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EXPECTED INFLATION AND INFLATION RISK PREMIUM IN THE EURO AREA AND IN THE UNITED STATES

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Abstract

This paper uses the celebrated no-arbitrage affine Gaussian term structure model applied to index-linked and standard government bonds to derive expected inflation rates and inflation risk premia, in the euro area and in the United States. Maximum likelihood estimates show that the model describes the evolution of the nominal and real term structures by using three latent factors which can be interpreted as two real factors and one inflation factor. These provide important information on expected inflation and inflation risk premia. The results highlight some striking differences between the euro area and the United States. In the United States, forward inflation risk premia become sizable around the start of the late-2000s financial crisis and considerably increase just before the adoption of the first unconventional monetary policy measures in March 2009. By contrast, in the euro area forward inflation risk premia remain unchanged even after the adoption of the unconventional monetary policy measures following the most acute phases of the financial crisis, in October 2008 and in May 2010. However, long-term inflation expectations have been well anchored over the past years.

JEL Classification: C02, G10, G12.
Keywords: real and nominal term structure, inflation risk premium, affine term structure, Kalman filter.

Contents

* Bank of Italy, Economics, Research and International Relations.

1 Introduction[1]

Over the past decade government-issued inflation-indexed bonds have become available in a number of euro-area countries and have provided a fundamentally new instrument sought after by institutional investors and by households, especially for retirement saving. From a policy perspective, inflation-indexed bonds can be used to infer inflation expectations at different maturities. In fact, returns on bonds linked to an inflation index differ from the corresponding returns on standard bonds in terms of the expected inflation and the inflation risk premium (some technical features of bonds such as maturity, coupon rate and cash-flow structure contribute to this difference). In addition, since index-linked bonds have different maturities, an entire spectrum of inflation expectations and inflation risk premia can be derived from a comparison with standard nominal bonds. Hence, stemming from the no-arbitrage affine Gaussian term structure literature developed for standard bonds, some recent papers have investigated a theoretical and empirical framework to jointly price standard and index-linked bonds on the basis of a small number of common factors; the innovation from this stream of literature, to which this paper belongs, is to have consistent, i.e. arbitrage-free, estimates of the real and nominal interest rates as well as expected inflation rates and inflation risk premia.

This paper presents estimates of a no-arbitrage affine Gaussian term structure model for the nominal and real zero-coupon interest rates implied in standard and index-linked government bonds, respectively, in the euro area and the United States. This class of models gives one the opportunity to divide a model-implied constant-maturity inflation compensation (or model-implied breakeven inflation rate), obtained as the difference between the estimated nominal and real zero-coupon rates, into the expected component (i.e. the expected inflation) and the premium requested by investors to hedge against unexpected changes in inflation (i.e. the inflation risk premium).

Forward breakeven inflation rates, i.e. the 5-year inflation rates in 5-years' time implied by nominal and real interest rates, are usually taken as proxies for inflation expectations and provide a measure of a central bank's credibility in targeting a specific inflation rate – in the case of the European Central Bank, for instance, this target is defined as a rate of inflation below, but close to, 2 per cent in the medium term. Inflation compensation provides a measure of inflation expectations since the payoff of a nominal bond in real terms should be close to that of an index-linked bond over its entire life. A forward-looking way to evaluate the success of monetary policy is to look at expectations of inflation; in fact, if monetary policy is successful at keeping expectations well-anchored, then financial market participants would tend to "look through" the cycles of inflation and not change expectations about the rate of inflation over the longer run. The low level of inflation and the non-conventional monetary policies observed over the last few years have raised concerns about the possibility that market participants were no longer seeing central bank policy as committed to long-run price stability.

The use of a model which jointly estimates expected inflation and the inflation risk premium presents three advantages with respect to the use of the plain-vanilla break-even inflation rates

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computed as the difference between nominal and real interest rates as well as with respect to inflation swaps. First, over longer time horizons, the breakeven inflation rate can substantially differ from expected inflation since the compensation requested by investors for uncertainty about future inflation rates – i.e. the inflation premium – can be significant. Second, real and nominal interest rates are estimated on the basis of a common set of factors which drive the entire nominal and real term structure; so this class of models is able to give an intuition of the economic drivers of the nominal and real term structures. Third, as expected inflation is a key ingredient in monetary policy decisions, it is important to have fresh and readily available updates; expected inflation obtained from the nominal and real term structures is available at a much higher frequency than the information provided by the Survey of Professional Forecasters (1-year and 5-year inflation forecasts are released quarterly in the United States, while 1-year, 2-year, and 5-year inflation forecasts are released quarterly in the euro area) and by Consensus Economics (long-term inflation forecasts are released half-yearly in the United States and in the euro area).

This paper also considers a simple correction to take into account the liquidity risk, due to the lower liquidity of index-linked with respect to standard bonds, and the seasonality bias of the consumer price reference index (Pericoli, 2012).

This paper innovates with respect to the previous macrofinance literature. First, this work uses weekly data for the euro area; previous works with weekly data are those by Risa (2001) for the United Kingdom and by Adrian and Wu (2010) for the United States. Second, the three latent factors are interpreted as a transformation of observable financial variables and this helps in assigning an economic interpretation to these factors which drive the shape of the nominal and real term structures. Third, the same methodology is applied to the euro area and to the United States allowing a consistent comparison between the two markets. The use of weekly data is essential when this class of models is used by monetary policy makers to evaluate inflation expectations and inflation risk premia.

Results show that: i) the model is capable of fitting quite well the nominal and the real term structures with small or negligible pricing errors; ii) the three latent factors, which drive the nominal and real term structures, can be interpreted as the cross section average of real interest rates, the slope of the real term structure and the breakeven inflation rate; iii) long-term inflation risk premia are positive and somewhat larger in the United States than in the euro area; iv) these premia present countercyclical dynamics and behave very similarly to other indicators of financial market risk.

The paper is organized as follows. Section 2 reviews the literature and Section 3 defines the inflation compensation. The model and the estimation problems are presented in Sections 4-6. Data are described in Section 7. Estimates of the term structure and the impulse response analysis are presented in Section 8, robustness checks in Section 9. Section 10 concludes.

2 The literature

Recent papers on the term structure of inflation can be divided between two broad groups. The first uses the standard set up of the no-arbitrage Gaussian affine term structure models of nominal and real interest rates with an identification assumption meant to increase the power of the estimates;

Evans (1998), Risa (2001), Joyce et al. (2010), Ang et al. (2008), D'amico et al. (2008), Christensen et al. (2010), Garcia and Werner (2010), Adrian and Wu (2010), and Haubrich et al. (2011a, 2011b) follow this line of research. Alternatively, a second stream of works uses standard new-Keynesian macrofinance models which encompass financial and macro variables; the works by Chernov and Mueller (2008) and Hördal and Tristani (2010) are along this line of research. This paper belongs to the former class of papers.

Evans (1998) and Risa (2001) use a no-arbitrage Gaussian affine term structure model and study the term structure of real and nominal rates, expected inflation, and inflation risk premia derived from the prices of index-linked and nominal debt in the United Kingdom. Both authors find strong evidence of variable inflation risk premia throughout the term structure and, furthermore, reject both the Fisher Hypothesis and versions of the Expectations Hypothesis for real rates. In these papers the variability of the nominal to real yield spread is mostly due to inflation at the short end and to its premium at the long end.

Ang et al. (2008) develop a term structure model with regime switches, time-varying prices of risk, and inflation to identify these components of the nominal yield curve. They find that the unconditional real rate curve in the United States is fairly flat around 1.3%. In one real rate regime, the real term structure is steeply downward sloping. An inflation risk premium that increases with maturity fully accounts for the generally upward sloping nominal term structure.

Christensen et al. (2010) show that the affine arbitrage-free Nelson–Siegel model can be estimated for a joint representation of nominal and real yield curves in the United States. Results suggest that long-term inflation expectations have been well anchored over the past few years in the United States and that the inflation risk premium, although volatile, has been close to zero on average. Haubrich et al. (2011a, 2011b) estimate the term structure of inflation expectations and inflation risk premia by means of data on inflation swap rates, nominal Treasury yields and survey forecasts of inflation; the use of inflation swap rates rules out the problems connected with the illiquidity of the index-linked Treasuries. They find that the short-term real interest rate is typically the most volatile component of the yield curve, that expected inflation over short horizons is also volatile, that investors' expectation of longer-term inflation declines substantially during the last twenty years, and that the 10-year inflation risk premium varies between 23 and 55 basis points. Joyce et al. (2010) and D'Amico et al. (2008) document the importance of using index-linked bonds for accurate predictions of inflation for the United Kingdom and for the United States, respectively.

Garcia and Werner (2010) document that no-arbitrage Gaussian affine term structure models fit data well in the euro area, but lack economic interpretation; so the authors introduce survey inflation risks and show that perceived asymmetries in inflation risks help interpret the dynamics of long-term inflation risk premia, even after controlling for a large number of macro and financial factors. Similarly Adrian and Wu (2010) present estimates of the term structure of inflation expectations, derived from an affine model of real and nominal yield curves for the United States. The model features stochastic covariation of inflation with the real pricing kernel. The authors fit the model not only to yields, but also to the yields' variance-covariance matrix, thus increasing the identification power, and find that model-implied inflation expectations can differ substantially from breakeven inflation rates when market volatility is high.

Within the second set of works, Chernov and Mueller (2008) use evidence from the term structure of inflation expectations to address the question of whether or not monetary policy is effective.

They show that the inflation premia and out-of-sample estimates of long-term inflation suggest that U.S. monetary policy became effective over time. As an implication, their model outperforms standard macro-finance models in inflation and yield forecasting. Hördal and Tristani (2010) extend a traditional new-Keynesian macro-finance model by encompassing the nominal and the real term structure and introduce survey data on inflation and interest rate expectations at various future horizons. They show that in the euro area and in the United States, inflation risk premia are relatively small, positive, and increasing in maturity. The cyclical dynamics of long-term inflation risk premia are mostly associated with changes in output gaps, while their high-frequency fluctuations seem to be aligned with variations in inflation. However, inflation premia are countercyclical in the euro area, while they are procyclical in the US.

3 The inflation compensation

The comparison between the nominal and real term structure gives the inflation compensation requested by investors to hold standard nominal bonds. This compensation, known as the break-even inflation rate (BEIR), is equal to the difference between the nominal and the real interest rates, namely

$$BEIR_t^n = y_t^n - r_t^n \tag{1}$$

where y_t^n is the nominal interest rate at time t for maturity n, and r_t^n is the corresponding real interest rate. However, the BEIR is not a pure expectation of the inflation rate since, as shown by Evans (1998), it can be thought of as the sum of the expected inflation rate at time t during the n periods to maturity, $\pi_t^{e,n} = \frac{1}{n} E_t \left[\Sigma_{i=1}^n \pi_{t+i}\right]$, and the inflation risk premium at period t, γ_t^n, namely

$$\begin{aligned} BEIR_t^n &= y_t^n - r_t^n \\ &= \pi_t^{e,n} + \gamma_t^n \end{aligned}$$

It can be shown that if variables are jointly lognormal, this risk premium is given by $\gamma_t^n = Cov(m_t^n, \pi_t^{e,n}) - \frac{1}{2} Var(\pi_t^{e,n})$, where m_t^n is the stochastic discount factor between period t and $t+n$ and $\pi_t^{e,n}$ the expected inflation rate over the same period; in other words, the premium requested by investors to hold indexed-linked bonds and to hedge against unexpected changes in inflation depends on the covariance between the marginal rate of substitution (the stochastic discount factor) and the inflation rate; the second term is a convexity adjustment, arising from the Jensen inequality. Sometimes, the first term of the inflation risk premium, $Cov(m_t^n, \pi_t^{e,n})$, is referred to as the 'pure inflation risk premium'.

The inflation risk premium, i.e. the compensation for risk due to the uncertainty of future inflation, can be evaluated by means of a joint model of the nominal and real term structure. This premium, in a standard representative-agent power-utility model, is positive when the covariance between the stochastic discount factor and inflation is negative, in other words when expected consumption growth is low and inflation is high.

4 The general model

This paper uses a no-arbitrage standard Gaussian affine term structure model, set in discrete time, as in the majority of the recent literature about macro term structure models. The term structures for nominal and real interest rates are linked through the pricing kernel corrected by the inflation rate. This model follows the original setup by Evans (1998), successively enhanced by Risa (2001), Garcia and Werner (2010) and Adrian and Wu (2010).

4.1 The real term structure

The model consists of three equations. The first equation describes the dynamics of the vector of state variables X_t (a k-dimensional vector, $k \in \mathbb{N}$):

$$X_t = \mu + \rho X_{t-1} + \Sigma \varepsilon_t \ , \tag{2}$$

where $\varepsilon_t \sim N(0, I_k)$, μ is a $k \times 1$ vector and ρ and Σ are $k \times k$ matrices. Without loss of generality, it can be assumed that Σ is lower triangular. Furthermore, to ensure stationarity of the process, we assume that all the eigenvalues of ρ lie strictly inside the unit circle. The probability measure associated to the above specification of X_t will be denoted by P. X_t is a matrix containing k latent factors, which can be thought of as $k-1$ real factors and one inflation factor.

The second equation relates the one-period interest rate $r_t^1 = r_t$ to the state variables (positing that it is an affine function of the state variables):

$$r_t = -\delta_0 - \delta_1^\intercal X_t \ , \tag{3}$$

where δ_0 is a scalar and δ_1 is a $k \times 1$ vector with the last element equal to zero since the real rate is not affected by the inflation rate.

The third equation is related to bond pricing in an arbitrage-free market. A sufficient condition for the absence of arbitrage on the bond market is that there exists a risk-neutral measure Q, equivalent to P, under which the process X_t follows the dynamics:

$$X_t = \overline{\mu} + \overline{\rho} X_{t-1} + \Sigma \eta_t \ , \tag{4}$$

where $\eta_t \sim N(0, I_k)$ under Q and such that the price at time t of a bond paying a unitary amount of consumption (or an amount of cash indexed to the inflation rate) at time $t+n$ equals:

$$p_t^n = \mathrm{E}_t^Q \left[\exp(-r_t) \, p_{t+1}^{n-1} \right] \ , \tag{5}$$

where E_t^Q denotes expectation under the probability measure Q, conditional upon the information available at time t.

The vector $\overline{\mu}$ and the matrix $\overline{\rho}$ are in general different from μ and ρ, while equivalence of P and Q guarantees that Σ is left unchanged. The link between the risk-neutral distribution Q and the physical distribution P is given by the (time-varying) price of risk which is affine in the state variables:

$$\lambda_t = \lambda_0 + \lambda_1 X_t \ ,$$

where $\lambda_0 = \Sigma^{-1}(\mu - \overline{\mu})$ and $\lambda_1 = \Sigma^{-1}(\rho - \overline{\rho})$. According to Cameron, Martin and Girsanov's theorem (e.g. Kallenberg - 1997)

$$\mathrm{E}_t^P\left[\frac{dQ}{dP}\right] = \prod_{j=1}^{\infty} \exp\left[-\frac{1}{2}\lambda_{t+j-1}^{\top}\lambda_{t+j-1} - \lambda_{t+j-1}^{\top}\varepsilon_{t+j}\right] \ ,$$

so that the real pricing kernel

$$m_{t+1} = \exp\left(-r_t - \frac{1}{2}\lambda_t^{\top}\lambda_t - \lambda_t^{\top}\varepsilon_{t+1}\right) \tag{6}$$

can be used to recursively price bonds:

$$p_t^n = \mathrm{E}_t^P\left[m_{t+1}p_{t+1}^{n-1}\right] \ . \tag{7}$$

Multifactor affine models of the term structure, such as the one just described, are very popular in the finance literature and their properties have long been studied by many researchers. Thorough specification analyses of these models have been conducted (e.g. Dai and Singleton, 2000) and their properties are now well-known. A distinguishing feature of these models is that they are able to describe the dynamics of yields in terms of a small set of unobservable state variables: typically three variables are deemed a sufficient number to describe the whole yield curve and this is supported also by empirical studies, such as the seminal paper by Litterman and Scheinkman (1991). Although such models are capable of describing accurately and parsimoniously the evolution of interest rates over time, the factors they identify as the driving forces of interest rates often lack economic intuition and are difficult to relate to the relevant economic variables. This is one of the reasons why recent studies have proposed augmenting the usual set of unobservable state variables with some observable variables. Typically, inflation and a measure of the output gap are the two observable variables, while a small number of unobservable factors, ranging from one to three, are included in the models: recent examples are Ang and Piazzesi (2003), Rudebusch and Wu (2008), Hördal, Tristani and Vestin (2004) and Ang et al. (2008). All these works impose a set of restrictions on the system of equations (2-4) and, after estimating the coefficients, derive bond prices using equation (5). When index-linked and standard bonds are considered, the actual inflation can be substituted by the breakeven inflation rate, i.e. the difference between the nominal and the real rates implied in bonds. Thus, this class of models does not consider the inflation rate among its state variables.

Note that within this Gaussian framework, bond yields are affine functions of the state variables:

$$r_t^n = -\frac{1}{n}\ln(p_t^n) = A_n + B_n^{\intercal}X_t \ ,$$

where r_t^n is the yield at time t of a bond maturing in n periods and A_n and B_n are coefficients obeying the following simple system of Riccati equations, derived from (5)[2]:

$$\begin{aligned} A_1 &= -\delta_0 \\ B_1 &= -\delta_1 \\ &\ldots \\ &\ldots \\ A_{n+1} &= -\delta_0 + A_n + B_n^{\intercal}(\mu - \Sigma\lambda_0) - \frac{1}{2}B_n^{\intercal}\Sigma\Sigma^{\intercal}B_n \\ B_{n+1} &= -\delta_1 + B_n^{\intercal}(\rho - \Sigma\lambda_1) \ . \end{aligned} \tag{8}$$

[2]A proof by induction for a more general case can be found, for example, in Dai, Singleton and Yang (2003).

The yields $\widetilde{r}_t^n$ and the bond prices $\widetilde{p}_t^n$ that would obtain in an arbitrage-free market populated by risk neutral investors are instead obtained setting the prices of risk to zero ($\lambda_t = 0$) in (6) and (7):

$$\widetilde{p}_t^n = \mathrm{E}_t^P \left[\exp\left(-r_t\right) \widetilde{p}_{t+1}^{n-1} \right] \ .$$

They obey the same system of recursive equations (8), where $\overline{\mu}$ and $\overline{\rho}$ are substituted by μ and ρ. Subtracting the risk-neutral yields $\widetilde{r}_t^n$ thus calculated from the actual yields r_t^n one obtains the term risk premia ϕ_t^n:

$$\phi_t^n = r_t^n - \widetilde{r}_t^n \ ,$$

which is the additional interest per unit of time required by investors for bearing the risk associated with the fluctuations of the price of a bond expiring in n periods. Such premia are time varying in general, and they are constant only when $\lambda_1 = 0$, i.e. for $\rho = \overline{\rho}$.

4.2 The nominal term structure

Nominal bond prices are priced by the nominal pricing kernel $\widehat{M}$ which is linked to the real pricing kernel through the inflation rate, Π, i.e. the change in the consumer price index. Given the following relation $\widehat{M}_{t+1} = M_{t+1}/\Pi_{t+1}$, the log nominal pricing kernel is given by

$$\begin{aligned} \log \widehat{M}_{t+1} &= \widehat{m}_{t+1} = m_{t+1} - \pi_{t+1} \\ &= m_{t+1} - \exp(e_K^\top X_{t+1}) \\ &= \exp\left(-r_t - \frac{1}{2}\lambda_t^\top \lambda_t - \lambda_t^\top \varepsilon_{t+1} - e_K^\top X_{t+1}\right) \ , \end{aligned}$$

where $e_K = (0, ..., 0, 1)^\top$ and thus $e_K^\top X_{t+1}$ picks the inflation rate. Using the affine pricing rule the price of a nominal bond is given by

$$\begin{aligned} \exp\left(\widehat{A}_{n+1} + \widehat{B}_{n+1}^T X\right) &= \exp\Big[-\delta_0 + \widehat{A}_n + \left(\widehat{B}_n^\intercal - e_K^\top\right)(\mu - \Sigma\lambda_0) \\ &\quad -\frac{1}{2}\left(\widehat{B}_n^\intercal - e_K^\top\right)\Sigma\Sigma^\intercal \left(\widehat{B}_n^\intercal - e_K^\top\right)^T \\ &\quad + \left(-\delta_1 + \left(\widehat{B}_n^\intercal - e_K^\top\right)(\rho - \Sigma\lambda_1)\right) X_t\Big] \ , \end{aligned}$$

where

$$\begin{aligned} &\widehat{A}_1 = -\delta_0 - e_K^\top \mu + \frac{1}{2} e_K^\top \Sigma\Sigma^\intercal e_K + e_K^\top \Sigma\lambda_0 \\ &\widehat{B}_1 = -\left(\delta_1 + e_K^\top \rho\right) + e_K^\top \Sigma\lambda_1 \\ &\ldots \\ &\ldots \\ &\widehat{A}_{n+1} = -\delta_0 + \widehat{A}_n + \left(\widehat{B}_n^\intercal - e_K^\top\right)(\mu - \Sigma\lambda_0) \\ &-\frac{1}{2}\left(\widehat{B}_n^\intercal - e_K^\top\right)\Sigma\Sigma^\intercal \left(\widehat{B}_n^\intercal - e_K^\top\right)^T \\ &\widehat{B}_{n+1} = -\delta_1 + \left(\widehat{B}_n^\intercal - e_K^\top\right)(\rho - \Sigma\lambda_1) \ . \end{aligned} \tag{9}$$

5 The estimation problem

The term structure model is expressed in the state-space form (Hamilton, 1989)

$$\begin{aligned} Y_t &= A + HX_t + R\eta_t \quad \text{(observation equation)}, \\ X_t &= \mu + \rho X_{t-1} + \Sigma\varepsilon_t \quad \text{(state equation)}, \\ R &\perp \Sigma \,, \end{aligned} \tag{10}$$

where $A = [\widehat{A}_1, ..., \widehat{A}_{N,} A_1, ..., A_R]$, $H = [\widehat{B}_1, ..., \widehat{B}_{N,} B_1, ..., B_R]$, N and R are the number of nominal and index-linked bonds used in the estimation, $\varepsilon_t \sim N(0, I_k)$, and $\eta_t \sim N(0, I_{N+R})$. The matrix Y_t contains the N nominal zero-coupon rates with annual maturity from 3 to 10 years, and the R real zero-coupon rates with annual maturity from 3 to 10 years. The matrix X_t contains three latent factors $[l_t^1, l_t^2, \pi_t]$, two real factors and the inflation rate.

Pericoli and Taboga (2008) show that, without loss of generality, it is possible to assume that ρ is lower triangular and that matrix Σ is diagonal with all diagonal elements equal to one but the last, namely

$$\begin{aligned} \rho &= \begin{bmatrix} \rho_{11} & 0 & 0 \\ \rho_{21} & \rho_{22} & 0 \\ \rho_{31} & \rho_{32} & \rho_{33} \end{bmatrix} \\ \Sigma &= \begin{bmatrix} 1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & \sigma_\pi \end{bmatrix} . \end{aligned}$$

matrix R is a 16×16 diagonal matrix whose main diagonal is given by

$$R = diag\left[\sigma_N(3), ..., \sigma_N(10), \sigma_R(3), ..., \sigma_R(10)\right] \,,$$

where $\sigma_N(\tau)$ and $\sigma_R(\tau)$ are the standard deviations of the nominal and real bond with maturity τ. Let's assume that the standard deviation of the observation errors are is non increasing in the term to maturity τ, i.e. the volatility is lower for bonds with longer maturities:

$$\begin{aligned} \sigma_N(\tau) &= c_N + d_N/\sqrt{\tau} \\ \sigma_R(\tau) &= c_R + d_R/\sqrt{\tau} \,. \end{aligned}$$

Note that this form can reflect several possible definitions of the observation error; when d_N and d_R are equal to zero the price errors are constant across maturities (Risa, 2001).

Based on the state space representation in (10), the factors are filtered according to the Kalman filter; given estimates of the latent factors $\widehat{X}_t$, the parameters can be estimated by maximum likelihood, based on the conditional distribution of $Y_t|Y_{t-1}$ for each observation.

Expected inflation for different horizons can be obtained from equation (10). The τ-period ahead conditional expectation of inflation $E_t(\pi_{t+\tau}) = E_t\left(e_K^\top X_{t+\tau}\right)$ is given by

$$\frac{1}{\tau}E_t(e_K^\top X_{t+\tau}) = [0 \;\; 0 \;\; 1] \cdot \left[(I-\rho)^{-1}(I-\rho^\tau)\,\mu + \rho^\tau \cdot X_t\right] \,. \tag{11}$$

5.1 Model specification

The complete model is defined by the following parameters

$$
\begin{aligned}
\rho &= \begin{bmatrix} \rho_{11} & 0 & 0 \\ \rho_{21} & \rho_{22} & 0 \\ \rho_{31} & \rho_{32} & \rho_{33} \end{bmatrix} , \\
\mu &= (0, 0, \mu_\pi)^\intercal , \\
& \delta_0 , \\
\delta_1 &= (\delta_1^1, \delta_1^2, 0)^\intercal , \\
\Sigma &= \begin{bmatrix} 1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & \sigma_\pi \end{bmatrix} , \\
\lambda_0 &= (\lambda_0^1, \lambda_0^2, \lambda_0^3), \quad \lambda_1 = \begin{bmatrix} \lambda_{1,11} & \lambda_{1,12} & \lambda_{1,13} \\ \lambda_{1,21} & \lambda_{1,22} & \lambda_{1,23} \\ \lambda_{1,31} & \lambda_{1,32} & 0 \end{bmatrix} , \\
\sigma_N(\tau) &= c_N + d_N/\sqrt{\tau}, \text{ for } \tau = 3, ..., 10 , \\
\sigma_R(\tau) &= c_R + d_R/\sqrt{\tau}, \text{ for } \tau = 3, ..., 10 .
\end{aligned}
$$

5.2 Choice of the number of factors

In order to test the performance of the 3-factor model, the same model with 2 and 4 factors has been estimated. The problem of selecting the number of factors is cumbersome; Likelyhood Ratio tests cannot be used to test for the number of statistically relevant factors, since some of the parameters become unidentified under the null. Previous studies use 3 factors (D'amico et al., 2008, Christensen et al., 2010, and Garcia and Werner, 2010), 4 factors (Risa, 2001) and 5 factors (Adrian and Wu, 2010). The criterion for using 3 factors in this paper is based on the cumulative explained variance of nominal and real interest rates obtained by a Principal Component Analysis. For the euro area bond market, variances explained by the first, second, third and fourth principal factors are 88.5%, 8.4%, 2.8% and 0.2%, respectively; those for the United States market are 91.3%, 7.1%, 1.4% and 0.2%. Then, in both markets the first three factors explain more than 99.5% of the total variance and this is deemed sufficient for choosing 3 factors. A more thorough analysis can be made either by comparing out-of-sample errors of the pricing equations (as in Risa, 2001) or by cross-validation. After a graphical inspection it appears that, in the case of 2 factors, the fit is not able to capture the dynamics of the term structure; and the unique real latent factor, which proxies the cross-sectional average of real rates, is not able to capture the cross-sectional dispersion among interest rates. With 4 factors, the model tends to overfit the term structures both for real and nominal interest rates.

6 The data

Nominal and real zero coupon interest rates for the euro area are estimated from end-of-week quotes of French government bonds by means of the methodology first introduced by Fisher at al. (1995). The nominal term structure is estimated by using the quotes of the euro repo rates with maturity at 1 week, 2 weeks, 3 weeks, 1 month, 2 months, 3 months, 6 months, 9 months, and 12 months for the short term, of the BTANs (*Bon à Taux Annuel Normalisé*) with maturity greater than 1 year and below 5 years, and of standard OATs (*Obligations Assimilables au Trésor*) with maturity longer than 1 year. Data run from January 2002 to August 2011.

The real term structure for the euro area is estimated by using OAT€i's, i.e. OATs indexed to the euro-area harmonised index of consumer prices (HICP) excluding tobacco, the reference price index of the eurozone. The euro-area index-linked bond market started in 1998 with the issue of French government bonds, OATi's, indexed to the domestic French Consumer Price Index (CPI). In 2002 there was the first issue of French government bonds, OAT€i's, indexed to the euro-area HICP excluding tobacco. This work considers French index-linked bonds; Italian index-linked bonds have a lower rating while German index-linked bonds are characterized by a much shorter history and very few issues.

End-of-week mid-quotes are obtained from Bloomberg and Thomson Financial Reuters. The daily consumer price index reference, available on Bloomberg, is obtained from the website of the European Central Bank (http://www.ecb.int) and from the website of the French Treasury (www.aft.gouv.fr). Weekly term structures are obtained by considering the last day of the week.

The issue of the different liquidity of standard and index-linked bonds is addressed by estimating the nominal term structure with off-the-run bonds and notes, i.e. those issued before the most recently issued bonds or notes of a particular maturity. To check the consistency of this measure of liquidity, estimated differences between off-the-run and on-the-run interest rates are compared with the differences between the zero-coupon rate extracted from bonds issued by CADES, a French government agency whose securities are fully backed by the French Treasury, and the corresponding interest rate extracted from the nominal French OATs. Results are very similar.

The nominal and real term structures for the United States are taken from the weekly data estimated by Gürkaynak et al. (2007) and Gürkaynak et al. (2010), respectively. Data run January 1998 to August 2011.[3]

7 Results

Results show that model (10) is capable of jointly estimating the nominal and the real term structures for the euro area and for the United States (Figures 1-2 and Table 1). Both in the euro area and in the United States, estimates track nominal and real interest rates quite closely; only at the shortest maturities are the yield pricing errors large, especially for real rates.

An important step towards a better understanding of the mechanics of a reduced-form no-arbitrage model like (10) consists in assigning an economic interpretation to the latent factors since it helps to provide a deeper insight into the economic forces driving bond prices. From a graphical inspection it emerges that latent factors can be interpreted as the cross-sectional average of the real term structure (first factor), the slope of the real term structure computed as difference between the 10-year and the 3-year real zero-coupon rate (second factor) and the 10-year breakeven inflation

[3]Weekly updates are available at http://www.federalreserve.gov/econresdata/researchdata.htm.

rate (third factor); this latter is a wedge between the nominal and the real interest rate (Figures 3-4). Table 3 documents that the correlation between the first latent factor and the average of the real rates is around 0.8 in both areas and increases to almost 0.9 in the period from 2007 to 2011; also the correlation between the third latent factor and the breakeven inflation rate is around 0.7 and jumps to 0.9 from 2007; the correlation between the second latent factor and the slope of the real term structure is much lower but close to 0.5 from 2007.

The standard three-factors model introduced by the seminal work of Litterman and Scheinkman (1991) introduces the nominal term structure average, the slope of the nominal term structure and the curvature of the nominal term structure as the main driving forces of the nominal term structure. In contrast, model (10) considers two real factors (the average of the real term structure and the real long-term interest rate) and an inflation factor; the latter summarizes the information embedded in the slope of the nominal term structure and in its curvature.

The estimates of long-term inflation expectations given by equation (11) are plotted in the top panels of Figures 5 and 6 for the United States and the euro area. In the United States the 10-year expected inflation rate envisages some swings around 2 per cent from 1998 until the middle of 2004 when it starts showing steady values with much smaller variations; from the end of 2004 until the middle of 2008 the average of the 10-year expected inflation is equal to 2.2 per cent. The 10-year expected inflation drops in the second half of 2008 to almost nil and steadily increases in the course of 2009 up to 2 per cent. The 10-year breakeven inflation rate quite closely tracks the corresponding expected inflation until the middle of 2001 but records higher values since then; this explains why the U.S. 10-year inflation risk premium, e.g. the difference between the breakeven inflation and expected inflation rates, is almost nil until the middle of 2001 while in the following years it surges to an average of 0.40 percentage points. An alternative indication of inflation expectations comes from the forward expected inflation rate (bottom panels of Figures 5 and 6). The expected inflation forward rates, e.g. the 5-year expected inflation 5 years ahead, is very stable at around an average of 2.1 per cent; only in the last quarter of 2008 does it decline below 2 per cent but it rapidly comes back to its long-term average. Correspondingly, the forward inflation risk premium, given by the difference between the 5-10 year forward breakeven and expected inflation forward rates, is nil on average from 1998 to the middle of 2001, around 0.4 percentage points from the beginning of 2001 and the middle of 2005 when it drops to 0.20 percentage points until the start of the subprime crisis on August 2007; in 2008, against the backdrop of an extremely expansive monetary stance, it increases above 0.5 percentage points and remains above this level.

In the euro area the picture differs slightly. The 10-year expected inflation rate is poorly estimated from 2002 until 2004. It averages above 1.8 per cent from 2004 until 2008 when it drops to 1.4 per cent; from the middle of 2009 and the end of 2010 it is close to 1.8 per cent. Conversely to the United States, in the euro area there is a strong correspondence between the 10-year expected inflation and the 10-year breakeven rates from 2005 until 2008. Accordingly, the 10-year inflation risk premium average is tiny in this period. The indication stemming from the expected forward rates is similar; the expected inflation forward rates, e.g. the 5-year expected inflation 5 years ahead, is stable at around an average of 1.8 per cent with a minor drop in the last quarter of 2008. The forward inflation risk premium, given by the difference between the 5-10 year forward breakeven and expected inflation forward rates, records wide oscillations given by the large variability of the 5-10 year forward breakeven. However, it is on average around 50 basis points from 2005. A caveat is in order; in fact results for the euro area for the 2002-2004 period can be biased by the small number of index-linked bonds as well as by their extremely low liquidity.

7.1 A breakdown of the results during the crisis

The comparison between expected inflation forward rates and forward risk premia in the euro area and in the United States brings to light some differences between the two areas. Against the background of a euro area expected inflation forward rate well anchored below 2 per cent, the forward inflation risk premium has recorded constant figures of around 0.5 percentage points; the two variables have barely changed after the unconventional monetary policy measures introduced in the aftermath of the financial crisis. A first bold wave of unconventional monetary policy measures put forward by the ECB starts at the beginning of October 2008; a second wave, that coincides with the deterioration of the euro-area government debt markets, starts in May 2010.[4] The spot 10-year inflation risk premium does not significantly change either in October 2008 or in the second half of 2010; similarly the 5-10 year forward risk premium temporarily decreases in the second half of 2010. All in all, there is not a clear effect on risk premia stemming from the unconventional measures.

Conversely, in the United States the sequence of unconventional monetary policy measures, meant to provide quantitative easing, has changed the perception of expected inflation forward rates by market participants and determined a substantial and not-negligible inflation risk premium. In particular, the forward inflation risk premium shows a sudden increase first in late-2008 when speculation about the first wave of unconventional measures (the so-called Quantitative Easing 1, QE1, operative from March 2009 until February 2010) first emerges and again in early August 2010 when speculation about the second wave of measures (the so-called Quantitative Easing 2, QE2, in place since November 2010) starts to intensify.

7.2 Impulse response analysis

Figures 7 and 8 plot the responses of the 5-year, 7-year and 10-year nominal and real interest rates to one standard deviation shocks to the three latent factors. As explained above, even if unobservable, the factors can be reconciled with measurable variables, e.g. two related to real rates and one to inflation, and thus a shock to one of the factors can be interpreted from an economic point of view.

A shock to the first latent factor (which can be thought of as the cross-section average of real rates) leads to an increase in real and nominal rates. The response of shortest maturity rates is greater for real than nominal rates, while for longer maturity rates it is greater for nominal than for real rates. Rates tend to converge to their long-term averages both in the euro area and in the United States. In general, an increase in the the cross-section average of real rates produces a parallel shift in nominal and real term structures; however the amplitude of the shifts is similar

[4] In the euro area the main measures are, in October 2008, the adoption of the Fixed-Rate Full-Allotment procedure (FRFA) on money market rates and the expansion of the collateral eligible for Eurosystem credit operations; in May 2009 the outright purchases in the primary and secondary market of covered bonds (Covered Bonds Purchasing Programme, CBPP) and the lengthening of the refinancing operations through the Long-term 12-month operations; in May 2010 the outright purchases of euro-area government bonds in the secondary market (Securities Markets Programme, SMP. See Cecioni et al. (2011) for a survey on unconventional measures.

across nominal and corresponding real rates in the euro area market while it is smaller for real rates in the United States.

Shocks to the second latent factor (which can be thought of as the slope of the real term structure defined by the difference between the 10-year and the 3-year real rate) have a very different impact on nominal and real rates. In both areas real rates tend to be largely unaffected while nominal rates, after a first drop, tend to increase with short-term nominal rates far above the long-term ones. This effect can be explained by the information content of long-term real rates which embed news about long-term output growth and hence about the reaction function of monetary policy.

Finally, a shock to the third factor (which can be thought of as the long-term breakeven inflation, i.e. the difference between the 10-year nominal and the 10-year real interest rates) leads to a tiny decrease in real rates in both markets and to a larger increase in nominal rates. The shock to breakeven inflation mostly impacts short-term nominal rates while longer maturity nominal rates tend to react less. The response of long-term nominal rates to breakeven inflation rate shocks with respect to that of short-term nominal rates is consistent with the greater stability of long-term forward inflation expectations with respect to the instantenous long-term inflation expectations.

7.3 Comparison with other works[5]

This section presents a comparison of results of this paper with those of Hördal and Tristani (2010) and Garcia and Werner (2010) for the euro area, and with those of Adrian and Wu (2010), Christensen et al. (2010), Hördal and Tristani (2010) and Haubrich et al. (2011a, 2011b) for the United States (Figures 9-10 and Table 4). Note that comparisons are not symmetrical since these authors compute different sets of inflation risk premia with different frequencies. In general, the values of inflation premia obtained in this work are consistent with those presented in the literature with the exception of those presented by Hördal and Tristani (2010) for both areas; this difference may be due to the different set-up of their framework.

For the euro area, the 5-year inflation risk premium 5 years ahead obtained in this paper strongly comoves with that of Garcia and Werner (2010); only in 2009-10 the two premia decouple as that of Garcia and Werner remains close and below 0.4 percentage points. The premium of Hördal and Tristani (2010) behaves quite differently from 2004 until 2008 when it records large and decreasing values; only since mid-2009 the Hördal-Tristani premium comoves with that of Garcia and Werner and with that of this paper. As far as the 10-year inflation risk premium is concerned, this paper's premium comoves with that of Hördal and Tristani (2010) which, however, has negative values since the beginning of 2007. The correlation of the 5-year inflation risk premium 5 years ahead of this paper with that of Garcia and Werner (2010) is around 0.8 and is stable across subsamples, while it drops to 0.2 with that of Hördal-Tristani.

For the United States, the 5-year inflation risk premium 5 years ahead obtained in this paper strongly comoves with that of Adrian and Wu (2010) and with that of Christensen et al. (2010), which nonetheless moves in the opposite direction during the autumn of 2008 after the collpase

[5] I would like to thank Tobias Adrian, Jens Christensen, Joseph Haubrich, George Pennacchi, Peter Ritchken, Oreste Tristani and Thomas Werner who made this comparison possible by allowing me to use the data of their papers.

of Lehman Brothers. Analogously with the euro area, the premium computed by Hördal-Tristani (2010) diverges from the other three and remains close to nil with even negative values for prolonged periods of time. From the beginning of 2004 to the middle of 2008 the 10-year inflation risk premium computed in this paper is on average very close to those of Adrian and Wu (2010) and Haubrich et al. (2011a, 2011b) and slightly larger than those of Christensen et al. (2010) and Hördal-Tristani (2010); these two premia show an extended synchronizaton during the 2004-2011 period with only major differences in levels since the beginning of 2009. In the autumn of 2008, the premium of Adrian and Wu (2010) temporarily increases and returns to the levels of Haubrich et al. (2011a, 2011b) and of this paper. The correlation of the 5-year inflation risk premium 5 years ahead of this paper is 0.5 with that obtained in Christensen et al. (2010) and 0.7 with that of Adrian and Wu (2010); it is negative, -0.4, with that of Hördal-Tristani (2010). The correlation of the 10-year inflation risk premium of this paper is over 0.6 with that of Christensen et al. (2010), it is 0.3 with that of Haubrich et al. (2011a, 2011b), and it is negative with that of Hördal-Tristani (2010). All these values remain similar in the crisis period.

8 Robustness

The goodness-of-fit of estimates have been tested by means of a number of robustness checks. First, the model has been enriched by introducing surveys of inflation expectations; second, a proxy of economic growth has been introduced. This section briefly reviews the main findings of the robustness checks.

Surveys of inflation expectation – Moreover, the surveys of inflation expectation are introduced in the model in order to improve its identification power, as in Chernov and Mueller (2008), D'Amico et al. (2008), Garcia and Werner (2010), and Hördal and Tristani (2010). Alternatively, Adrian and Wu (2010) use time-varying conditional covariation between real and inflation factors to increase the identification power of the model. Haubrich et al. (2011a, 2011b) combine the use of surveys of inflation expectation with four volatility state variables which completely determine the risk premia. Model (10) has been estimated with the quarterly surveys published by the Survey of Professional Forecasters for the United States and for the euro area. Results are very similar to those presented. For the euro area the observation equation of model (10) becomes

$$Y_{S,t} = A_S + H_S X_t + R_S \eta_{S,t} \tag{12}$$

where

$$Y_{S,t} = \begin{bmatrix} Y_t \\ \pi_t \\ \frac{1}{52} E_t\left(\pi_{t+52}\right) \\ \frac{1}{104} E_t\left(\pi_{t+104}\right) \\ \frac{1}{260} E_t\left(\pi_{t+260}\right) \end{bmatrix}, \quad A_S = \begin{bmatrix} A \\ 0 \\ e_K^\top (I-\rho)^{-1}\left(I-\rho^{52}\right)\mu \\ e_K^\top (I-\rho)^{-1}\left(I-\rho^{104}\right)\mu \\ e_K^\top (I-\rho)^{-1}\left(I-\rho^{260}\right)\mu \end{bmatrix}, \quad H_S = \begin{bmatrix} H \\ 0 \\ e_K^\top \rho^{52} \\ e_K^\top \rho^{104} \\ e_K^\top \rho^{260} \end{bmatrix},$$

$$R_S = \begin{bmatrix} R & \underset{(N+R)\times 4}{0} \\ \underset{4\times(N+R)}{0} & \begin{bmatrix} 0 & 0 & 0 & 0 \\ 0 & \sigma_S & 0 & 0 \\ 0 & 0 & \sigma_S & 0 \\ 0 & 0 & 0 & \sigma_S \end{bmatrix} \end{bmatrix},$$

where $E_t(\pi_{t+\tau})$ is the τ-week ahead expected inflation from the Survey of Professional Forecasters, X_t is the same set of latent variables of model (10), σ_S is the variance of the forecasts and $\eta_{S,t} \sim N(0, I_{N+R+4})$. In this derivation equation (11) is directly inserted into the observation equation of model (10). Results are very similar to those presented above. Since for the United States the Survey of Professional Forecasters publish quarterly surveys for the one-year-ahead and 10-year-ahead inflation forecasts, the second last lines in model (12) are cancelled.

Short-term interest rate – A natural way to increase the identification power of the model is to use the short-term interest rate in the estimates. Model (10) is then estimated by inserting the 3-month repo interest rate in the Y matrix. The repo rate is preferred to the interbank rate and to the eurocurrency rate since it does not contain premia for counterparty risks. Results are very similar to those presented above.

8.1 Further research

Macroeconomic activity – A natural extension of the current framework would be to insert the real data on macroeconomic activity in a set up like that used by Pericoli and Taboga (2008); thus the state equation contains latent as well as observable variables. In this case, as in the case of survey data, the frequency of the macroeconomic variable is lower than that of the bond quotes, typically monthly or quarterly. The state equation in model (10) becomes $X_t = \mu + \rho X_{t-1} + hZ_{t-1} + \Sigma\varepsilon_t$, where Z is the monthly rate of growth of the industrial production index and h is a vector of parameters.

9 Conclusion

This paper presents a no-arbitrage affine Gaussian term structure model for the euro area and the United States. This model is capable of describing the evolution of the nominal and of the real term structure by means of a small number of latent factors which can be interpreted as two real-rate factors and one inflation factor. The model is also able to provide the spot long-term inflation expectation and the forward long-term inflation expectation implied in the nominal and real term structure together with the corresponding inflation risk premia. Inflation risk premia show large values and ample variability in the United States while they are smaller and more stable in the euro area.

Long-term expected forward inflation rates, a common indicator of inflation expectations, are on average below forward breakeven inflation rates in the United States, at around 2.1 per cent from 2002 until 2010; this implies that the forward inflation risk premium is on average positive in a range of 20 to 40 basis points in the United States; the forward inflation risk premia become sizable around the start of the late-2000s financial crisis and considerably increase in the United States just before the adoption of the first unconventional measures of monetary policy, known as QE1, in March 2009. In contrast, in the euro area expected forward inflation rates remain well anchored around 1.8 per cent and the forward inflation risk premium is unchanged even after the adoption of the unconventional monetary policy measures following the peaks of the financial crisis, in October 2008 and in May 2010.

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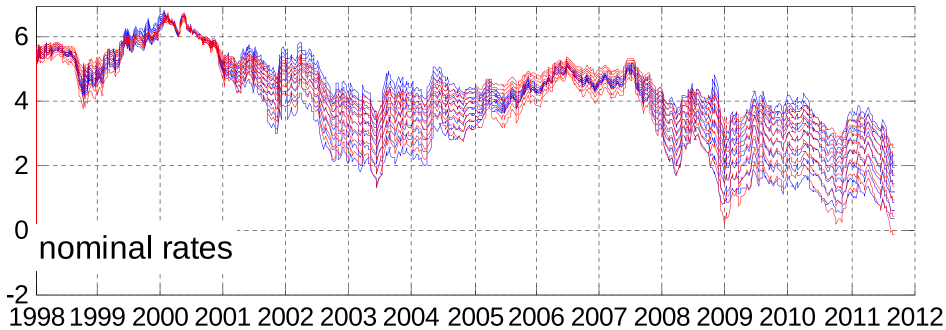
10 Tables and Figures

Figure 1 – United States: current (blue) and estimated (red) nominal and real rates



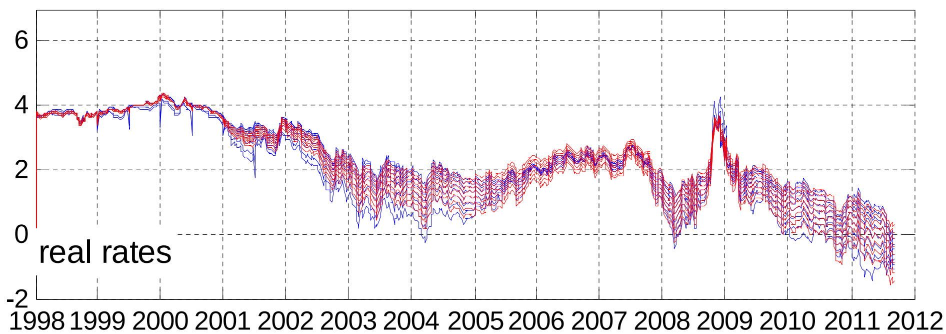


Figure 2 – Euro area: current (blue) and estimated (red) nominal and real rates

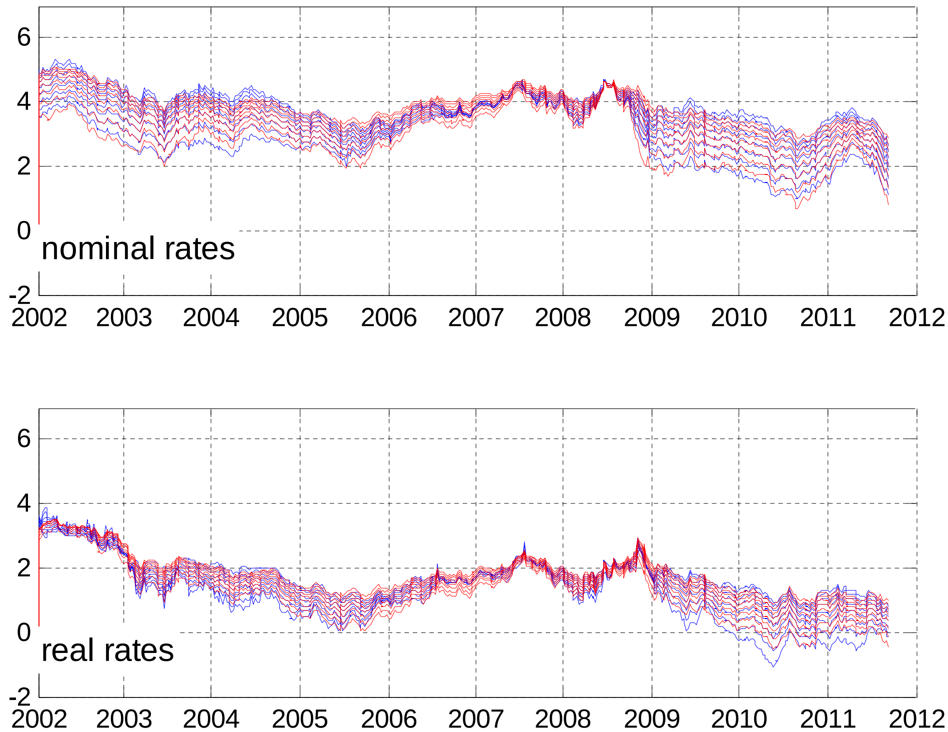


Figure 3 – United States: latent factors and observable variables

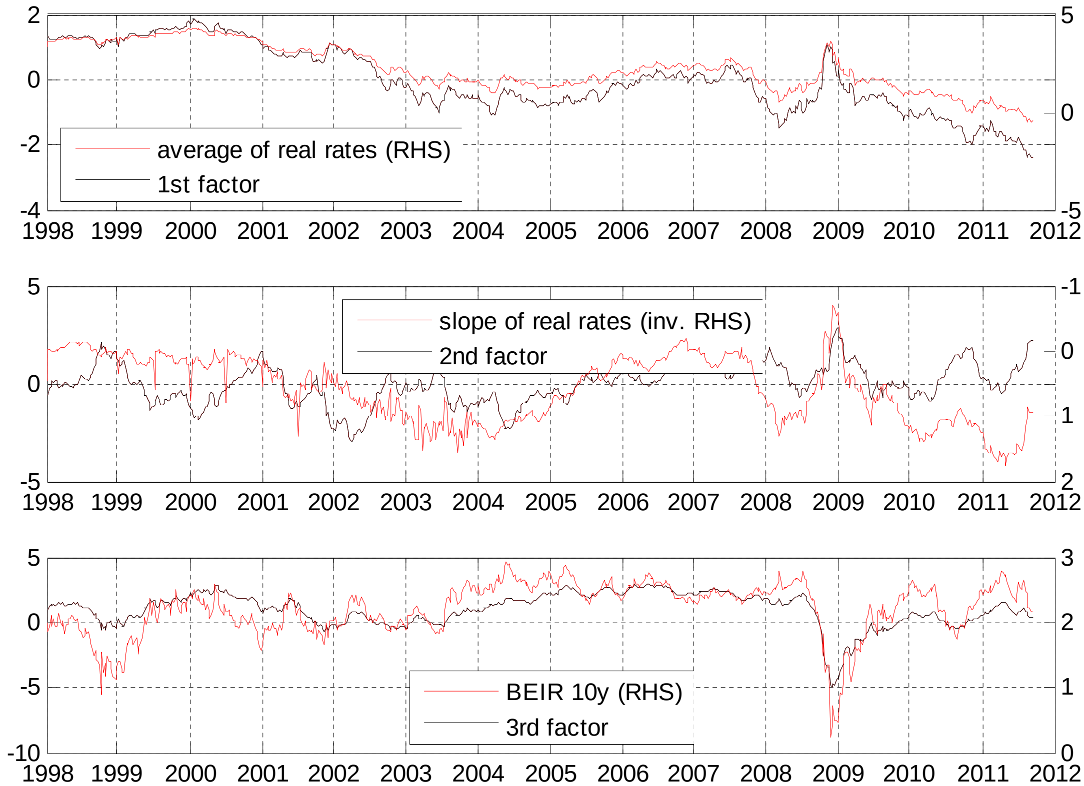


Figure 4 – Euro area: latent factors and observable variables

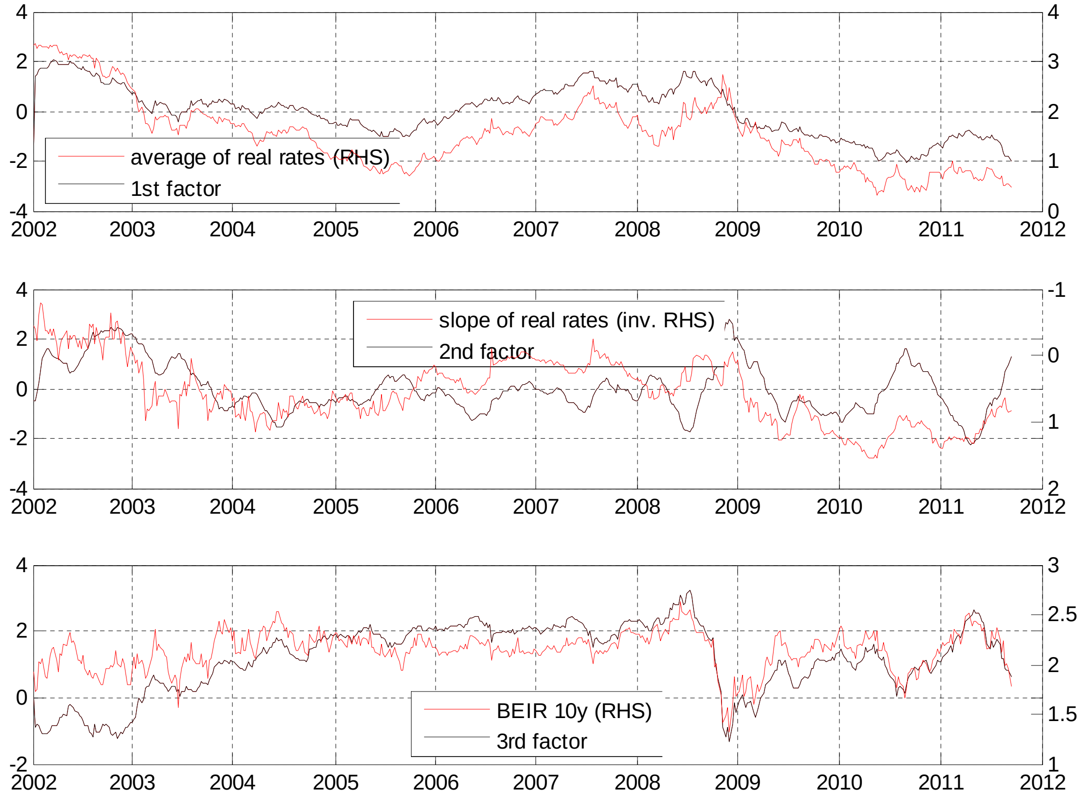
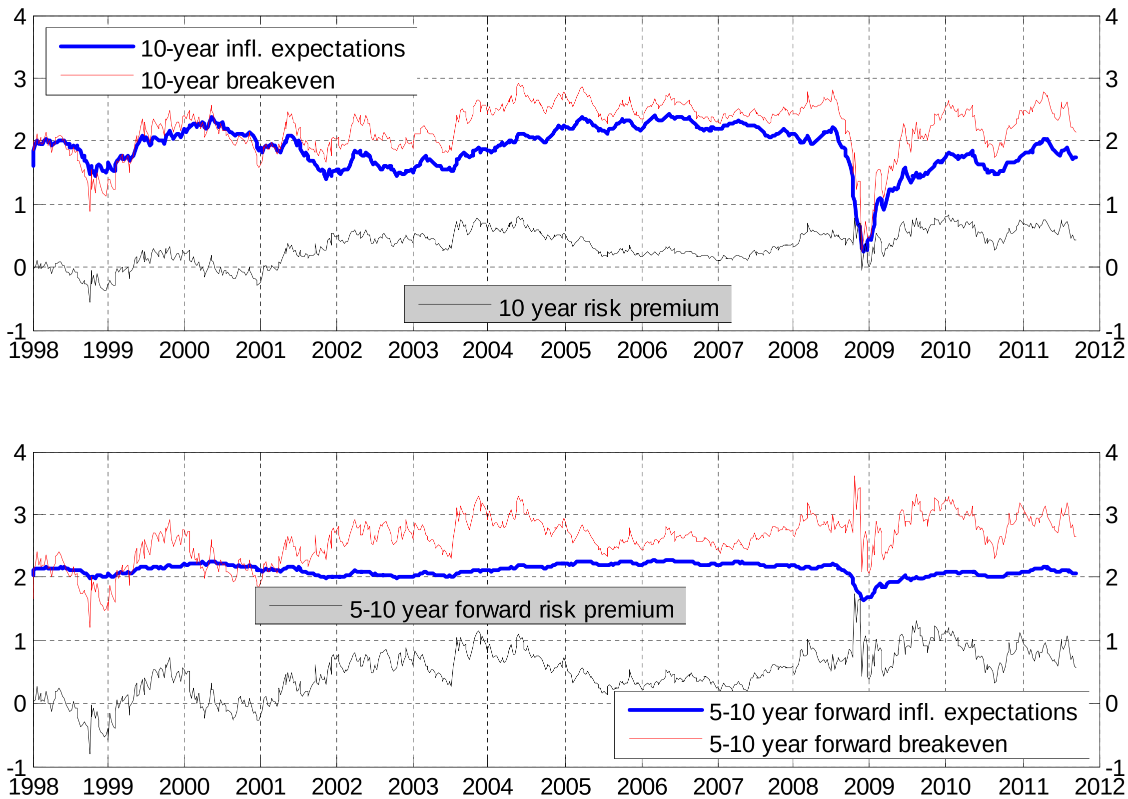


Figure 5 – United States: breakeven inflation rates, expected inflation rates and risk premia



The 10-year expected inflation rate is given by equation (11), namely

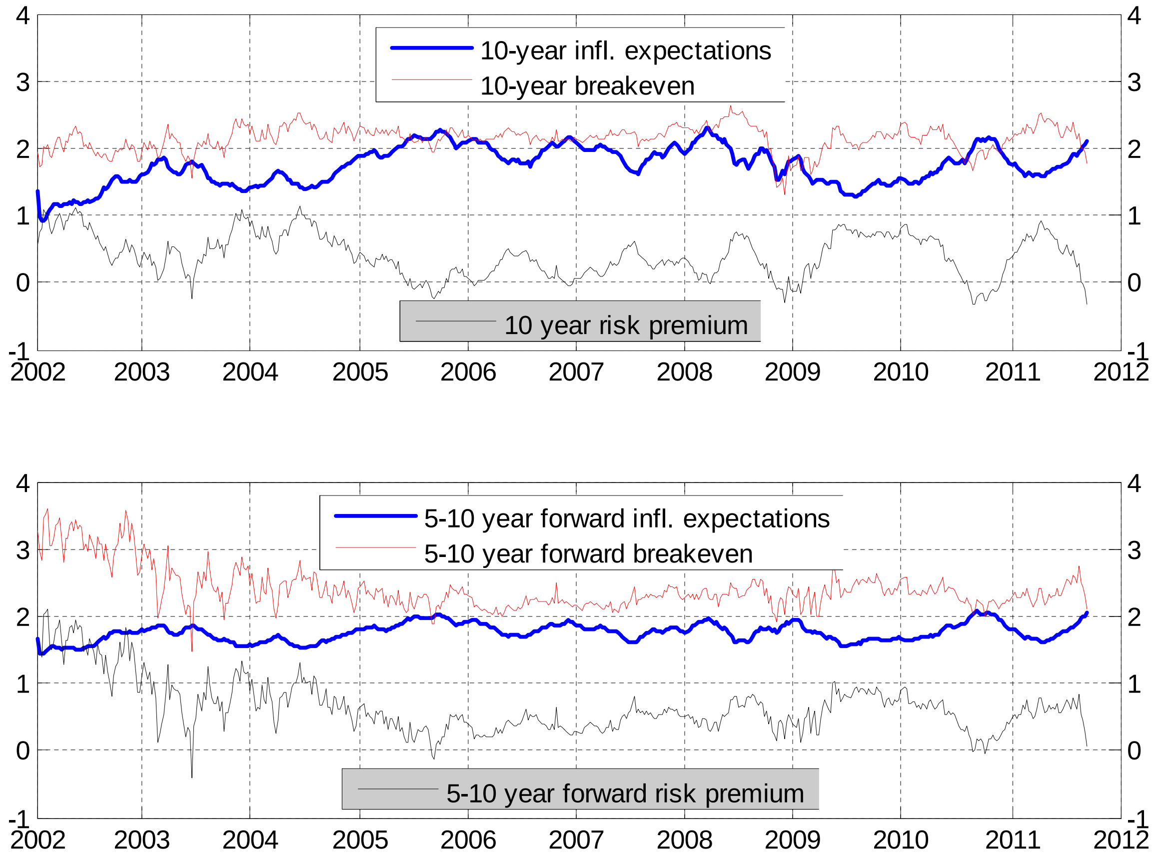
$$\frac{1}{520}E_t(e_K^\top \widehat{X}_{t+520}) = [0\ 0\ 1] \cdot \left[(I - \widehat{\rho})^{-1} \left(I - \widehat{\rho}^{520} \right) \widehat{\mu} + \widehat{\rho}^{520} \cdot \widehat{X}_t \right]$$

where 520 is the 10-year forecasting period in weeks and $\widehat{}$ above a parameter stands for its estimate. The 5-10 year forward expected inflation is given by

$$2 \times \frac{1}{520}E_t(e_K^\top \widehat{X}_{t+520}) - \frac{1}{260}E_t(e_K^\top \widehat{X}_{t+260})$$

where 260 is the 5-year forecasting period in weeks. The 10-year inflation risk premium is the difference between the 10-year breakeven inflation rate and (11); the 5-10 year forward risk premium is the difference between the 5-10 year forward breakeven inflation rate and the 5-10 year forward expected inflation.

Figure 6 – Euro area: breakeven inflation rates, expected inflation rates and risk premia



The 10-year expected inflation rate is given by equation (11), namely

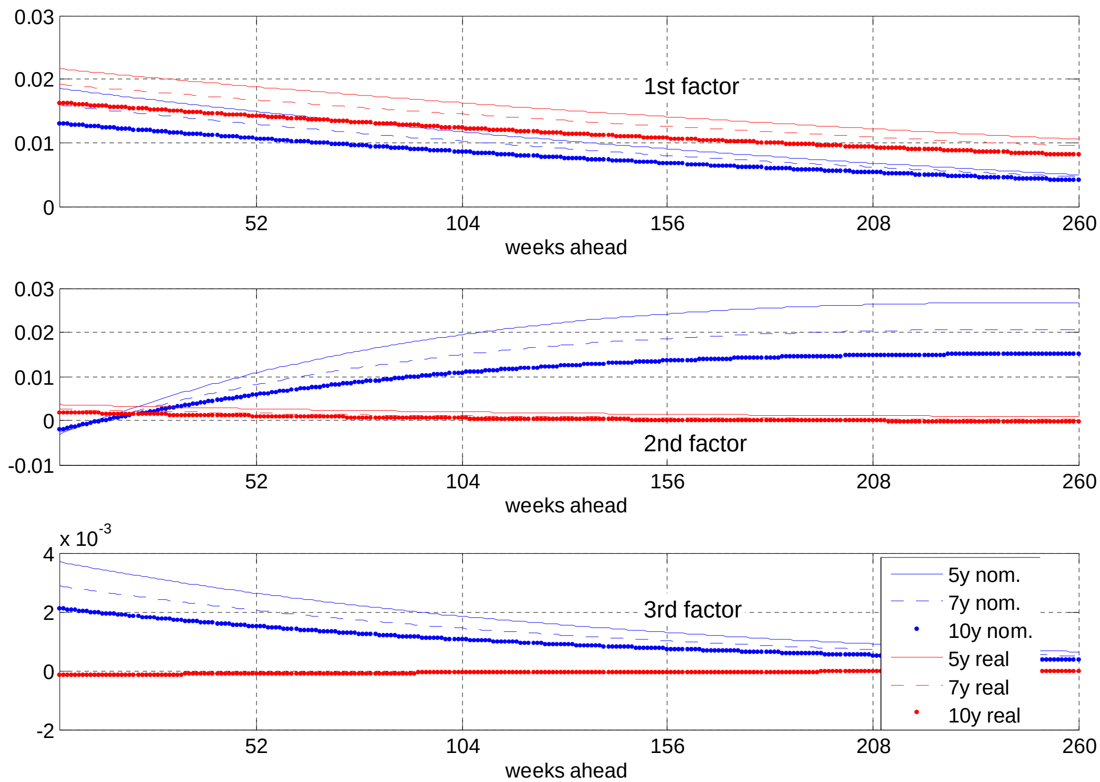
$$\frac{1}{520}E_t(e_K^\top \widehat{X}_{t+520}) = [0\ 0\ 1] \cdot \left[(I-\widehat{\rho})^{-1} \left(I-\widehat{\rho}^{520}\right)\widehat{\mu} + \widehat{\rho}^{520}\cdot \widehat{X}_t \right]$$

where 520 is the 10-year forecasting period in weeks and $\widehat{}$ above a parameter stands for its estimate. The 5-10 year forward expected inflation is given by

$$2 \times \frac{1}{520}E_t(e_K^\top \widehat{X}_{t+520}) - \frac{1}{260}E_t(e_K^\top \widehat{X}_{t+260})$$

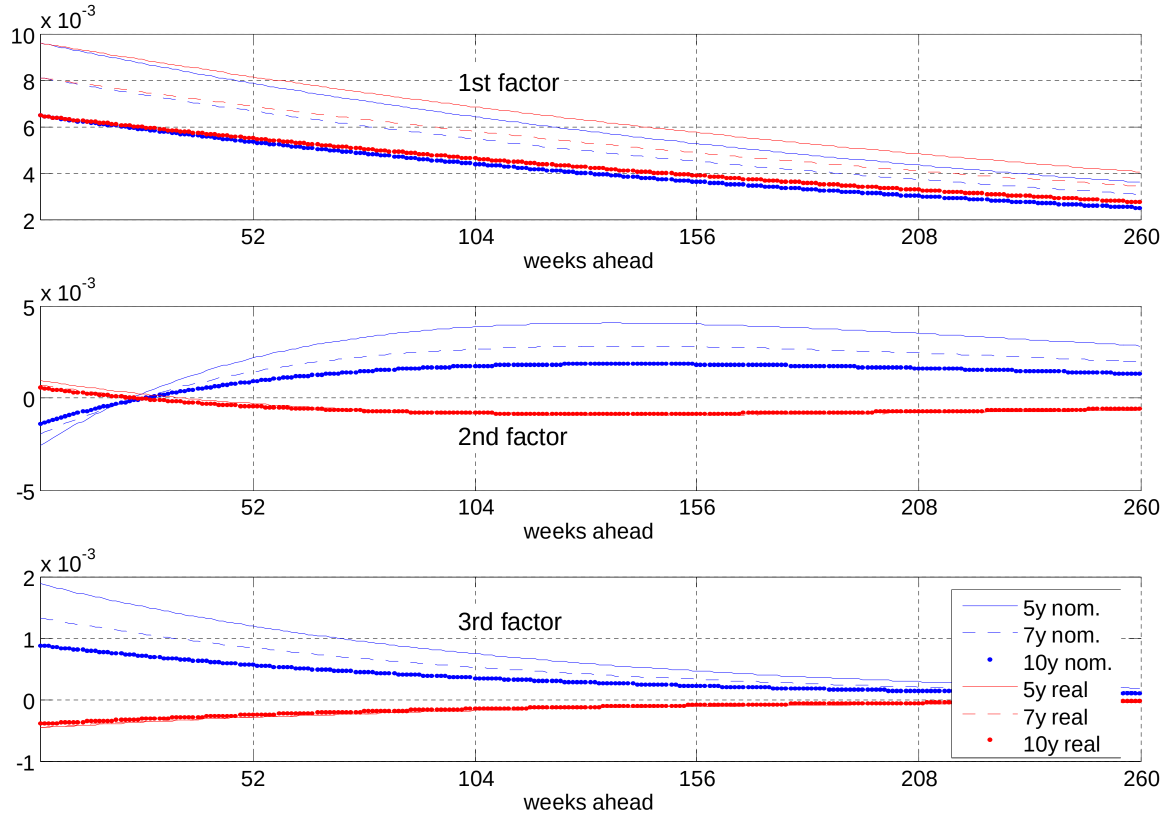
where 260 is the 5-year forecasting period in weeks. The 10-year inflation risk premium is the difference between the 10-year breakeven inflation rate and (11); the 5-10 year forward risk premium is the difference between the 5-10 year forward breakeven inflation rate and the 5-10 year forward expected inflation.

Figure 7 – United States: impact of one standard deviation shocks on state variables



The Figures show the expected path of the 5-year, 7-year, 10-year nominal and real rates following a one standard deviation move in each of the state variables.

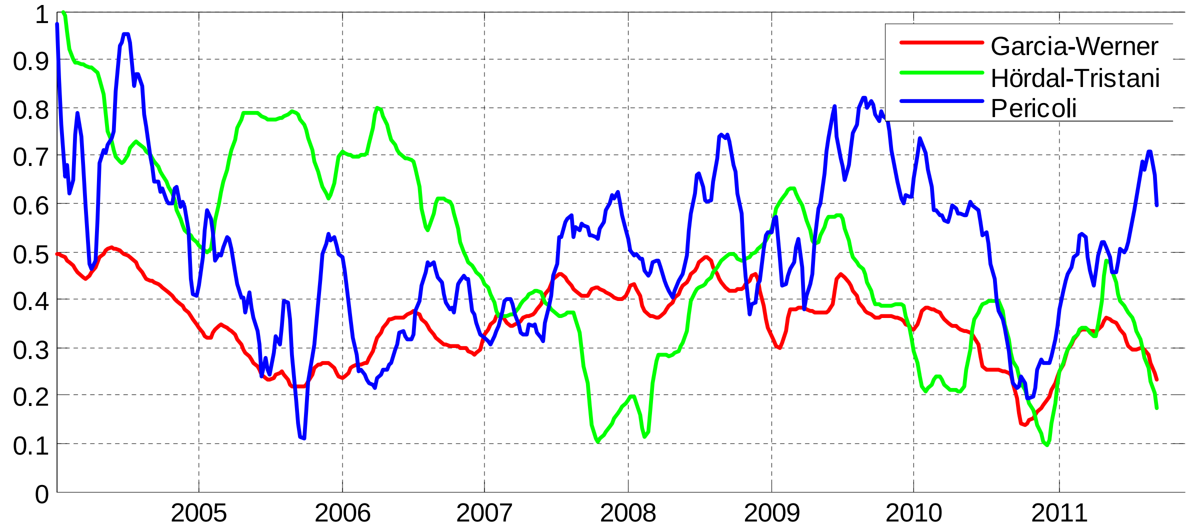
Figure 8 – Euro area: impact of one standard deviation shocks on state variables



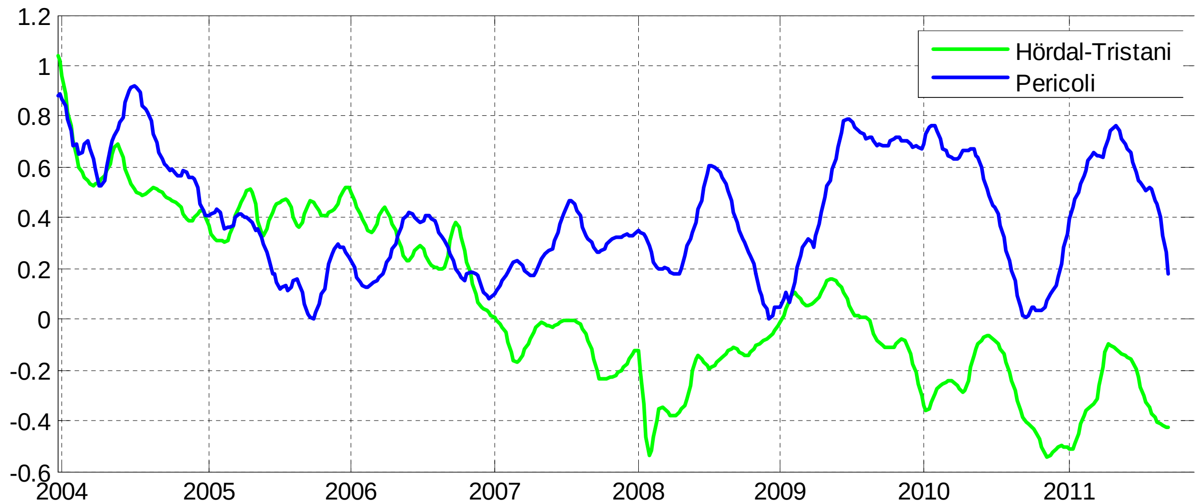
The Figures show the expected path of the 5-year, 7-year, 10-year nominal and real rates following a one standard deviation move in each of the state variables.

Figure 9 – Inflation risk premia in the euro area in different models

5y-5y forward



10y



The Figures report the inflation risk premia as reported in the papers by Garcia and Werner (2010), Hördal and Tristani (2010) and by this paper; the data of Garcia and Werner (2010) and Hördal and Tristani (2010) are transformed to weekly data through a polynomial interpolation of the monthly original data where each original data point is set in the last week of the corresponding month. For comparison, this paper's data are represented as a 5-week moving average.

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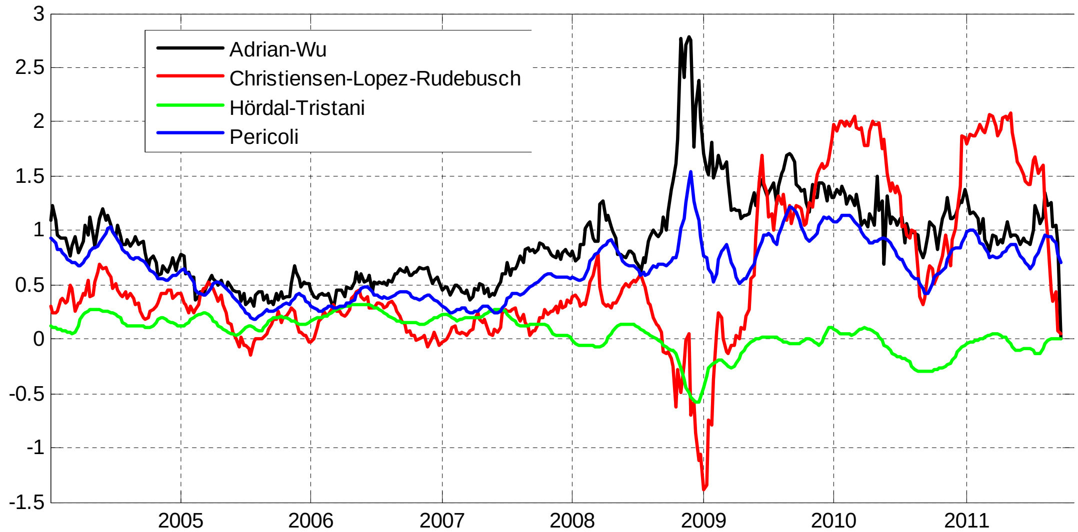
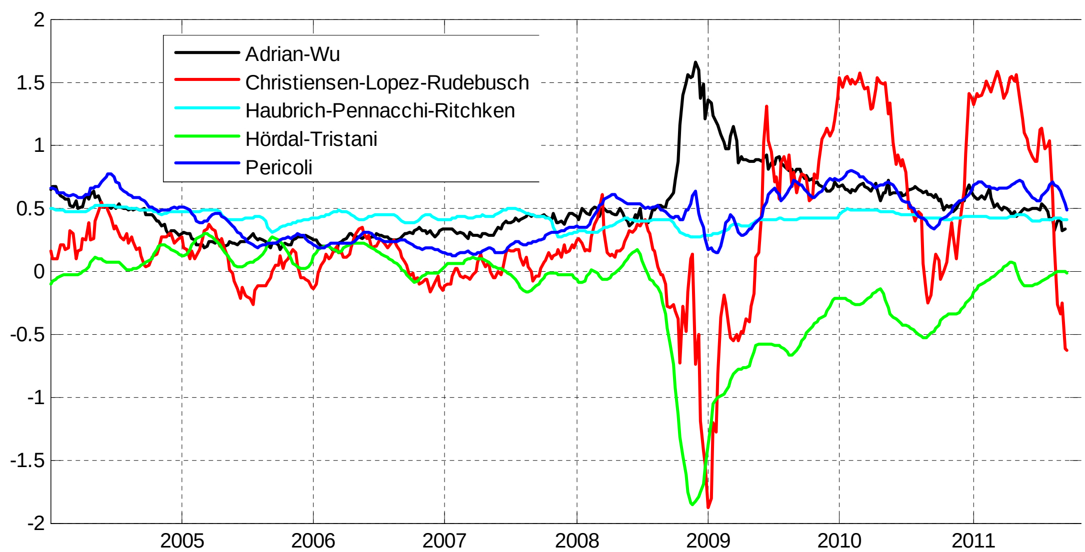
Figure 10 – Inflation risk premia in the United States in different models

5y-5y forward



10y



The Figures report the inflation risk premia as reported in the papers by Adrian and Wu (2010), Christensen, Lopez and Rudebusch (2010), Haubrich, Pennacchi and Ritchken (2011a), Hördal and Tristani (2010) and by this paper; the data of Haubrich, Pennacchi and Ritchken (2011a) and Hördal and Tristani (2010) are transformed to weekly data through a polynomial interpolation of the monthly original data where each original data point is set in the last week of the corresponding month. For comparison, data of this paper are represented as a 5-week moving average; data of Adrian and Wu (2010) and Christensen, Lopez and Rudebusch (2010) are represented in their original weekly frequency.

Table 1a – United States: yield pricing errors

	nominal			real		
	mean	median	std. dev.	mean	median	std. dev.
3-yr	-0.01	-0.01	0.04	-0.30	-0.03	9.96
4-yr	-0.01	-0.01	0.04	0.01	-0.01	0.76
5-yr	-0.01	-0.00	0.03	0.01	0.00	0.15
6-yr	-0.01	-0.00	0.02	-0.03	0.01	1.02
7-yr	-0.00	-0.00	0.02	-0.01	0.00	0.07
8-yr	-0.00	-0.01	0.02	-0.01	-0.00	0.06
9-yr	-0.00	-0.00	0.02	-0.02	-0.01	0.06
10-yr	0.00	0.00	0.02	-0.02	-0.02	0.06

Note: statistics of weekly data from 2 January 1998 to 25 August 2011. Pricing error is defined as the percentage difference between the current and the estimated yield.

Table 1b – Euro area: yield pricing errors

	nominal			real		
	mean	median	std. dev.	mean	median	std. dev.
3-yr	-0.03	-0.02	0.03	0.02	-0.01	0.97
4-yr	-0.03	-0.02	0.02	-0.09	-0.02	1.74
5-yr	-0.02	-0.02	0.02	-0.03	-0.01	0.11
6-yr	-0.02	-0.02	0.02	-0.00	-0.00	0.04
7-yr	-0.01	-0.01	0.02	0.00	0.00	0.04
8-yr	-0.01	-0.01	0.02	0.01	0.01	0.05
9-yr	-0.01	-0.01	0.02	0.00	0.01	0.05
10-yr	-0.01	-0.01	0.02	0.00	0.01	0.05

Note: statistics of weekly data from 4 January 2002 to 25 August 2011. Pricing error is defined as the percentage difference between the current and the estimated yield.

Table 2 – Parameter estimates

US	coefficient	std. err.	euro area	coefficient	std. err.
ρ_{11}	0.998	0.304	ρ_{11}	0.989	0.134
ρ_{21}	-0.004	0.151	ρ_{21}	-0.001	0.340
ρ_{22}	0.999	0.027	ρ_{22}	0.996	1.064
ρ_{31}	-0.000	0.086	ρ_{31}	0.008	0.028
ρ_{32}	-0.000	0.013	ρ_{32}	-0.007	0.061
ρ_{33}	0.993	0.008	ρ_{33}	0.989	0.005
μ_π	0.030	48.341	μ_π	0.060	39.359
σ_π	-0.268	0.494	σ_π	-0.114	0.086
δ_0	-40.895	5.439	δ_0	-72.063	48.945
$\delta_{1,1}$	0.246	216.415	$\delta_{1,1}$	0.133	0.744
$\delta_{1,2}$	-66.370	56.304	$\delta_{1,2}$	-122.733	312.417
$\delta_{1,3}$	12.352	0.000	$\delta_{1,3}$	-18.485	0.000
$\lambda_{0,1}$	0.016	4.205	$\lambda_{0,1}$	-0.025	3.918
$\lambda_{0,2}$	1.372	1.181	$\lambda_{0,2}$	1.173	1.320
$\lambda_{0,3}$	0.023	13.151	$\lambda_{0,3}$	0.008	40.025
$\lambda_{1,11}$	0.006	0.307	$\lambda_{1,11}$	-0.008	0.134
$\lambda_{1,12}$	1.996	0.026	$\lambda_{1,22}$	1.940	2.374
$\lambda_{1,13}$	6.930	0.000	$\lambda_{1,33}$	172.666	0.000
$\lambda_{1,21}$	0.284	0.298	$\lambda_{1,12}$	-0.005	0.005
$\lambda_{1,22}$	-0.000	3.890	$\lambda_{1,21}$	-0.161	6.228
$\lambda_{1,23}$	-0.002	16.362	$\lambda_{1,13}$	-0.000	0.002
$\lambda_{1,31}$	0.091	0.025	$\lambda_{1,23}$	0.042	0.680
$\lambda_{1,32}$	0.003	6.492	$\lambda_{1,31}$	-1.424	6.602
$\lambda_{1,33}$	-0.006	6.087	$\lambda_{1,32}$	-0.021	4.867
c_N	0.001	0.075	c_N	0.089	0.151
d_N	0.023	1.124	c_R	0.010	0.645
c_R	0.015	1.192	d_N	0.046	0.118
d_R	0.005	1.035	d_R	0.049	0.053

Table 3 – Correlation of latent factors with observable variables

euro area	2002-2011			2007-2011		
	1st	2nd	3rd	1st	2nd	3rd
average of real rates	0.82	–	–	0.88	–	–
slope of real rates	–	0.53	–	–	0.55	–
10-year BEIR	–	–	0.69	–	–	0.85
United States	**1998-2011**			**2007-2011**		
	1st	2nd	3rd	1st	2nd	3rd
average of real rates	0.75	–	–	0.89	–	–
slope of real rates.	–	0.27	–	–	0.48	–
10-year BEIR	–	–	0.76	–	–	0.91

The Table reports the absolute value of the correlation between the three latent factors, defined in column as 1st, 2nd and 3rd, and the observable variables, defined as the average of real rates, the slope of real rates – the difference between the 10-year and the 3-year real rate – and the 10-year BEIR.

Table 4 – Comparison of inflation risk premia

	2004-2011		2004-2007		2008-2011	
	5y-5y	10y	5y-5y	10y	5y-5y	10y
euro area	mean	mean	mean	mean	mean	mean
Garcia Werner	0.38	–	0.43	–	0.35	–
Hördal-Tristani	0.55	0.16	0.76	0.50	0.40	0.12
Pericoli	0.63	0.44	0.69	0.44	0.54	0.44
United States						
Adrian-Wu	0.95	0.63	0.84	0.61	1.19	0.67
Christensen et al.	0.50	0.25	0.18	0.04	1.20	0.70
Haubrich et al.	–	0.42	–	0.45	–	0.40
Hördal-Tristani	0.05	-0.17	0.15	0.01	-0.03	-0.32
Pericoli	0.65	0.45	0.57	0.39	0.84	0.58

The Table reports the averages of inflation risk premia in percentage points as reported in the papers by Adrian and Wu (2010), Christensen, Lopez and Rudebusch (2010), Haubrich, Pennacchi and Ritchken (2011a), Hördal and Tristani (2010) and by this paper; the data of Garcia and Werner (2010), Haubrich, Pennacchi and Ritchken (2011) and Hördal and Tristani (2010) are monthly, those of Adrian and Wu (2010), Christensen, Lopez and Rudebusch (2010) and Pericoli are weekly. 5y-5y stands for 5-year inflation risk premium 5 years ahead; 10y stands for 10-year inflation risk premium.

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2012

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