Expected inflation, inflation risk premium and the term structure of macroeconomic announcements in the euro area and in the United States

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Abstract

In the first part this paper uses the celebrated no-arbitrage affine Gaussian term structure model applied to index-linked and standard government bonds to obtain expected inflation rates and the corresponding inflation risk premia, in the euro area and in the United States. After estimating the model using maximum likelihood, results also show that the model performs surprisingly well in generating a number of dynamic features of both index-linked and standard bonds. Results show that the model describes the evolution of the nominal and real term structures by mean of a small number of latent factors which can be interpreted as two real factors and one inflation factor and carry substantial information related to expected inflation and inflation risk premia in the euro area and in the United States. In the second part, the paper analyses the impact of macroeconomics news on the nominal and real term structure.

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Part I
The term structure model

1 Introduction

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 extrapolate inflation expectations at different maturities. In fact, bonds linked to an inflation index differ from the corresponding standard bonds as to expected inflation and inflation risk premium as well as to maturities, coupon rates and cash-flow structures. In addition, since index-linked bonds have different maturities, an entire spectrum of inflation expectations and inflation risk premia can be derived from the 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 numbers of common factors; the novelty of this stream of literature, which this paper belongs to, 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 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 the opportunity to divide a model-implied constant-maturity inflation compensation (or model-implied break-even inflation rate), obtained as 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).

Break-even inflation rates are usually taken as proxies for inflation expectations and provide a measure of central bank credibility about 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. Inflation compensation provides a measure of inflation expectation 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 unorthodox monetary policy recorded over the past years has raised concerns about the possibility that market participants were still seeing central bank policy as consistent with longer-run price stability.

The use of a model which jointly estimates the expected inflation and the inflation risk premium presents two advantages with respect to use of the plain-vanilla break-even inflation rates computed as difference between nominal and real interest rates. First, over longer time horizons, the break-even inflation rate can substantially differ from the expected inflation since the compensation requested by investors for uncertainty about future inflation rates – i.e. the inflation premium – can be relevant. Second, real and nominal interest rates are estimated on the base 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 economic intuition of the drivers of the nominal and real term structures. Third, being expected inflation a key ingredient in the decision of monetary policy 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 that given by the Survey of Professional Forecasters (released quarterly in the United States and in the euro area) and by Consensus Economics (released twice a year).

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, 2011).

This paper innovates with respect to the previous literature in two fields. First, it is the first work which uses weekly data for the euro area; previous works with weekly data are those of Risa (2001) for the United Kingdom and of Adrian and Wu (2010) for the United States. The use of weekly data is fundamental when this class of model is used to evaluate inflation expectations and inflation risk premia by monetary policy makers. Results show that i) long-term inflation risk premia are relevant in the United States and smaller in the euro area; ii) these premia present a countercyclical dynamics and behave very similarly to other indicators of financial market risk; iii) the three latent factors, which drives the nominal and real term structures, can be interpreted as the cross section average of real interest rates, the 10-year real interest rate and the breakeven inflation rate.

Moreover, the second part of this paper analyses the impact of surprises from macroeconomic news on the nominal and real term structure, evaluate the factor loadings associated to each news and their impulse response function. When the surprise of macroeocnomic announcements are plugged into the term structure model, it is possible to estimate the impact of unexpected news onto the entire term structure, both nominal and real, which is consistent with the absence of arbitrage in the bond markets.

The paper is organized as follows. Section 2 reviews the literature. Section 3 presents the data. The model is presented in Section 4. Estimates of the real term structure are presented in Section 5. Section 6 concludes.

2 The literature

Recent papers on the term structure of inflation can be divided in 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 some identification assumption to increase the power of the esimates; along this line of research there are Evans (1998), Risa (2000), Ang et al. (2008), Christensen et al. (2010), Garcia and Werner (2010) and Adrian and Wu (2010). Alternatively, a second group of works uses standard new keynesian macrofinance model which encompasses financial and macro variables in a framework à la Ang and Piazzesi (2003); the work by Hördal and Tristani (2010) is along this line of research. This paper belongs to the first class of papers.

Evans (1998) and Risa (2000) 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 premium 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, have been close to zero on average.

Garcia and Werner (2010) document that no-arbitrage Gaussian affine term structure models 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.

Among the second set of works, Hördal and Tristani (2010) extend their new keynesian macro-finance model by encompassing the nominal and the real term structure and introduce data and 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 indexed-linled 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, $\pi_t^{e,n}$, 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 $\gamma_t^n = Cov(m_t^n, \pi_t^n) - \frac{1}{2}Var(\pi_t^n)$, where m is the stochastic discount factor and π the inflation rate; 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, coming out from a Jensen inequality. Sometimes, the first term of the inflation risk premium, $Cov(m_t^n, \pi_t^n)$, is referred to as the 'pure inflation risk premium'.

The inflation risk premium, i.e. the compensation for risk due to uncertainty of future inflation, can be evaluated by mean 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 in 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 of the inflation component. This model follows the original setup by Evans (1998), successively enriched 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 strictly lie 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 cash at time $t+n$ (denoted by p_t^n) equals:

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

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 λ_t:

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

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 relevant economic variables. This is one of the reasons why recent studies have proposed to augment 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 into the models: recent examples are Ang and Piazzesi (2003), Rudebusch and Wu (2005), Hördal, Tristani and Vestin (2005) and Ang et al. (2005). All these works impose some 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 nominal and real rates implied in bonds. Thus, this class of models do not consider the inflation rate among its state variables.

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

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

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 (**??**)[1]:

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

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 to the fluctuations of the price of a bond expiring in n periods. Such premia are in general time varying, and they are constant only when $\rho = \overline{\rho}$.

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

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

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

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, 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.

Without loss of generality, let's assume that ρ il lower triangular and that the matrix Σ is diagonal with all diagonal elements equal to one but the last, namely

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

the 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 non decreasing in the term to maturity τ, i.e. the volatility is higher 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_N are equal to zero the price errors are constant across maturities (Risa, 2000).

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

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

6 Model specification

The complete model is defined by the following parameters

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

7 The data

Nominal and real term structures for the euro area are estimated from daily quotes of French government bonds by mean of the methodology first introduced by Fisher at al. (1996). The nominal term structure is estimated by using the quotes of the euro repo rates with maturity 1 week, 2 weeks, 3 weeks, 1 month, 2 months, 3 months, 6 months, 9 months, 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 *Obligations Assimilables au Trésor* (OAT) with maturity greater than 1 year.

The real term structure for the euro area is estimated by using OAT€i, i.e. OAT indexed to the euro-area harmonised index of consumer prices (HICP) ex-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, indexed to the domestic French Consumer Price Index (CPI). In 2002 there was the first issue of French government bonds, OAT€i, indexed to the euro-area HICP ex-tobacco. This work considers only French index-linked bonds since they are given an AAA rating, against the AA+ rating given to the Italian government securities.

Daily 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 (http://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, we compare my estimated difference between off-the-run and on-the-run interest rates with the difference 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. (2006) and Gürkaynak et al. (2007).[2]

8 Results

Results show that model (10) is capable to jointly estimate 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 follow quote closely the nominal as well as real interest rates; only at the shortest maturities the yield pricing errors are large.

In both areas the latent factors can be interpreted as the cross-sectional average of the real term structure (1st factor), the 10-year real interest rate (2nd factor) and the 10-year breakeven inflation rate (3rd factor); this latter is a wedge between the nominal and the real interest rate (Figures 3-4).

The model gives an estimate of the spot 10-year inflation risk premia (top panel of Figures 5 and 6) and the 5-10 year forward inflation risk premia, defined as the difference between the corresponding breakeven inflation rates and inflation expectations computed by equation (11). 10-year inflation risk premia are quite small, generally positive but on average between 10 and 40 basis points, with the exception of the last quarter of 2008 when the bond market froze. 5-10 year forward inflation risk premia are much more volatile due to the stability of the 5-10 year forward inflation expectation; in the United States the average was around zero from 1998 to 2006 while it recorded positive value since 2007. It has reached a peak of almost 2 per cent at the end of 2008 and since then has remained above 50 basis points. Conversely, in the euro area the 10-year inflation risk premium has been stable around 30 basis points, while the 5-10 year forward inflation risk premium has averaged around zero.

The comparison between forward expected inflation and forward risk premia in the euro area and in the United States gives evidence to some differences in the two areas. In the euro area, forward expected inflation has remained well anchored to the targeted level, 2 per cent, and has scarcely changed after the introduction of the unconventional monetary policy measures introduced in aftermath of the financial crisis (the change in the eligible collateral in the refinancing operations, the sterilized purchase of government bonds and some others). In the United States, the waves of unconventional monetary policy measures, defined quantitative easing, has changed the perception of expected inflation by market participants.

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

8.1 Impulse response function

The impulse response function of model (10) are shown in figures 7 and 8. They show that shocks to the first latent factor (which is proxied by the cross-section average of real rates) provoke an increase to the real rates and to the nominal rates. As far as real rates are concerned, the response is larger for shortest maturity rates while is larger for longer maturity rates for nominal rates. The rates tend to converge to their long-term averages both in the euro area and in the United States; however in the United States the response to the shock is very similar among nominal and real rates.

The shock to the second latent factor (which is proxied by the ten-year real rate) has very different impact in the euro area with respect to the United States. In the former market nominal rates decrease while real rates increase; conversely in the United States both nominal and real rate decrease.

The shock to the third factor (which is proxied by the long-term breakeven inflation) show that in both markets real rates decrease and nominal rates increase and all converge with the same pace towards zero. As expected, a shock to the breakeven inflation tends to increase nominal rates which are not sheltered by increase in the inflation rate.

9 Robustness

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. In the case of 2 factors, the fit was not able to capture the dynamics of the term structure; the unique real latent factor is not capable to point out the cross-sectional dispersion among interest rates. With four factors, the model overfits the term structures both for real and nominal interest rates.

Surveys of inflation expectation – Moreover, the surveys of inflation expectation are introduced in the model in order to improve its identification power, as in Garcia and Werner (2010) and in D'Amico et al. (2008). Alternatively, Adrian and Wu (2010) use time-varying conditional covariation between real and inflation factors to increase the identification power. The model (10) has been estimated with the quarterly surveys published by the Survey pf Professional Forecasters for the United States and for the euro area. Results are very similar to those presented. The observation equation in the model (10) becomes

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

where

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

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

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.

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. The 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 premiums for counterparty risks. Results are very similar to those presented above.

Macroeconomic activity – A natural extension 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 of quarterly. The state equation in the 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. Results for this robustness check are still preliminary.

10 Conclusion

The 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 mean of a small number of latent factors which can be interpreted as two real factors and one inflation factor. The model is capable to give the spot long-term inflation expectation and the forward long-term inflation expectation implied in the nominal and real term structure and the corresponding inflation risk premia.

Long-term expected inflation in the euro area and in the United States is on average below breakeven inflation; this suggests that inflation risk premia is on average positive and between 10 and 40 basis points. Forward inflation risk premia are larger than long-term premia and record large figures around the recent episodes of tension in the last financial crisis.

Part II

The term structure of macroeconomic news

VERY PRELIMINARY DRAFT, PLEASE DO NOT QUOTE

11 Macroeconomic news release

This Section of the paper considers the simple model (10) and adds to it the surprises coming out from the macroeconomic data releases. Thus the model becomes a state-space system with unobservable and observable variables and can be treated according to the specification of Pericoli and Taboga (2008).

12 The data for macroeconomic news

Surprises are defined as an unexpected release with respect to the median forecasts released by Bloomberg; they are built such that good news provoke an increase in long-term yields, for instance a decrease in the unemployment rate provokes an increase in long-term yields, and conversely negative news provoke a decrease in long-term yields. This paper uses the first estimates of the macroeconomic news even if most of the macroeconomic figures released in the euro area and in the United States are initially preliminary estimates and on the next announcement for the same figure, a revised estimate of the preceding figure is released. Most of the macroeconomic datasets used in empirical papers are made of the revised estimates of every macroeconomic figures.[3]

Let $S_{t,i}$ denote the surprise at time t in the figures indexed by i as follows:

$$S_{t,i} = \frac{R_{t,i} - F_{t,i}}{\sigma_{S_i}} \ , \tag{12}$$

where $F_{t,i}$ is the market median consensus about the upcoming figures i for t, the date of release; $R_{t,i}$ is the announcement (the first estimate) at time t of the same figure i. To make surprises comparable, they are scaled using their historical standard deviation, σ_{S_i}. This way of proceeding is very common, see e.g. Fleming and Remolona (1999).The Bloomberg median survey forecasts as a measure of the market consensus for a given figure at a given date; thus, $F_{t,i}$ will be proxyed by the last forecast in the Bloomberg database for each announcement.

[3] What is more, the Bloomberg calendar also contains the Bloomberg forecasts regarding each of these figures. Bloomberg forecasts are formed using the fiftieth empirical percentile of the distribution of a survey made of the forecasts of several bank economists, regarding a precise figure. The use of the median as a measure of the expectations makes the forecast robust to the influence of badly intentionned economists that would want to shift the forecast in order to make the most of it. What is more, this forecast is extensively used by market participants. For each figure that is predicted by Bloomberg's collection of economists' forecasts, the median is regularly updated until every economist answers the survey, which can take up to two weeks. We retained the last median computed by the Bloomberg services, so as to match both the practioners and academic ways of doing things. Some of the eliminated series were discarded because there was no available forecast.

Table 3 presents the macroeconomic news most followedby bond traders in the euro area and in the United States. However the list is not exhaustive since some important news have been excluded since they have been discontinued in the sample period. News can be grouped in four main categories: news on future economic activity (such as consumer confidence indices, purchasing manager indices etc.), news on inflation (consumer price and producer price indices), news on growth (GDP rate of growth, industrial production rate of growth, new orders etc.) and, finally, news on the labor market (jobless claims, unemployment rate etc.).

13 The model with macroeconomic news

The model which encompasses macroeconomic news is obtained by (10) by adding a new set of variable in the state equation, namely

$$\begin{array}{rcll} Y_t & = & \overline{A} + \overline{H}\overline{X}_t + R\eta_t & \text{(observation equation)} \\ \overline{X}_t & = & \overline{\mu} + \overline{\rho}\overline{X}_{t-1} + \overline{\Sigma}\varepsilon_t & \text{(state equation)} \\ R & \perp & \Sigma \, , & \end{array} \tag{13}$$

where the usual notation holds and

$$\begin{array}{rcl} \overline{A} & = & \left[A + \left[\widehat{E}, E\right]^\top\right], \overline{H} = \begin{bmatrix} H & 0 \\ 0 & \begin{bmatrix} \widehat{G} & 0 \\ 0 & G \end{bmatrix} \end{bmatrix}, \overline{\rho} = \begin{bmatrix} \rho & 0 \\ 0 & \phi \end{bmatrix}, \\ \overline{\mu} & = & [\mu, \mu_S]^\top, \overline{X} = [X, S], \overline{\Sigma} = \begin{bmatrix} \Sigma & 0 \\ 0 & \Sigma_S \end{bmatrix}. \end{array}$$

Where X is the usual set of K latent factos, S is the set of M known surprises (a M-dimensional vector, $M \in \mathbb{N}$), ϕ is a $M \times M$ diagonal, Σ_S is a $M \times M$ diagonal matrix, μ_S is a $M \times 1$ vector of drifts, $\widehat{E}$ and E are the $(N+R) \times 1$ vector of drifts for the nominal and real rates in the observation equation associated to factors S, $\widehat{G}$ and G are $(N+R) \times M$ matrix of loadings for nominal and real rates in the observation equation associated to factors S.

The matrix $S = [s_1, s_2, ..., s_M]$ is made by the time-series vector of surprises which are equal to zero when there is no release and equal to the surprise in the week when the news is releases, namely

$$S = \begin{bmatrix} s_{1,1} & 0 & \cdots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & s_{2,t-1} & \cdots & 0 \\ s_{1,t} & 0 & \cdots & 0 \\ 0 & 0 & \cdots & s_{M,t+1} \\ \vdots & \vdots & \ddots & \vdots \\ 0 & s_{2,T} & \cdots & 0 \end{bmatrix}.$$

The drifts, E, and the loadings, G, for real rates associated with surprises S are defined by the usual system of Riccati equations

$$
\begin{aligned}
E_1 &= -\gamma_0 \\
G_1 &= -\gamma_1 \\
& \dots \\
& \dots \\
E_{n+1} &= -\gamma_0 + E_n + G_n^{\intercal}(\mu_S - \Sigma_S\theta_0) - \frac{1}{2}G_n^{\intercal}\Sigma_S^{\intercal}\Sigma_S G_n \\
G_{n+1} &= -\gamma_1 + G_n^{\intercal}(\phi - \Sigma_S\theta_1) \ .
\end{aligned}
\tag{14}
$$

The equivalent equations for nominal rates have the same recursive structure, namely

$$
\begin{aligned}
\widehat{E}_1 &= -\widehat{\gamma}_0 - \iota_M^\top \mu_S + \frac{1}{2}\iota_M^\top \Sigma_S \Sigma_S^{\intercal}\iota_M + \iota_M^\top \Sigma\theta_0 \\
\widehat{G}_1 &= -\left(\widehat{\gamma}_1 + \iota_M^\top \phi\right) + \iota_M^\top \Sigma_S \theta_1 \\
& \dots \\
& \dots \\
\widehat{E}_{n+1} &= -\widehat{\gamma}_0 + \widehat{E}_n + \left(\widehat{G}_n^{\intercal} - \iota_M^\top\right)(\mu_S - \Sigma_S\theta_0) \\
& \quad -\frac{1}{2}\left(\widehat{G}_n^{\intercal} - \iota_M^\top\right)\Sigma_S^{\intercal}\Sigma_S\left(\widehat{G}_n^{\intercal} - \iota_M^\top\right)^{\intercal} \\
\widehat{G}_{n+1} &= -\widehat{\gamma}_1 + \left(\widehat{G}_n^{\intercal} - \iota_M^\top\right)(\phi - \Sigma_S\theta_1) \ ,
\end{aligned}
\tag{15}
$$

where ι_M is a $(M \times 1)$ vector of 1s. Note that the equation of the real pricing kernel becomes

$$
m_{t+1} = \exp\left(-r_t - \frac{1}{2}\overline{\lambda}_t^\top \overline{\lambda}_t - \overline{\lambda}_t^\top \varepsilon_{t+1}\right) \ , \tag{16}
$$

where

$$
\overline{\lambda}_t = [\lambda_0, \theta_0] + \begin{bmatrix} \lambda_1 & 0 \\ 0 & \theta_1 \end{bmatrix} \begin{bmatrix} X_t \\ S_t \end{bmatrix} \ .
$$

Based on the state space representation in (13), 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.

14 Preliminary results for U.S. jobless claims

Results are very similar to those obtained by the original model (10). However the new version of the state-space representation given in (13) gives the opportunity of analyzing the term structure of the announcement effect on the entire nominal and real term structure as well as the impulse response function of the nominal and real interest rates to an unexpected change in the macroeconomic news, i.e. when the release greater or smaller than the median forecast. This Section presents the results for the case of the United States jobless claims, which measures emerging unemployment by tracking how many people have filed for unemployment benefits in the previous week. Investors can use this report to gather pertinent information about the economy, even if it's a very volatile data, so the four week average of jobless claims is also monitored.

Figure 9 plots the estimated factor loadings for nominal and real interest rates (i.e. $\widehat{G}$ and G). The factor loadings of the surprise in the U.S. jobless claims are increasing in maturity for both nominal and real interest rates. A surprise affects more the short end than the long end of the term structure and its impact is stronger on the real term structure than on the nominal one. The impulse response function behaves accordingly (figure 10); a surprise in the U.S. jobless claims impacts the entire term structure but its effects halves in 7 weeks and completely disappears after 52 weeks. As expected from the factor loadings the impacts is stronger for real rates.

15 Further extension

[TO BE WRITTEN AND EXTRACTED FROM A WORK IN PROGRESS]

- The same analysis has to be made for the macroeconomic news in table 3 for the euro area and for the United States;
- Comparison between the factor loadings of news of the same group (inflation, future economic activity, growth and labor market) in the euro area and in the United States;
- Comparison between pre-crisis (from January 1998 to July 2007 in the United States and from January 2002 to July 2007 in the euro area) and post-crisis (from August 2007 to December 2010) periods.

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Part III
Figures and tables

16 Figures

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

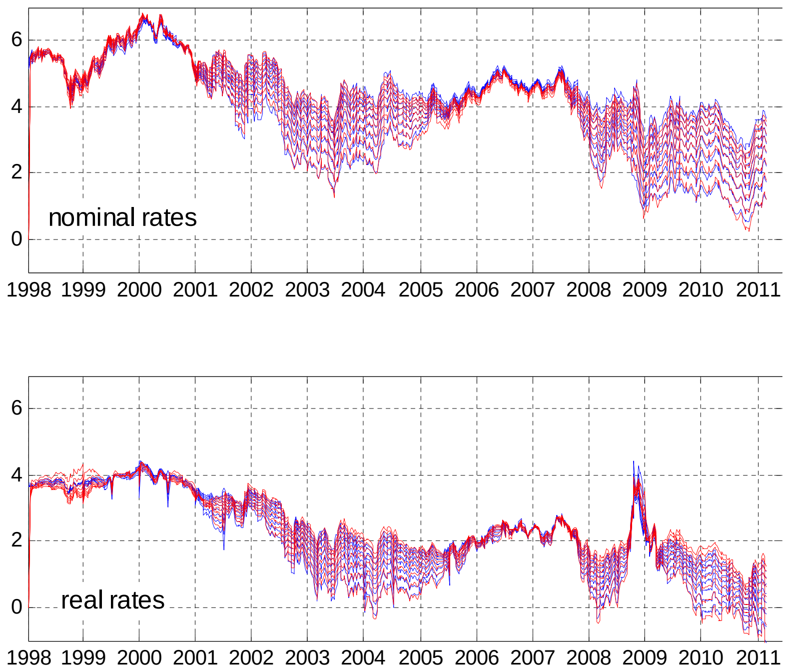


Figure 2 – Euro area: current (blue) and estimated (red) nominal and real rates

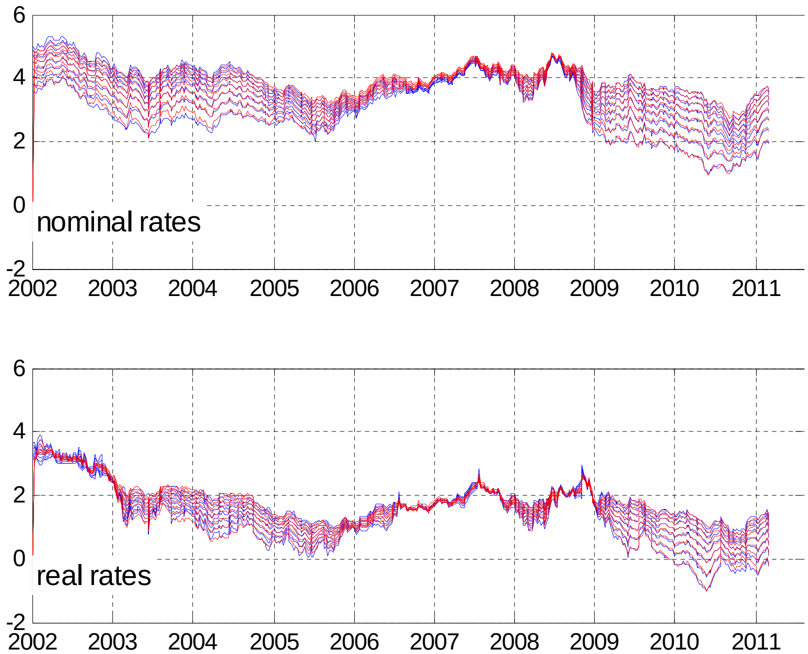


Figure 3 – euro area: latent factors and observable variables

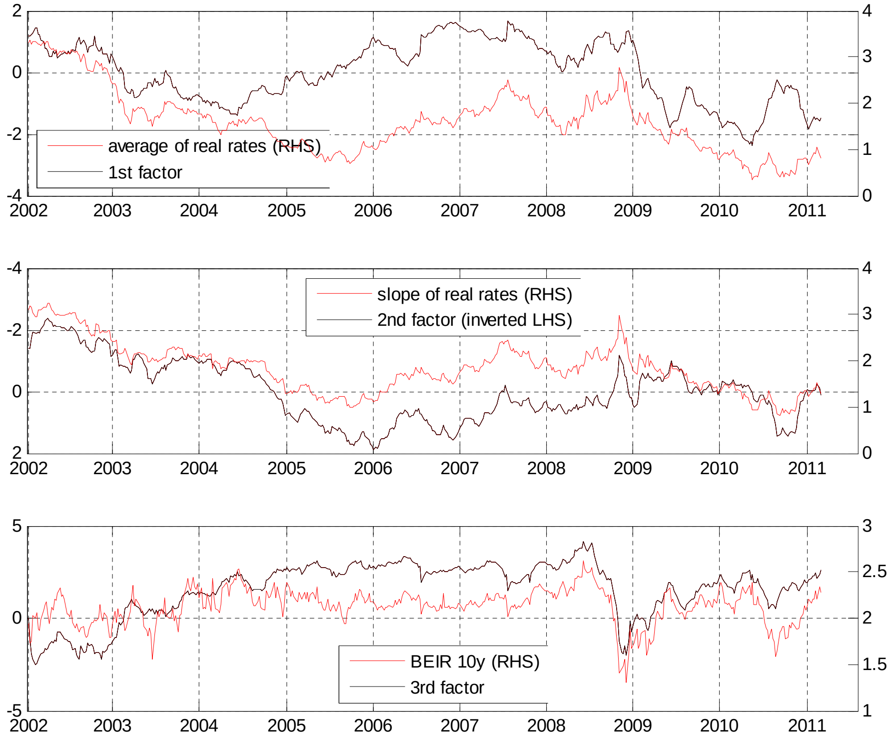


Figure 4 – United States: latent factors and observable variables

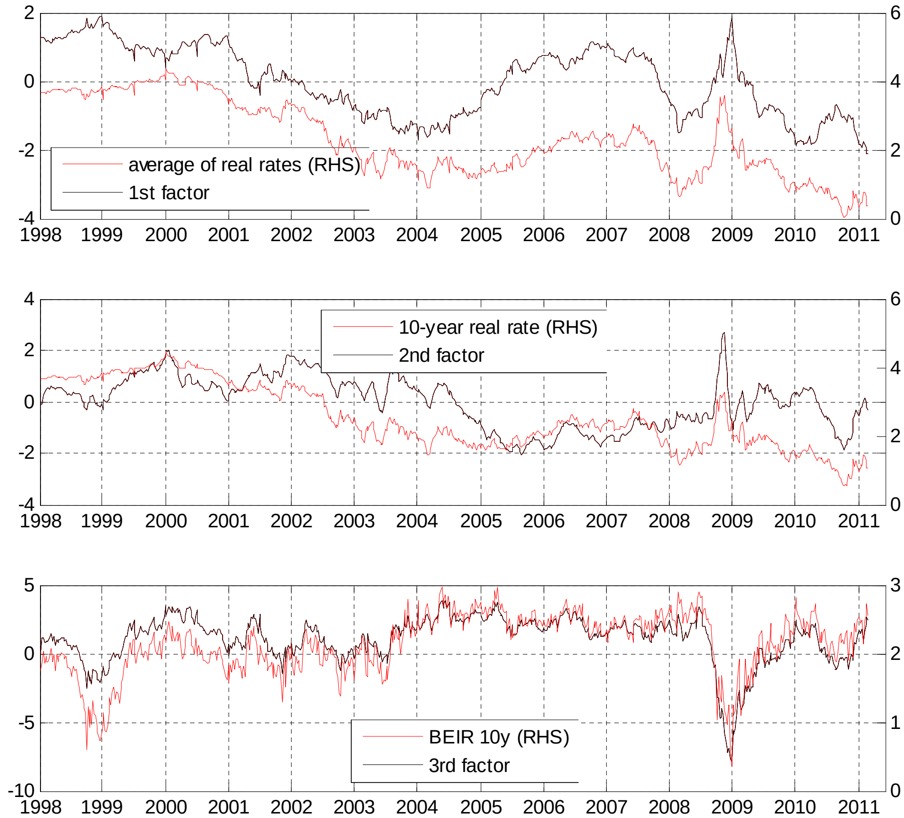
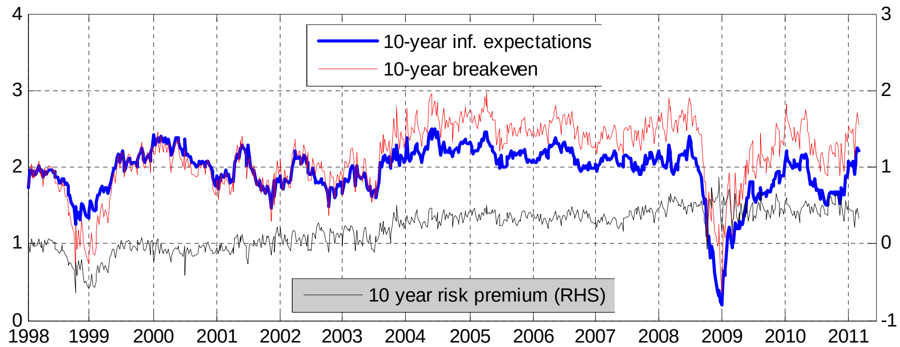


Figure 5 – United States: BEIR, expected inflation and risk premium



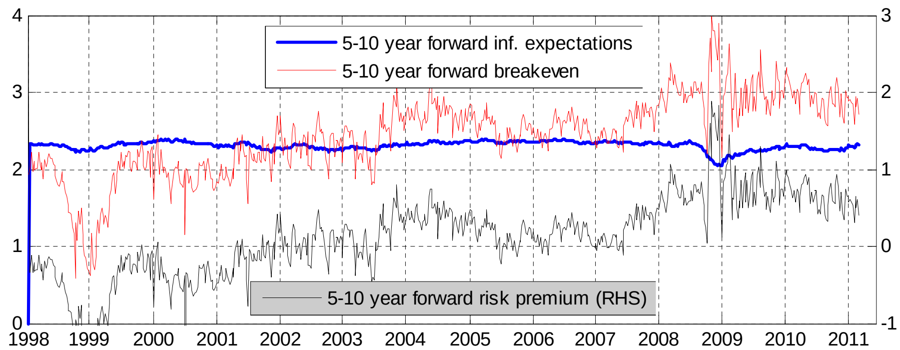


Figure 6 – Euro area: BEIR, expected inflation and risk premium

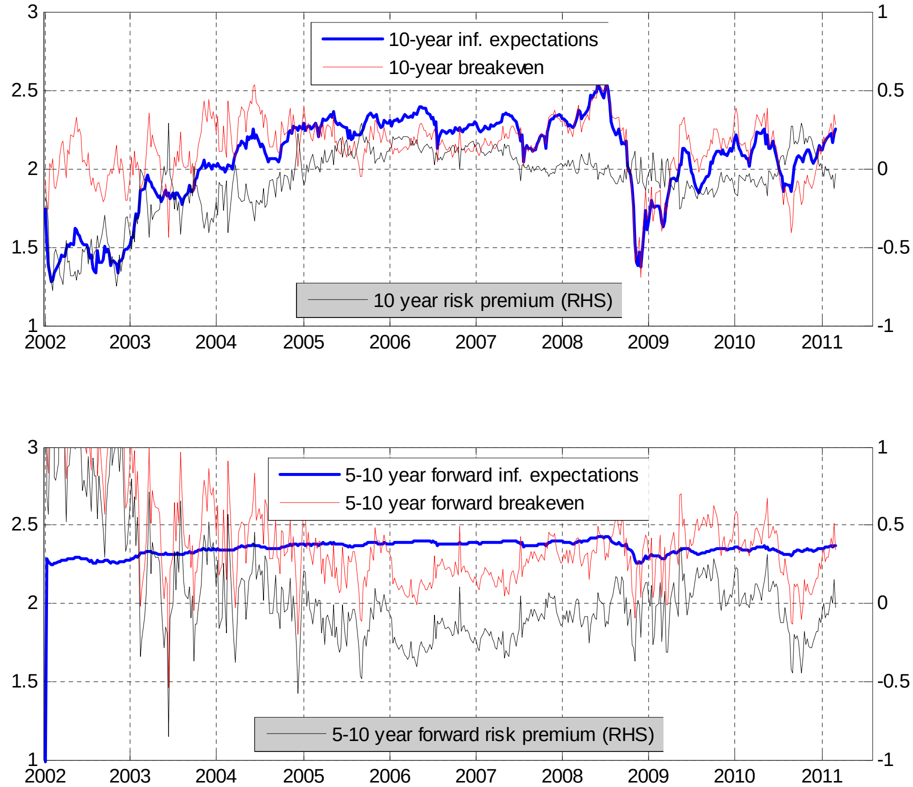


Figure 7 – euro area: impulse response function

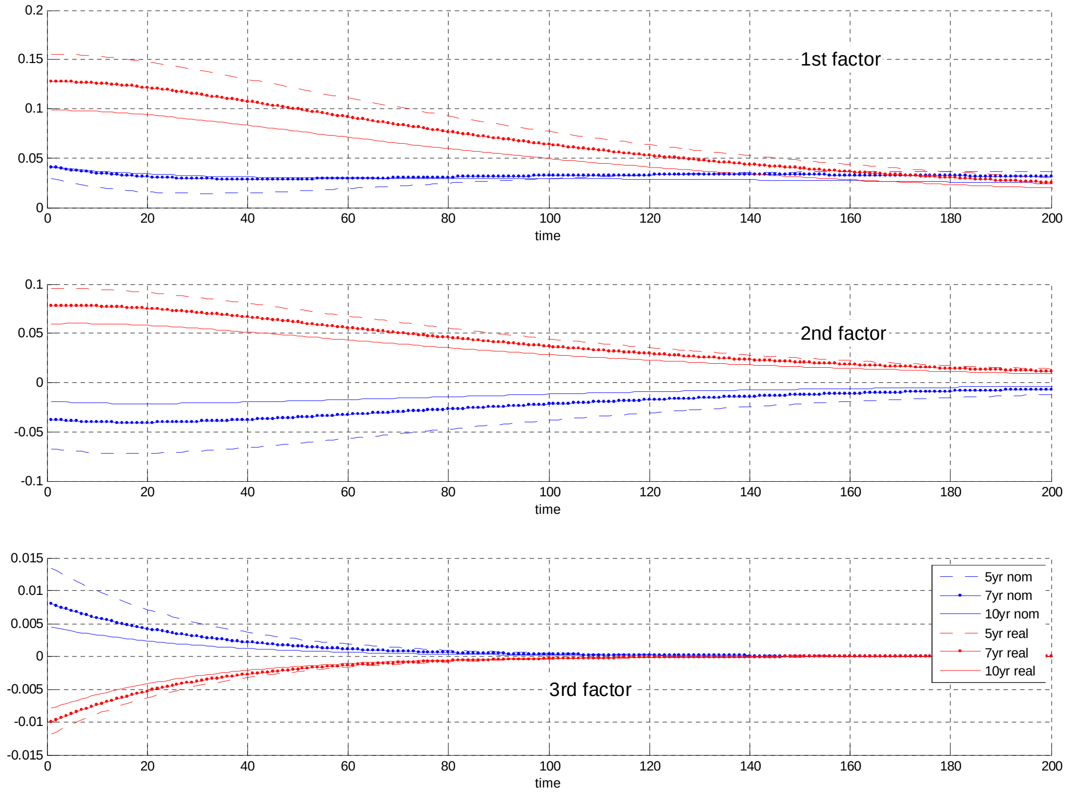


Figure 8 – United States: impulse response function

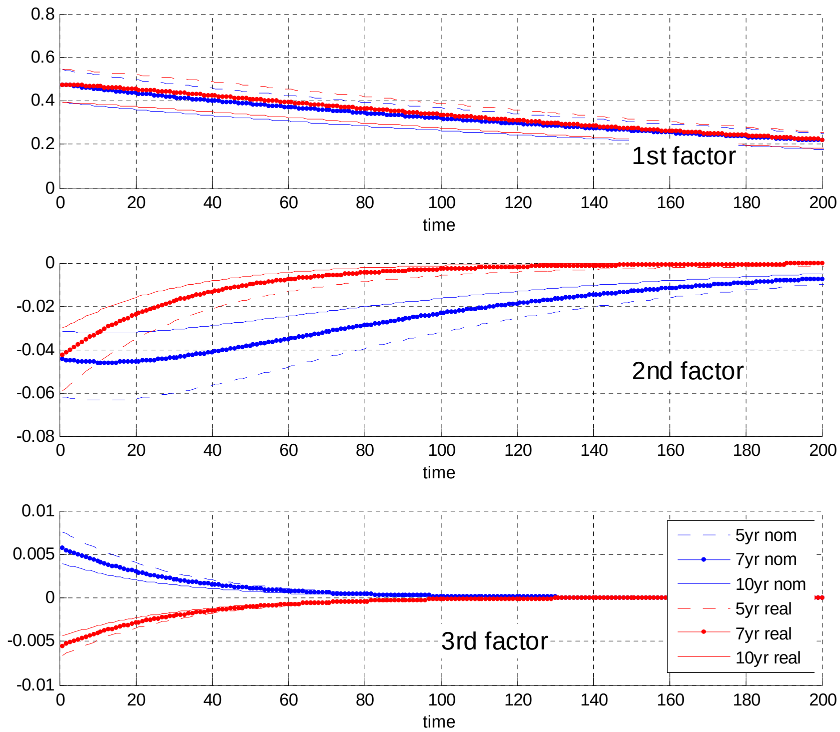


Figure 9 – factor loadings for the US jobless claim surprise ($\widehat{G}$ and G)

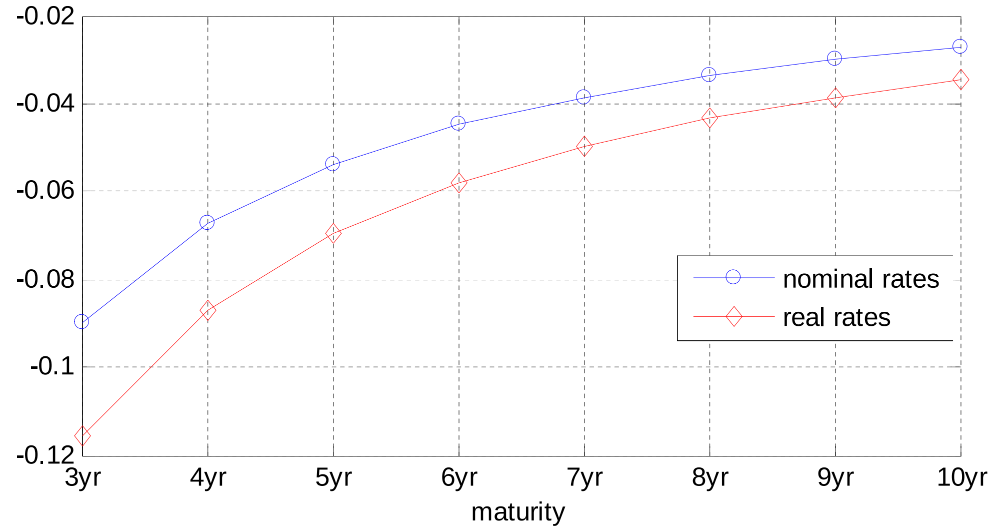
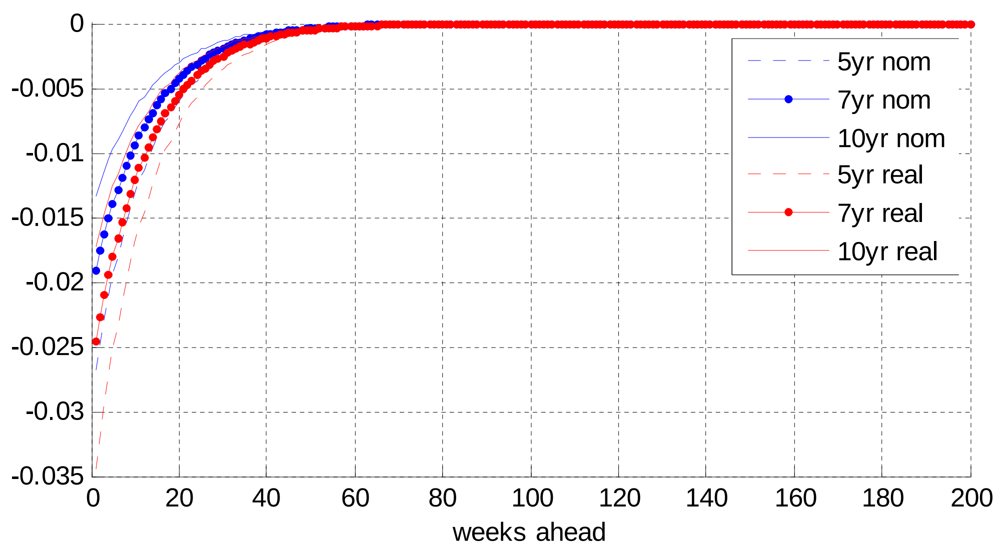


Figure 10 – response function of interest rates to a shock in the US jobless claim



17 Tables

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 February 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 February 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 - Economic variable news

Bloomberg code	description	freq.	mean	std	min	max	field
United States							
NAPMPMI	ISM Manufacturing PMI SA	M	5.96	1.99	-3.00	3.70	future ec. act.
CONCCONF	Conference Board Cons. Conf.	M	-0.85	4.94	-2.82	2.49	future ec. act.
NFP TCH	Employees on Nonfarm Payrolls	M	-24.73	89.61	-3.54	2.09	labor market
GDP CQOQ	GDP Chained 2005 Dollars	Q	-3.45	0.28	-4.17	2.08	growth
PPI CHNG	PPI By Proc. Stage Finish. Goods	M	6.64	0.48	-2.46	3.49	inflation
CPI CHNG	CPI Urban Consumers	M	-3.26	0.13	-2.99	2.99	inflation
IP CHNG	Industrial Production	M	-10.59	0.36	-5.49	3.02	growth
USTBTOT	Trade Balance Bal. Of Payments	M	-1.48	2.96	-2.96	3.57	growth
INJCJC	Initial Jobless Claims	W	26.82	21.24	-3.90	3.76	labor market
USURTOT	Unemployment Rate	M	-13.32	0.14	-3.35	2.68	labor market
RSTAMOM	Adj. Retail - Food Serv. Sales	M	2.10	0.70	-2.25	6.49	growth
DGNOCHNG	Durable Goods New Orders Ind.	M	-2.66	2.70	-3.03	3.99	growth
euro area							
ECPMICOU	EC Composite PMI Output	M	9.89	0.37	-2.66	3.19	future ec. act.
EUGNEMUQ	Eurostat GDP constant prices	Q	-20.25	0.05	-5.02	1.67	growth
EUITEMUM	Eurostat Ind. Prod. ex constr.	M	-8.86	0.53	-3.35	1.86	growth
EUNOEZM	Eurostat New Orders	M	6.75	2.04	-2.78	3.32	growth
PMITMEZ	Manufact PMI Markit Survey	M	5.08	0.30	-2.32	3.65	future ec. act.
ECPMICOU	EC Composite PMI Output	M	9.89	0.37	-2.66	3.19	future ec. act.
EUBCI	EC Business Climate Indicator	M	9.61	0.19	-2.80	2.50	future ec. act.
EUCCEMU	EC Economic Sentiment Indicator	M	-8.62	1.28	-3.12	3.90	future ec. act.
EUESEMU	EC Economic Sentiment Indicator	M	4.02	1.50	-3.72	2.46	future ec. act.
PMITMEZ	Manufact PMI Markit Survey	M	5.08	0.30	-2.32	3.65	future ec. act.
EUPPEMUM	Eurostat PPI Industry Ex constr.	M	-10.26	0.14	-6.02	2.67	inflation
RSSAEMUM	Eurostat Retail Sales Volume	M	-24.90	0.63	-2.69	4.90	growth
ECCPEMUM	Eurostat MUICP All Items	M	2.15	0.07	-4.00	2.66	inflation
GRZEEUEX	ZEW Expectation of Ec. Growth	M	1.39	7.74	-2.60	2.42	future ec. act.
PMITSEZ	Services PMI Markit Survey	M	3.73	0.56	-3.91	3.02	future ec. act.
PMITMEZ	Manufact PMI Markit Survey	M	5.08	0.30	-2.32	3.65	future ec. act.
ECCPEMUM	Eurostat HICP All Items	M	2.15	0.07	-4.00	2.66	inflation

Source: Bloomberg. For each economic variable the surprise is computed as difference between the actual release (Bloomberg datatype is ACTUAL_RELEASE) and the median of the survey (Bloomberg datatype is BN_MEDIAN_SURVEY) and is standardized by its standard deviation. EC stands for European Commission. Q/M/W indicates that data are released at a quarterly, monthly and weekly frequency. Mean is the arithmetic average of the standardized surprise as in equation (12), std. is the standard deviation of the surprise, e.g. the denominator in equation (12), min and max are the minimum and the maximum of the standardized surprise given by equation (12).