex-ante measure of perceived inflation uncertainty and describe its stylized facts over time. Section 4 documents the relation of inflation uncertainty with bond premia and section 5 examines in greater detail the link between inflation uncertainty and bond premia across alternative regimes. Concluding remarks appear in section 6.

[4] We establish these facts by first, interacting our measures of expected inflation, inflation uncertainty, and disagreement with the probability of a decline in real GDP in the following quarter from the SPF, and then analyzing the coefficients of the interacted variables across three different subsamples characterized by quite distinct composite regimes.

[5] Ulrich (2013) develops a term-structure model with Knightian inflation ambiguity, which he approximates with inflation disagreement from survey data based on the evidence provided by Patton and Timmermann (2010) showing that dispersion among forecasters cannot be explained by different information sets but mainly arises from heterogeneity in models.

## 2 Measuring ex-ante inflation uncertainty

The SPF presents a useful exception to the typical survey structure. Since its inception, this quarterly survey has asked respondents to provide probabilistic assessments of the outlook for inflation. Focusing on these probabilistic responses allows construction and comparison of aggregate measures of inflation uncertainty and disagreement. Starting with the important study by Zarnowitz and Lambros (1987) a number of authors have suggested various approaches to measure uncertainty as reflected in these survey responses.[6] In this paper, we build on the methodology originally developed in D'Amico and Orphanides (2008) (D&O) that extend earlier work and present various measures of inflation uncertainty and disagreement that correct some of the biases that might arise due to the survey's imperfections. (This is described in the appendix.) The methodology yields quarterly time series estimates of an approximate year-ahead inflation uncertainty and disagreement from 1968Q4 to 2013Q2. The availability of a long history of these measures allows us to study their dynamic over business cycles and relate them to bond risk premia, this relation is the primary focus of our study.

Following D&O, our starting point is the estimation of the characteristics of the probability densities for each forecaster in each period available in the survey. In particular, we begin our analysis with direct measures of individual uncertainty and disagreement that do not assume any specific continuous distributions for the probabilistic beliefs but are simply based on computing sample means and variances. Assuming that the underlying distribution for each forecaster is continuous, the discretization of the probabilistic responses in only a few intervals introduces an upward bias in the estimates of the individual variances that depends on the width of the intervals. To compensate for this bias, we apply the Sheppard's correction to variance estimates (see Kendall and Stuart (1977)).

To compute the mean, $\mu$, and standard deviation, $\sigma$, of the individual histograms we apply the following formulas, respectively:

$$\mu_{i,h,t} = \sum_{j=1}^{n} \frac{(u_{j,h,t} - l_{j,h,t})}{2} \cdot p_{i,j,h,t},$$

$$\sigma_{i,h,t} = \sqrt{\left[\sum_{j=1}^{n} \left(\frac{(u_{j,h,t} - l_{j,h,t})}{2} - \mu_{i,h,t}\right)^2 \cdot p_{i,j,h,t} - \frac{w_t^2}{12}\right]_+},$$

where $p_{i,j,h,t}$ is the probability that the forecaster $i$ assigns to the $j$th interval in the survey conducted at time $t$ with horizon $h$; $u_{j,h,t}$ and $l_{j,h,t}$ are the upper and lower limits of the $j$th interval; and $w_t$ is the width of the central interval, thus the term $\frac{w_t^2}{12}$ represents the Sheppard's correction for the second moment.

The individual means across all $N_t$ respondents in each quarter may be used to obtain summary characteristics of their cross sectional distribution. In particular, averaging across respondents provides an aggregate measure of the mean expectation of inflation for a specific quarter:

[6] Recent studies examining these density forecasts and their properties include, Andrade, Ghysel and Idier (2014), D'Amico and Orphanides (2008), Engelberg, Manski and Williams (2009), Giordani and Soderlind (2003), Lahiri and Liu (2004), Rich and Tracy (2010), and Wallis (2005).

$$\mu_{h,t} = \frac{1}{N_t} \sum_{i=1}^{N_t} \mu_{i,h,t};$$

The cross-sectional standard deviation of the individual means provides a measure of disagreement regarding the mean forecast:

$$d_{h,t} = \sqrt{\frac{1}{N_t} \sum_{i=1}^{N_t} \left(\mu_{i,h,t} - \mu_{h,t}\right)^2}.$$

Obtaining a reliable aggregate characterization of inflation uncertainty is not as simple as averaging across individual standard deviations, $\sigma_{i,h,t}$.[7] An important reason for this is the tendency of some survey respondents to round and concentrate their probabilistic responses in just a few bins, which introduces errors in the estimation. These errors appear more severe when respondents have to utilize larger intervals and the forecast horizon is short, suggesting that it may be exceedingly difficult to obtain an accurate measure when uncertainty is relatively small.

To mitigate this difficulty, D&O propose to model directly the distribution of the individual uncertainties and treat the uncertainties of forecasters that assign a positive probability only to one or two bins as small but unobserved**.** Specifically, let us suppose that the individual uncertainty in quarter $t$ for horizon $h$ originates from a distribution of the individual variances, $v_{i,h,t}$ :

$$v_{i,h,t} \tilde{} G(v, \theta_{h,t}).$$

Then the characteristics of this distribution, including the mean and variance of inflation uncertainty can be recovered from estimation of the parameters $\theta_{h,t}$. To the extent this distribution could be fit by treating the variances of responses with one or two bins as small but without knowledge of their precise values, biases arising from errors in the estimation with few bins could be avoided. This can be achieved by fitting a truncated distribution to the empirical CDF of the individual non-parametric variances ($v_{i,h,t} = \sigma^2_{i,h,t}$) that exceed a certain threshold, incorporating the information we have about the mass of the CDF at the threshold, but treating variances below that threshold as unobservable.

More precisely, given a threshold $C$, consider the individual variances above the threshold $v_{i,h,t}(c_\tau)$ for all $c_\tau > C$, where $c_1, ...c_{N_C}$ are the right endpoints of the intervals in which the range of uncertainty values has been discretized, and $F(v_{i,h,t}(c_1)), ...F(v_{i,h,t}(c_{N_C}))$ are the empirical CDF of $v_{i,h,t}(c_\tau)$ defined at these endpoints. Given a candidate distribution $G$, we can obtain the parameters $\theta$ that provide the best fit by minimizing:

$$\min_\theta \sum_{\tau=1}^{N_C} \left[G\left(v_{i,h,t}\left(c_\tau\right); \theta\right) - F\left(v_{i,h,t}\left(c_\tau\right)\right)\right]^2.$$

Since uncertainty cannot take negative values, the choice of candidate models is limited to distributions with non-negative support. We examined the Gamma, Chi-square and

[7] Often, even the median would not be the most appropriate measure since, as shown in Table 1A of the appendix, there are periods when more than 50 percent of respondents concentrate their answers in less than 3 bins.

Lognormal, but in the following we concentrate our attention on the Gamma, which provided the best fit among them.

Estimating the two parameters of the Gamma, $(\alpha, \beta)$, in each quarter, allows us to track the characteristics of the distribution of uncertainty over time. Since we are interested in average uncertainty expressed in standard deviation units, $\sigma_{i,h,t} = \sqrt{v_{i,h,t}}$ and we model the variances, $v_{i,h,t}$, to be Gamma distributed, we apply a change in variable to recover the mean of $\sigma_{i,h,t}$. If $v^G_{h,t} = E^G_{h,t}(v_{i,h,t}) = \widehat{\alpha}_{h,t} \cdot \widehat{\beta}_{h,t}$, where $E^G$ indicates the expected value under the Gamma distribution, then the mean of the cross-sectional distribution for the standard deviations may be obtained from:

$$\widehat{\sigma}^G_{h,t} = E^G_{h,t}(\sigma_{i,h,t}) = \widehat{\beta}^{1/2}_{h,t} \cdot \frac{\Gamma(\widehat{\alpha}_{h,t} + 1/2)}{\Gamma(\widehat{\alpha}_{h,t})}.$$

Fitting directly the distribution of the individual non-parametric uncertainties in this manner provides a way to summarize the quarterly uncertainty that improves upon taking a simple average of individual standard deviations by circumventing the problems associated with respondents with just one or two bins.

## 3 Stylized facts about inflation uncertainty

We present estimates of the quarterly time series of the mean and dispersion of expected inflation as well as the inflation uncertainty reflected in the subjective inflation probability distributions of the SPF. Figure 1 plots the three measures at the approximate one-year horizon over the whole 1968-2013 sample, along with the NBER recessions indicated by the shaded regions. Specifically, the top panel shows the mean inflation forecast, $\mu_{h,t}$, the middle panel shows the inflation uncertainty derived using the truncated Gamma distribution, $\widehat{\sigma}^G_{h,t}$, and the bottom panel shows the disagreement regarding the mean inflation forecast, $d_{h,t}$.

As illustrated in the middle panel, the ex-ante measure of inflation uncertainty rises sharply just before or during most recessions, with the last recession being a notable exception.[8] It increases even during recessions characterized by declining inflation expectations as in 1981-82, 1990-91 and 2001, suggesting that inflation falling below the expected central tendency for future inflation is also perceived as a risk.[9] The estimates suggest three episodes of particularly elevated uncertainty: 1973Q4-1975Q4 that includes the first oil shock of the 1970s, 1978Q4-1980Q2 that reflects the second oil shock and a period of low credibility by the Federal Reserve, and 1980Q4-1992Q1 that reflects the disinflation period under Fed Chairmen Volcker and Greenspan. It also reveals a long period of relatively lower uncertainty, from 1992Q2 until the end of the sample.

As can be seen by comparing the middle and bottom panels, inflation uncertainty and disagreement co-move, but their evolution differs substantially over time. Disagreement about the mean inflation forecast rose somewhat with inflation during the 1970s but has trended downward since then, although has increased somewhat during the last recession.

[8] However, this particular finding is broadly consistent with the estimated dynamics of beliefs about composite growth/inflation regimes in David and Veronesi (2013), showing that during the last recession the probability assigned to the low growth/zero inflation regime is about 90 percent suggesting almost certainty about a very extreme regime.

[9] Other studies such as Pfajfar and Zakelj(2011) have found similar evidence from experimantal data on the importance of the downside risk to inflation.

This simple graphical comparison would suggest that dispersion of beliefs about inflation is not a good measure of ex-ante inflation uncertainty, as it would be expected considering that not only these are two quite different economic concepts but also because other forecasters' expectations are not observable in real time.[10]

Next, we analyze the simple correlations between the mean inflation forecast, inflation disagreement and uncertainty, as well as their correlations with other commonly-used ex-post measures of inflation/macroeconomic risk. This analysis is useful to understand whether widely-used ex-post measures of macroeconomic risk are a good proxy of ex-ante inflation uncertainty or macroeconomic uncertainty more in general. If not, then maybe it should not be surprising that they are not found to be strong predictors of risk premia in financial markets, as they would not seem to reflect the macroeconomic uncertainty that investors are trying to price in real-time.

Table 1 shows the correlation matrix, which in addition to $\mu_{h,t}$, $\widehat{\sigma}^{G}_{h,t}$, and $d_{h,t}$, also includes the following variables: realized inflation volatility estimated with a GARCH (1,1), one-year-ahead SPF inflation forecast errors computed using the final releases of inflation, one-year-ahead SPF inflation forecast errors computed using the advance releases of inflation (see Stark 2010 for more details on real-time forecast evaluation of the SPF), Real GDP disagreement measured by the dispersion (inter-quantile range) of one-year-ahead Real GDP point forecasts in the SPF, the VXO stock market volatility index implied by one-month option contracts on the S&P100 index used by Bloom (2009) as proxy of macro uncertainty, and the broad macroeconomic uncertainty factor that Jurado, Ludvigson, and Ng (2013) obtain as the principal component of the forecast error variances for 279 macro and financial indicators.[11] It should be stressed that all the measures of disagreement and forecast accuracy are not observable in real time, as they can not be computed until either the other panelists return their projections and the survey results are released or the advance estimates and various revisions of the data become publicly available. Finally, for all the series that are available at a higher frequency than quarterly, we have used the average of the data over the first month of each quarter, as this would roughly corresponds to the observations available to the forecasters when they generated their projections for the SPF, making the comparison across various measures more accurate. Indeed, the questionnaire is mailed to the panelists at the end of the first month and has to be returned within few days.

All correlations with inflation uncertainty ($\widehat{\sigma}^{G}_{h,t}$) are positive but none exceeds 0.43, which is the correlation with the mean inflation forecast. In particular, the correlation between inflation disagreement and inflation uncertainty is only 0.34, suggesting, in line with the graphical analysis, that inflation disagreement may not be a good proxy for inflation uncertainty. Similarly, the correlations with all other ex-post measures of macroeconomic risks are quite low. In contrast, the correlation between the inflation disagreement and mean inflation forecast is 0.60, suggesting generally greater disagreement about the inflation outlook when inflation is expected to be higher. And also the correlation between inflation disagreement and Real GDP disagreement is about 0.60, indicating that dispersion of beliefs about macro fundamentals seem to move together. Finally, it should be noted that some of

[10] This was already clearly pointed out in Zarnowitz and Lambros (1987), D'Amico and Orphanides (2008), and Rich and Tracy (2010).

[11] Before 1986, since option contracts on the S&P100 Index are not available, Bloom (2009) relies on the realized volatility of the actual S&P100 Index returns. And the broad macroeconomic uncertainty factor is available from 1960 to 2011on Sydney C. Ludvigson's website: http://www.econ.nyu.edu/user/ludvigsons/.

the extremely low or even negative correlations can be due to the shorter-term horizon of the risk measure.

# 4 Relation with bond risk premia

In this section, we examine the relation of bond premia with both our ex-ante measure of inflation uncertainty and the disagreement regarding inflation forecasts. We first obtain various measures of term and risk premia in the U.S. Treasury and corporate bond markets going back to 1968, and then ran regressions of these bond premia on the inflation uncertainty and disagreement estimates described in Section 2, while also controlling for the mean inflation forecast. This allows us to examine more formally the direct link between bond risk premia, ex-ante expected inflation and inflation uncertainty, as well as dispersion of beliefs about inflation, within a unified and consistent framework, as all three attributes of the inflation outlook are derived from the same subjective inflation probability distributions. Further, since the SPF forecasters come largely from the business world and Wall Street, they may be paying particular attention to the evaluation and pricing of macroeconomic risks.[12]

For ease of comparison to Wright (2011), our first measure of Treasury bond term premium is the 10-year ahead forward term premium ($10yFTP$) obtained from an arbitrage-free three-factor model of the term-structure augmented with survey data on short-term rates (see Kim and Orphanides, 2012, and Kim and Wright, 2005).[13] The second measure of the Treasury term-premium is the spread between the 10- and 7-year Treasury forward rates ($10y{-}7yFS$) obtained from the Gurkaynak, Sack, and Wright (2007) (GSW) dataset.[14] We include this alternative measure of the term-premium because is model-free, being based on the simple observation that at distant horizons the expectation components of forward rates should be very similar in magnitude and therefore the spread between those rates should mainly reflect the term-premium component of the 10-year forward rate.

However, since Treasury term premium estimates are quite persistent and we also want to facilitate comparison to previous studies on the predictability of bond risk premia (e.g., Ludvigson and Ng, 2009, and Buraschi and Whelan, 2012), we compute the standard excess returns on the 2-, 5-, and 7-year maturity Treasury bonds and focus on one-year holding period returns such that the length of the investment horizon is about the same as the length of the inflation forecasting horizon. In particular, let $p_t^n$ be the log price of the n-year zero-coupon bond, then the yield of the bond is defined as $y_t^n = -\frac{1}{n}p_t^n$. The one-year holding period return is the realized return on an $n-$year maturity bond bought at date $t$ and sold as an $(n-1)$-year maturity bond at date $t+1$:

$$r_{t,t+1}^n = p_{t+1}^{n-1} - p_t^n,$$

and the excess holding period returns are obtained by subtracting the yield on the one-year bond:

[12] As reported in the SPF documentation, among the survey participants about 1/3 are Wall Street financial firms, the second largest group are banks and economic consulting firms, and then there are a few university research centers and chief economists at many Fortune 500 companies.

[13] We are grateful to the staff of the Monetary and Financial Market Analysis section at the Federal Reserve Board for having extended the estimates of the Treasury term premium back to 1960.

[14] The GSW dataset contains the 10-year forward rate starting in 1973.

$$rx_{t,t+1}^n = r_{t,t+1}^n - y_t^1.$$

The excess holding period returns are also derived from the Treasury zero-coupon bond yields dataset of GSW.

Finally, to verify the robustness of our investigation for other fixed-income securities, we also use standard measures of corporate bond risk premia such as the spread between the 10-year BAA- and AAA-rated corporate bond yields and the excess bond premium (EBP) of Gilchrist and Zakrajsek (2012).

Our first set of regressions are based on projections of excess bond returns at a one-year horizon for different maturity bonds on the mean inflation forecast, $\mu_{h,t}$, inflation uncertainty, $\widehat{\sigma}_{h,t}^G$, and disagreement regarding the mean inflation forecast, $d_{h,t}$:

$$rx_{t,t+1}^n = \alpha + \beta_1 \mu_{1,t} + \beta_2 \widehat{\sigma}_{1,t}^G + \beta_3 d_{1,t} + \epsilon_t.$$

We also run regressions of the same form but using as dependent variables either the Treasury term premia or corporate bond risk premia described above. Further, since all the financial variables we use are available at daily or monthly frequency, the quarterly time-series for each premium selects the premium prevailing on average on the last month of each quarter, so the premia are always measured after the SPF release, which occurs in the middle month of each quarter.

Addressing our main empirical question amounts to asking whether the coefficient $\beta_2$ is positive, statistically significant, and has an explanatory power larger than $\beta_3$. Table 2 summarizes the results for our baseline specification over the whole sample period, and Newey and West (1987) corrected t-statistics with a lag order of 4 quarters are reported in parentheses. Coefficients that are significant at the 5% or better level are highlighted in bold. As can be seen, the coefficient on the inflation uncertainty is consistently significant and positively related to all measures of bond risk and term premia, while the coefficient on inflation disagreement is not. This indicates that in this sample and differently from previous studies, when a measure of ex-ante inflation uncertainty is included in the regression, the inflation disagreement loses its marginal predictive power. In contrast, the mean inflation forecast stays mostly statistically significant, with the exception of the corporate bond premia specifications. Further, the magnitude of $\beta_2$ increases with the bond maturity. The $R^2$ statistics indicate that these attributes of the inflation forecast explain 10 to 13 percent of next year's excess returns on Treasury bonds and up to about 20 percent of excess returns on corporate bonds, in line with previous studies on the predictive power of macroeconomic disagreement for bond risk premia (see Buraschi and Whelan, 2012). The $R^2$ statistics are notably higher for the Treasury term premia, varying between 27 and 50 percent.

To verify the stability of our estimates, we also report the results after excluding the most recent financial crisis. As shown in Table 3, the coefficients' size and statistical significance is hardly affected by ending the sample period in 2007Q4. However, the $R^2$ for the corporate bond risk premia improve considerably.

## 4.1 Is it the inflation uncertainty that really matters?

To gain further insights about the economic determinants of bond risk premia and investigate whether inflation uncertainty contains substantial information above and beyond that contained in broad indicators of macroeconomic activity in general, and of macroeconomic disagreement more specifically, we include variables that prior works found to have strong predictive power for excess bond returns. In particular, we control for the macroeconomic factors of Ludvigson and Ng (2009) (L&N) and the proxies for disagreement about real economic activity similar to Buraschi and Whelan (2012) (B&W).[15]

L&N is the first study that finds strong predictable variation in excess bond returns that is associated to macroeconomic activity, which underscores the importance of including macro factors to explain the cyclical behavior of risk premia. Specifically, they use dynamic factor analysis to estimate common factors from a monthly panel of 132 measures of economic activity. It is evident that while their focus is on capturing broad variation in actual realizations of macroeconomic fundamentals, our focus is on building controls for uncertainty about macroeconomic forecasts and verify whether they contain additional information. To this end, using their same notation, we augment our baseline specification first with $f_1$, the estimated factor that loads heavily on measures of employment and real output, and then with $F5$, which is constructed as the linear combination of five individual factors, including $f_1$, two inflation factors, and a factor that loads heavily on aggregate measures of the stock market. The results of this additional set of regressions are shown in Tables 4.

In particular, the top panel of Table 4 shows the results for our baseline regressions but over the same sample period of L&N that ends in 2003, and the middle panel shows the baseline specification augmented with $f_1$. Again, coefficients that are significant at the 5% or better level are highlighted in bold. As can be seen, not only the coefficients on inflation uncertainty remain statistically significant and are little changed in magnitude, but the estimated factor summarizing information on employment and real output, differently from L&N, loses its marginal explanatory power.

The bottom panel shows the results for the regression specifications including $F5$, which indicate that, even when we control for a broad macroeconomic indicator that includes information from a large number of economic variables, inflation uncertainty remains significant and the size of its coefficients is only slightly smaller. This confirms that our ex-ante measure of inflation risk is not just proxying for business cycle fluctuations but contains additional information that seems separately priced in bond risk premia. It is also interesting to note that, even though $F5$ contains the individual factors that load most heavily on measures of inflation and price pressures, the mean inflation forecast mostly retains its marginal explanatory power. This is perhaps not surprising, as real-time measures of one-year-ahead expected inflation can convey information quite different from that provided by actual realizations of past inflation, and therefore useful in predicting inherently forward-looking variables such as asset prices.

Next, similarly to B&W, who find that disagreement about the real economy (measured by the cross-sectional mean-absolute-deviation in Real GDP growth forecasts) loads positively on excess bond returns, is statistically significant, and is more important than disagreement about inflation, we add to the baseline specification the disagreement about one-year-ahead Real GDP forecasts from the SPF, measured by the inter-quantile range of the point forecasts. The results from these regressions are summarized in Table 5. It can be noted that also in this case, the inflation uncertainty preserves its marginal predictive power and the magnitude of its coefficients is hardly changed, while the Real GDP dis-

[15] The disagreement proxies are similar but not identical because ours are computed from the SPF while theirs are based on the monthly Blue Chip survey data.

agreement is not statistically significant and its coefficients have a counterintuitive negative sign. However, once we control for disagreement about the real economy, the mean inflation forecast loses its marginal explanatory power.

### 4.2 Additional Controls: the Broad Uncertainty Factor and Inflation Forecast Errors.

As additional robustness check, we also include in our regression specifications the broad macroeconomic uncertainty factor that Jurado, Ludvigson, and Ng (2013) obtain as the principal component of the forecast error variances for 279 macro and financial indicators, and the one-year-ahead SPF inflation forecast errors computed as the average of the 4-quarter-ahead individual squared forecast errors for each quarter. These two variables are reasonable alternative measures of broad macroeconomic and inflation risk, respectively, as they both proxy for the unforcastable component of future values of macroeconomic series. However, differently from our measure of inflation uncertainty, they are not directly observable in real time.[16] For this reason, it is important to verify whether ex-ante inflation uncertainty retains its statistical significance once one controls for these important proxy of macroeconomic risks. Table 6 reports the results for the baseline regressions augmented with the one-year-ahead broad macroeconomic uncertainty factor ($U(4q)$) and one-year-ahead SPF inflation forecast errors ($InflationF.E.$), respectively.

There are few notable features of these results: (i) Inflation uncertainty preserves its explanatory power as it stays consistently significant and positively related to bond risk premia. (ii) The broad macroeconomic uncertainty factor ($U(4q)$) is not statistically significant indicating that it does not contain additional information for bond risk premia relative to the ex-ante inflation uncertainty. (This result holds for different horizons of $U(\cdot)$, not reported). (iii) The SPF inflation forecast errors are not statistically significant. This is not surprising since these forecast errors are only observed ex-post, and thus represent a rather noisy indicator of uncertainty compared to the ex-ante uncertainty measure captured in $\widehat{\sigma}^G_{h,t}$.

To summarize, so far we have shown that in the full sample period ( either 1968-2013 or 1968-2003), inflation uncertainty is consistently significant and positively related to bond risk premia, and the magnitude of its coefficients increases with the bond maturity. In contrast, the dispersion of inflation forecasts is never significant once we control for mean inflation forecast and inflation uncertainty. In addition, these results are robust to the inclusion of a number of controls such as the L&N macroeconomic factors, the disagreement about Real GDP that B&W found to be significant and more important than inflation disagreement, and alternative measures of inflation and broad macroeconomic uncertainty.

## 5 Does the importance of inflation uncertainty for bond risk premia change across regimes?

David and Veronesi (2013) (D&V) identify a novel mechanism that help explaining the importance of inflation uncertainty for bond risk properties, that is, the time-varying signaling role of inflation. Specifically, in a general equilibrium model in which agents learn about

[16] Obviously forecasters can observe the realization of their forecast errors only once the actual data are released.

composite growth and inflation regimes, inflation surprises (i.e., inflation news above or below expectations) may signal bad or good times depending on current beliefs about the composite regime, generating very different implications for bond risks, measured by bond return volatility and their covariance with stock returns. In particular, they show that uncertainty between a low growth/low inflation regime and a low growth/high inflation regime is extremely important for bond risk, as this inflation uncertainty generates both very high bond return volatility and a positive stock-bond covariance making bonds very poor hedges during recessions.

To account for variations in beliefs about these composite regimes, we interact our measures of mean inflation forecast, inflation uncertainty, and disagreement, with the probability of decline in real GDP in the following quarter, obtained from the SPF. We use this variable because D&V show that is highly correlated with the marginal probability of the low-growth regime implied by their model, and therefore interacted with our ex-ante measures of expected inflation and inflation uncertainty should capture quite well the changes in the composite beliefs about growth and inflation. Specifically, we run the following set of regressions:

$$rx^n_{t,t+1} = \alpha + \beta_1 \mu_{1,t} P_L + \beta_2 \mu_{1,t} P_H + \beta_3 \widehat{\sigma}^G_{1,t} P_L + \beta_4 \widehat{\sigma}^G_{1,t} P_H + \beta_5 d_{1,t} P_L + \beta_6 d_{1,t} P_H + \epsilon_t,$$

where $P_L$ is the probability of decline in real GDP in the following quarter from the SPF, and $P_H = (1 - P_L)$. Table 7 reports the regression results for the whole 1968-2013 sample period. As can be seen, inflation uncertainty interacted with the probability of low growth enters positively and is the only statistically significant explanatory variable. Compared to the baseline specification results reported in Table 2, using the interacted variables improves the $R^2$ statistics, indicating that this is a better specification. This finding is consistent with the predictions of the D&V study. Estimates of their model parameters reveal that the two extreme inflation regimes, very high and very low, only occur with low economic growth in the data, and more importantly, inflation uncertainty varies the most when inflation beliefs transition into these composite regimes, becoming a very strong driver of bond risks.[17]

To investigate further the impact of extreme inflation regimes accompanied by low growth on inflation uncertainty and its relation with bond risk premia, we choose to break our sample into three subperiods: 1968Q4-1981Q4, which is mainly dominated by very high inflation and frequent recessions, 1982Q1-1996Q4, which is mostly characterized by medium inflation and medium growth, and 1997Q1-2013Q2, during which inflation is remarkably low, with one-year expected inflation consistently below 3%, and growth is quite low as well, especially due to the last two recessions. Results for the regressions estimated in the three subsamples are reported in Table 8. As can be seen from the top panel, in the first subperiod, inflation uncertainty enters positively, is highly statistically significant, and the coefficient is increasing with the bond maturity. The $R^2$ for these regressions ranges between 27 and 33 percent, indicating an economically large degree of predictability of future bond returns, especially for longer-maturity bonds. In contrast, in the second subperiod (middle panel) none of the variables is statistically significant. Inflation uncertainty becomes a statistically significant driver again in the third subsample. Interestingly, inflation disagreement is also significant in this case.

[17] In their simulations, inflation uncertainty is positively related to the conditional covariance of stock and bond returns and to the bond return volatility and is their strongest driver, as by itself explains over 72% and 86% of their variation, respectively.

# 6 Concluding Remarks

This paper provides evidence of a direct link between variations in perceived inflation uncertainty and bond risk premia. Using the subjective inflation probability distributions available in the Survey of Professional Forecasters we construct a quarterly time series of the average individual uncertainty about inflation forecasts since 1968.

Consistent with the predictions of standard theoretical models we document that risk premia fluctuations relate to real-time, forward-looking information about second moments. Our survey-based measure of uncertainty about possible future inflation realizations is an important driver of bond risk premia. This uncertainty measure differs importantly from measures of disagreement regarding inflation forecasts and other proxies, such as model-based ex-post measures of macroeconomic risk. Inflation uncertainty is an important driver of bond premia in the overall sample, but the relation varies across inflation regimes. It is most important in the high inflation regime early in the sample and the low inflation regime over the last 15 years, lending support to the regime-dependent role of inflation uncertainty for bond risk properties suggested by David and Veronesi (2013).

Our results suggest that incorporating survey-based measures of inflation uncertainty may also help the assessment of alternative asset pricing models. Using survey data to distinguish between uncertainty and disagreement of point forecasts can help differentiate among competing theories that examine the possible impact of learning about composite regimes and heterogeneity in investor beliefs. In addition, including information from our measure of inflation uncertainty may improve the estimation of dynamic term-structure models that already include information from first moments of survey expectations but not second moments.

Overall, efforts toward the challenging task of extracting accurate ex-ante measures of perceived macroeconomic risk from survey data appear to be fruitful for our understanding of subjective risk assessment and asset pricing, facilitating the continuing progress in research at the crossroads of macroeconomics and finance.

Table 1
Correlations with ex-post measures of uncertainty

| | $\widehat{\sigma}^G_{1y,t}$ | $\mu_{1y,t}$ | $d_{1y,t}$ |
|---|---|---|---|
| 1-year SPF Inflation Uncertainty ($\widehat{\sigma}^G_{1y,t}$) | 1.00 | | |
| 1-year SPF Mean Inflation Forecast ($\mu_{1y,t}$) | 0.43 | 1.00 | |
| 1-year SPF Inflation Disagreement ($d_{1y,t}$) | 0.34 | 0.60 | 1.00 |
| Inflation Conditional Volatility GARCH (1,1) | 0.00 | −0.04 | 0.12 |
| 1-year SPF Inflation Forecast Errors (final) | 0.41 | 0.33 | 0.54 |
| 1-year SPF Inflation Forecast Errors (advance) | 0.31 | 0.58 | 0.53 |
| 1-year SPF Real GDP Disagreement | 0.21 | 0.76 | 0.62 |
| 1-month VXO Index | 0.07 | −0.06 | 0.02 |
| 1-year Broad Uncertainty Factor | 0.24 | 0.44 | 0.30 |

Notes: One-year-ahead SPF inflation forecast errors are computed as the average of the 4-quarter-ahead individual squared forecast errors for each quarter. One-year-ahead Real GDP disagreement is measured by the inter-quantile range of 4-quarter-ahead Real GDP point forecasts in the SPF. Regarding the VXO Index, before 1986, since option contracts on the S&P100 Index are not available, Bloom (2009) relies on the realized volatility of the actual S&P100 Index returns. The broad macroeconomic uncertainty factor is obtained by Jurado, Ludvigson, and Ng (2013) as the principal component of the forecast error variances for 279 macro and financial indicators, and is available from 1960 to 2011.

Table 2
Relation with bond risk and term premia in the full sample

| Market | | Sample 1968Q4:2013Q2 $\mu_{1y,t}$ | $\widehat{\sigma}^G_{1y,t}$ | $d_{1y,t}$ | $R^2$ |
|---|---|---|---|---|---|
| Treasury | $rx^2_{t,t+1}$ | $-0.22$ | **4.69** | $-0.40$ | 0.10 |
| | | $(-1.66)$ | (2.81) | $(-0.55)$ | |
| | $rx^5_{t,t+1}$ | $\mathbf{-0.96}$ | **12.90** | $-1.06$ | 0.12 |
| | | $(-2.51)$ | (2.59) | $(-0.43)$ | |
| | $rx^7_{t,t+1}$ | $\mathbf{-1.38}$ | **16.79** | $-1.51$ | 0.13 |
| | | $(-2.67)$ | (2.44) | $(-0.44)$ | |
| | $10yFTP_{t+1}$ | **0.23** | **4.94** | $-0.27$ | 0.50 |
| | | (3.48) | (4.85) | $(-0.61)$ | |
| | $10y - 7yFS^*_{t+1}$ | **0.05** | **0.46** | $-0.08$ | 0.27 |
| | | (3.08) | (2.37) | $(-0.52)$ | |
| Corporate | $10yBAA - AAA_{t+1}$ | 0.04 | **0.94** | 0.20 | 0.19 |
| | | (1.11) | (2.40) | (0.95) | |
| | $EBP^*_{t+1}$ | $-0.01$ | **0.68** | 0.38 | 0.03 |
| | | $(-0.42)$ | (2.18) | (1.48) | |

Notes: The table reports estimates from OLS regressions of Treasury excess bond returns, Treasury term premia, and corporate bond premia, on the variables named in row 1, over the whole 1968-2013 sample period. Newey and West (1987) corrected t-statistics have a lag order of 4 quarters and are reported in parentheses. Coefficients that are significant at the 5% or better level are highlighted in bold. And coefficients for the variables marked by (*) are estimated starting in 1973 because of data availability. The regressions include also quarterly dummy variables (not shown) to control for possible seasonality, however, results obtained without these dummy variables are practically identical.

Table 3
Relation with bond risk and term premia after excluding the most recent financial crisis

| Market | | Sample 1968Q4:2007Q4 $\mu_{1y,t}$ | $\widehat{\sigma}^{G}_{1y,t}$ | $d_{1y,t}$ | $R^2$ |
|---|---|---|---|---|---|
| Treasury | $rx^{2}_{t,t+1}$ | −0.22 | **4.68** | −0.49 | 0.11 |
| | | (−1.52) | (2.81) | (−0.66) | |
| | $rx^{5}_{t,t+1}$ | **−0.88** | **12.87** | −1.72 | 0.13 |
| | | (−2.12) | (2.59) | (−0.71) | |
| | $rx^{7}_{t,t+1}$ | **−1.25** | **16.68** | −2.59 | 0.13 |
| | | (−2.24) | (2.43) | (−0.75) | |
| | $10yFTP$ | **0.19** | **4.78** | −0.40 | 0.47 |
| | | (2.82) | (4.75) | (−0.95) | |
| | $10y - 7yFS^{*}$ | **0.04** | **0.40** | −0.10 | 0.23 |
| | | (2.45) | (2.30) | (−0.69) | |
| Corporate | $10yBAA - AAA$ | **0.07** | **1.09** | 0.08 | 0.44 |
| | | (2.83) | (3.03) | (0.60) | |
| | $EBP^{*}$ | 0.01 | **0.85** | 0.16 | 0.12 |
| | | (0.42) | (2.95) | (0.88) | |

Notes: The table reports estimates from OLS regressions of Treasury excess bond returns, Treasury term premia, and corporate bond premia, on the variables named in row 1 over the 1968-2007 sample period, in order to exclude the most recent financial crisis. Newey and West (1987) corrected t-statistics have a lag order of 4 quarters and are reported in parentheses. Coefficients that are significant at the 5% or better level are highlighted in bold. And coefficients for the variables marked by (*) are estimated starting in 1973 because of data availability. The regressions include also quarterly dummy variables (not shown) to control for possible seasonality, however, results obtained without these dummy variables are practically identical.

Table 4
Relation with bond risk premia while controlling for macroeconomic factors

| | Sample 1968Q4:2003Q4 | | | | | |
|---|---|---|---|---|---|---|
| | $\mu_{1y,t}$ | $\widehat{\sigma}^{G}_{1y,t}$ | $d_{1y,t}$ | $f_{1t}$ | $F5_t$ | $R^2$ |
| $rx^2_{t,t+1}$ | $-0.23$ | **4.67** | $-0.69$ | | | 0.12 |
| | $(-1.59)$ | (2.76) | $(-0.94)$ | | | |
| $rx^5_{t,t+1}$ | $\mathbf{-0.90}$ | **12.84** | $-2.33$ | | | 0.14 |
| | $(-2.10)$ | (2.55) | $(-0.96)$ | | | |
| $rx^7_{t,t+1}$ | $\mathbf{-1.26}$ | **16.61** | $-3.44$ | | | 0.14 |
| | $(-2.18)$ | (2.38) | $(-1.00)$ | | | |
| $rx^2_{t,t+1}$ | $-0.26$ | **4.20** | $-0.78$ | $-0.40$ | | 0.16 |
| | $(-1.90)$ | (2.57) | $(-1.17)$ | $(-1.54)$ | | |
| $rx^5_{t,t+1}$ | $\mathbf{-0.95}$ | **11.92** | $-2.51$ | $-0.80$ | | 0.15 |
| | $(-2.34)$ | (2.35) | $(-1.08)$ | $(-1.14)$ | | |
| $rx^7_{t,t+1}$ | $\mathbf{-1.31}$ | **15.62** | $-3.63$ | $-0.86$ | | 0.14 |
| | $(-2.38)$ | (2.21) | $(-1.09)$ | $(-0.95)$ | | |
| $rx^2_{t,t+1}$ | $-0.24$ | **3.83** | $-0.33$ | | **0.41** | 0.22 |
| | $(-1.75)$ | (2.61) | $(-0.54)$ | | (3.62) | |
| $rx^5_{t,t+1}$ | $\mathbf{-0.91}$ | **10.92** | $-1.51$ | | **0.94** | 0.19 |
| | $(-2.20)$ | (2.28) | $(-0.67)$ | | (2.76) | |
| $rx^7_{t,t+1}$ | $\mathbf{-1.27}$ | **14.31** | $-2.45$ | | **1.12** | 0.18 |
| | $(-2.26)$ | (2.10) | $(-0.76)$ | | (2.44) | |

Notes: The table reports estimates from OLS regressions of excess bond returns on the variables named in row 1 over the 1968-2003 sample period in order to compare our results to those obtained by Ludvigson and Ng (2009), whose macroeconomic factors are included in the regression specifications. Specifically, $f_{1t}$ is the estimated factor that loads heavily on measures of employment and real output, and $F5_t$ is constructed as the linear combination of five individual factors, including $f_{1t}$, two inflation factors, and a factor that loads heavily on aggregate measures of the stock market. Newey and West (1987) corrected t-statistics have a lag order of 4 quarters and are reported in parentheses. Coefficients that are significant at the 5% or better level are highlighted in bold.

Table 5

Relation with bond risk premia while controlling for Real GDP disagreement

| | Sample 1968Q4:2013Q2 | | | | |
|---|---|---|---|---|---|
| | $\mu_{1y,t}$ | $\widehat{\sigma}^{G}_{1y,t}$ | $d_{1y,t}$ | $RGDPDisagr.$ | $R^2$ |
| $rx^2_{t,t+1}$ | $-0.09$ | **4.25** | 0.38 | $-0.68$ | 0.15 |
| | $(-0.52)$ | (2.72) | (0.49) | $(-1.79)$ | |
| $rx^5_{t,t+1}$ | $-0.65$ | **11.70** | 1.28 | $-1.72$ | 0.16 |
| | $(-1.32)$ | (2.45) | (0.49) | $(-1.65)$ | |
| $rx^7_{t,t+1}$ | $-0.97$ | **15.14** | 1.76 | $-2.27$ | 0.16 |
| | $(-1.47)$ | (2.29) | (0.48) | $(-1.64)$ | |

Notes: The table reports estimates from OLS regressions of excess bond returns on the variables named in row 1 over the 1968-2013 sample period. Specifically, Real GDP disagreement is measured by the inter-quantile range of 4-quarter-ahead Real GDP point forecasts. Newey and West (1987) corrected t-statistics have a lag order of 4 quarters and are reported in parentheses. Coefficients that are significant at the 5% or better level are highlighted in bold.

Table 6
Relation with bond risk premia while controlling for alternative measures of broad macro and inflation uncertainty

| | Sample 1968Q4:2011Q4 | | | | | |
|---|---|---|---|---|---|---|
| | $\mu_{1y,t}$ | $\widehat{\sigma}^G_{1y,t}$ | $d_{1y,t}$ | $U(4q)$ | $Inflation F.E.$ | $R^2$ |
| $rx^2_{t,t+1}$ | **−0.26** | **4.56** | −0.50 | 3.01 | | 0.12 |
| | (−1.96) | (2.78) | (−0.66) | (0.73) | | |
| $rx^5_{t,t+1}$ | **−1.03** | **12.49** | −1.56 | 5.41 | | 0.14 |
| | (−2.87) | (2.52) | (−0.63) | (0.46) | | |
| $rx^7_{t,t+1}$ | **−1.44** | **16.23** | −2.34 | 4.09 | | 0.15 |
| | (−3.03) | (2.37) | (−0.67) | (0.26) | | |
| | | | | | | |
| $rx^2_{t,t+1}$ | −0.21 | **3.89** | −1.31 | | 0.93 | 0.14 |
| | (−1.63) | (2.36) | (−1.54) | | (1.87) | |
| $rx^5_{t,t+1}$ | **−0.92** | **10.82** | −3.43 | | 2.38 | 0.15 |
| | (−2.54) | (2.25) | (−1.21) | | (1.55) | |
| $rx^7_{t,t+1}$ | **−1.32** | **14.17** | −4.50 | | 2.97 | 0.15 |
| | (−2.70) | (2.15) | (−1.14) | | (1.41) | |

Notes: The table reports estimates from OLS regressions of excess bond returns on the variables named in row 1 over the 1968-2011 sample period because the broad uncertainty factor is available until 2011. Specifically, the broad macroeconomic uncertainty factor is obtained by Jurado, Ludvigson, and Ng (2013) as the principal component of the forecast error variances for 279 macro and financial indicators. One-year-ahead SPF inflation forecast errors are computed as the average of the 4-quarter-ahead individual squared forecast errors for each quarter. Newey and West (1987) corrected t-statistics have a lag order of 4 quarters and are reported in parentheses. Coefficients that are significant at the 5% or better level are highlighted in bold.

Table 7
Relation with bond risk premia interacting with probability that real GDP will decline in the quarter

| | Sample 1968Q4:2013Q2 | | | | | | |
|---|---|---|---|---|---|---|---|
| | $\mu_{1y,t}P_L$ | $\mu_{1y,t}P_H$ | $\widehat{\sigma}^G_{1y,t}P_L$ | $\widehat{\sigma}^G_{1y,t}P_H$ | $d_{1y,t}P_L$ | $d_{1y,t}P_H$ | $R^2$ |
| $rx^2_{t,t+1}$ | $-0.54$ | $-0.17$ | **8.96** | 3.03 | $-1.32$ | $-0.36$ | 0.18 |
| | $(-1.62)$ | $(-0.95)$ | (2.83) | (1.86) | $(-0.56)$ | $(-0.39)$ | |
| $rx^5_{t,t+1}$ | $-1.86$ | $-0.77$ | **23.44** | 8.84 | $-2.64$ | $-1.29$ | 0.17 |
| | $(-1.93)$ | $(-1.64)$ | (2.51) | (1.67) | $(-0.37)$ | $(-0.41)$ | |
| $rx^7_{t,t+1}$ | $-2.32$ | $-1.20$ | **28.45** | 12.26 | $-3.53$ | $-1.65$ | 0.16 |
| | $(-1.80)$ | $(-1.93)$ | (2.27) | (1.65) | $(-0.37)$ | $(-0.36)$ | |

Notes: The table reports estimates from OLS regressions of excess bond returns on the interacted variables named in row 1 over the 1968-2013 sample period. $P_L$ is the probability of decline in real GDP in the following quarter from the SPF, and $P_H = 1 - P_L$. Newey and West (1987) corrected t-statistics have a lag order of 4 quarters and are reported in parentheses. Coefficients that are significant at the 5% or better level are highlighted in bold.

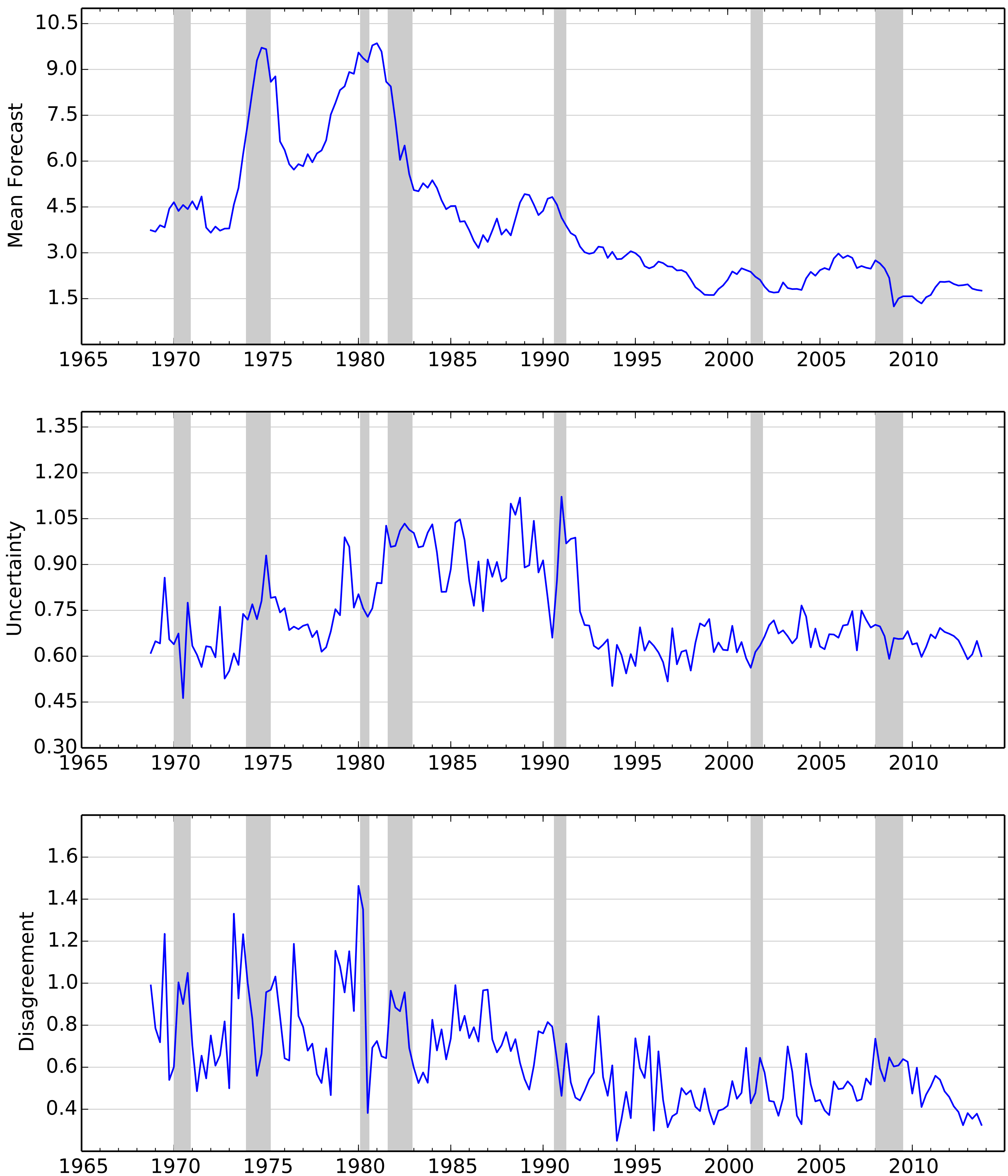
Table 8

Relation with bond risk premia interacting with probability that real GDP will decline in the following quarter across three different regimes

| | Three Sub-Periods | | | | | | |
|---|---|---|---|---|---|---|---|
| | $\mu_{1y,t}P_L$ | $\mu_{1y,t}P_H$ | $\widehat{\sigma}^G_{1y,t}P_L$ | $\widehat{\sigma}^G_{1y,t}P_H$ | $d_{1y,t}P_L$ | $d_{1y,t}P_H$ | $R^2$ |
| 68Q4:81Q4 | | | | | | | |
| $rx^2_{t,t+1}$ | **−2.02** | 0.03 | **25.65** | −0.34 | −3.65 | −1.01 | 0.27 |
| | (−3.55) | (0.10) | (5.02) | (−0.09) | (−1.63) | (−0.72) | |
| $rx^5_{t,t+1}$ | **−6.11** | 0.08 | **70.24** | −3.30 | −9.52 | −3.68 | 0.32 |
| | (−3.81) | (0.11) | (4.79) | (−0.37) | (−1.30) | (−1.19) | |
| $rx^7_{t,t+1}$ | **−8.11** | 0.09 | **91.96** | −5.17 | −12.96 | −5.18 | 0.33 |
| | (−3.77) | (0.10) | (4.35) | (−0.44) | (−1.24) | (−1.34) | |
| 82Q1:96Q4 | | | | | | | |
| $rx^2_{t,t+1}$ | 0.76 | 0.57 | 2.03 | −1.00 | −1.47 | 0.74 | 0.15 |
| | (0.62) | (1.03) | (1.24) | (−0.40) | (−0.15) | (0.38) | |
| $rx^5_{t,t+1}$ | 2.93 | 0.51 | −1.16 | 1.98 | −3.32 | 0.98 | 0.00 |
| | (0.82) | (0.25) | (−0.20) | (0.24) | (−0.11) | (0.13) | |
| $rx^7_{t,t+1}$ | 4.72 | 0.02 | −6.88 | 4.81 | −6.36 | 1.06 | 0.00 |
| | (1.07) | (0.01) | (−0.84) | (0.44) | (−0.18) | (0.10) | |
| 97Q1:13Q2 | | | | | | | |
| $rx^2_{t,t+1}$ | **2.10** | 0.47 | **18.93** | **−5.92** | **−23.99** | **5.20** | 0.21 |
| | (2.62) | (0.79) | (3.27) | (−2.16) | (−3.96) | (3.29) | |
| $rx^5_{t,t+1}$ | 4.42 | 0.10 | **58.11** | −16.47 | **−68.36** | **21.90** | 0.12 |
| | (1.73) | (0.05) | (3.51) | (−1.51) | (−3.92) | (2.91) | |
| $rx^7_{t,t+1}$ | 4.11 | −0.78 | **65.56** | −19.93 | **−75.88** | **31.42** | 0.07 |
| | (1.20) | (−0.27) | (2.88) | (−1.23) | (−3.06) | (2.76) | |

Notes: The table reports estimates from the same OLS regressions described in Table 7, but over three subperiods: 1968Q4-1981Q4, 1982Q1-1996Q4, and 1997Q1-2013Q2. $P_L$ is the probability of decline in real GDP in the following quarter from the SPF, and $P_H = 1 - P_L$. Newey and West (1987) corrected t-statistics have a lag order of 4 quarters and are reported in parentheses. Coefficients that are significant at the 5% or better level are highlighted in bold.

Figure 1: Inflation Forecast Attributes



# A Appendix

## A.1 Key features of SPF inflation density forecasts

This section summarizes some of the complications dealt with in the construction of our inflation uncertainty measure relating to the SPF density forecasts. For more details regarding the methodology see D'Amico and Orphanides (2008).

The definition of the inflation probability variable, PRPGDP, in the SPF is the probability that the annual-average over annual-average percent change in the price index for GDP (for GNP prior to 1992:Q1) falls in a particular range in the current year (PRPGDP-A) and the following year (PRPGDP-B).

The panelists are provided with a certain number of intervals of a specific width, so that their probabilistic assessments are summarized by histograms. However, since the survey's inception: first, the width and number of intervals changed over time (i.e., in 1981Q3; 1992Q1; and 2014Q1); second, these changes coincided also with a change in the range spanned by the intervals[18]; and third, the number of annual forecast horizons changed once (i.e., prior to 1981Q3 there was only one forecasting horizon, PRPGDP-A).

As indicated by the definition, because the survey question refers to the calendar year, the forecasting horizon is not fixed but varies in each quarter, as a result the underlying inflation uncertainty and disagreement in the individual responses have a seasonal pattern that needs to be corrected. This in conjuction with the changing number and width of the intervals over time introduces an important complicating factor: As the horizon gets shorter, a sizable percentage of respondents concentrate the probability mass in less than 3 intervals making it difficult to fit continuous distributions to the individual histograms and obtain higher moments of the distributions. This is clearly illustrated in Table 1A, which shows the percentage of responses in less than 3 intervals as function of the forecasting horizon. The numbers in the second and third columns show that when the forecasting horizon becomes 1 or 2 quarters ahead, in periods with larger and fewer bins as in 1981Q3-1991Q4, up to almost 73 percent of respondents place their probability mass in less than 3 bins. And even in other periods, values can range between 30 and 50 percent.

Table 1A
Percentage of responses in less than 3 bins as function of the forecasting horizon

| | $1-Q$ | $2-Q$ | $3-Q$ | $4-Q$ | $5-Q$ | $6-Q$ | $7-Q$ | $8-Q$ |
|---|---|---|---|---|---|---|---|---|
| $68Q4-81Q2$ | 12.6 | 18.2 | 17.3 | 10.7 | | | | |
| $81Q3-91Q4$ | 72.8 | 47.7 | 38.5 | 32.5 | 30.3 | 24.6 | 24.5 | 29.3 |
| $92Q1-13Q4$ | 51.3 | 30.9 | 23.0 | 17.1 | 16.0 | 14.5 | 13.3 | 12.9 |
| $68Q4-13Q4$ | 40.9 | 29.0 | 23.4 | 16.5 | 19.9 | 17.1 | 16.1 | 16.3 |

To correct for the bias introduced by respondents that concentrate their answers in less than 3 intervals, first of all, in each quarter, we compute non-parametric moments of the individual histograms to avoid fitting continuous distributions that can not be uniquely

[18] However, who did not increase appropriately the number of bins used, also ehibits unreasonable changes in the range of possible inflation values. Example: there is one individual that uses 3 bins in both quarters and expected in 1991Q4 one-year ahead inlfation to fall between 1%-5% with 60% probability associated to 3% ad 30% probability to 5% (mean=3.4), in 1992Q1 also belives that inflation will fall between 0.5-2.5% with 70% probability on 1.5% and 20% on 0.5% (mean=1.4%), so the 1991Q4 mode (3%) is not even included in the new range of possible inflation values.

identified with fewer than 3 bins. Second, the truncated Gamma methodology is then applied to current-year and following-year individual non-parametric variances as described in Section 2. Further, differently from D&O, the threshold for the truncation is not fixed but is set equal to the maximum empirical variance (after applying the Sheppard's correction) that a forecaster using only two bins can have, given a specific bin's size in the histogram. This change hardly affected the measure.

To obtain a roughly constant-horizon, one-year ahead measure of mean inflation, inflation uncertainty, and disagreement, we use a weighted average of the responses for PRPGDP-A and PRPGDP-B, where in each quarter the weights are function of the distance of the current-year and following-year responses from the targeted one-year horizon. And then a seasonal adjustment (X-12 ARIMA filter) is applied. This weighted average can be computed only starting in 1981Q3, when the questions in the SPF refer to both current and following year forecasts. Prior to 1981Q3, the probability variables usually referred to percent changes for the current year, which makes the problem of the shrinking forecasting horizon more severe, especially in the fourth quarter. However, during this period there were some exceptions where the probability variables referred to the percent change in the following year, rather than the current year. The surveys for which this is true are 1968Q4, 1969Q4, 1970Q4, 1971Q4, 1972Q3 and Q4, 1973Q4, 1975Q4, 1976Q4, 1977Q4, 1978Q4, and 1979Q2-Q4. For example, the probability ranges in the 1968Q4 survey should pertain to inflation in 1968, but in fact the survey asked about probabilities for inflation in 1969. This makes the forecasting horizon closer to one year and allows us to exclude only two observations (1974Q4 and 1980Q4), interpolate over the missing values, and then apply the seasonal adjustment.

Finally, to try to address any additional distortion in our measurement of inflation uncertainty and disagreement in the particular quarters in which a change in the number and width of the intervals is introduced, we exclude the histograms of those respondents who did not appropriately increase (decrease) the number of bins they used when the bin size was halved (doubled). This resulted in a sizable correction for 1992Q1 but not for 1981Q3.

On a different note, since the Philadelphia Fed is uncertain about the years referred to in the surveys of the first quarters of 1985 and 1986, which likely did not follow the standard of the current and following year, these quarters are excluded from the analysis.

### A.2 Additional robustness checks

To verify that the result pertaining to the last recession (i.e., the inflation uncertainty does not increase differently from previous recessions) is not due to smoothing issues related to the *annual-average* over *annual-average* percent change, we also compute the inflation uncertainty and disagreement from other two variables that are available in the SPF starting from 2007:Q1 and refer to *fourth-quarter* over *fourth-quarter* percent change. The first one, PRCCPI, gives the mean probability that the fourth-quarter over fourth-quarter percent change in core CPI falls in a particular range, and the second one, PRCPCE, gives the mean probability that the fourth-quarter over fourth-quarter percent change in core PCE falls in a particular range.

Using these probabilistic responses we obtain the same result: inflation disagreement increased during the last recession but inflation uncertainty did not.

Further, we even recomputed all the individual histograms keeping the size and the number of bins constant through the entire history of the SPF. This is done by cumulating the probability mass only in 6 bins with size of 2%, which is the lowest number of bins ever used in the survey. Inflation uncertainty and disagreement are then measured from these new histograms and contrasted to those obtained from the original histograms. Since these new measures covary very closely with the original measures, the magnitude of the correction does not have any significant impact on our analysis.

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