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UNDERSTANDING POLICY RATES AT THE ZERO LOWER BOUND: INSIGHTS FROM A BAYESIAN SHADOW RATE MODEL

by Marcello Pericoli* and Marco Taboga *

Abstract

Term structure models are routinely used by central banks to assess the impact of their communication on market participants' expectations for interest rates. But some recent studies have shown that the traditional term structure models may be misleading when policy rates are at the zero lower bound, one reason being that such models cannot reproduce the stylized fact that once policy rates reach the ZLB they tend to remain there for a prolonged period. A consensus has now emerged that shadow rate models, pioneered by Black (1995), can solve this problem. The thesis is that the “shadow rate” (the short-term interest rate that would prevail in the absence of the ZLB) can stay in negative territory for long time spans even when the actual rate remains close to the ZLB. Since they are strongly non-linear, shadow rate models are especially hard to estimate, and to date the literature has used only approximate methods to this end. Instead, we propose an exact Bayesian method of estimation and apply it to developments in the euro and dollar yield curves since the end of the 1990s. Our estimates confirm and provide a quantitative measure of quantify the significant divergence of monetary policies between the euro area and the US: between 2009 and 2013, the shadow rate was much lower in the US than in the euro area, and since then the opposite has been the case. At the end of our sample period in January 2015, according to our model the most likely date for the the first increase in policy rates was estimated to be mid-2015 in the US and 2020 in the euro area.

JEL Classification: C32, E43, G12.
Keywords: zero lower bound, shadow rate term structure model.

Contents

* Bank of Italy, DG Economics, Statistics and Research

1 Introduction[1]

Central banks, as well as other public and private entities, routinely use term structure models to gauge market participants' views about future interest rate developments. For central banks, one of the main motivations is the need to assess the effectiveness of their communication in terms of its impact on market participants' expectations. This is important, for example, when central banks explicitly communicate to the public their expectations regarding the path of future policy rates. Often, such "forward guidance" is phrased in terms of the time during which the central bank intends to keep policy rates near certain levels. In other cases, central banks announce that future policy rates will be linked to certain macroeconomic outcomes. In any case, the term structure of interest rates contains potentially useful information about market participants' expectations, which is usually extracted with the help of term structure models.

Recent studies (e.g., Kim and Singleton 2012, Krippner 2013a) have pointed out that the indications provided by several popular term structure models can be misleading when policy rates are – as they have been for many currencies in recent years – near the zero lower bound (ZLB). As a matter of fact, many models are unable to reproduce the stylized fact that policy rates tend to remain at the ZLB for prolonged periods of time once they reach it. Furthermore, when policy rates are at the ZLB, some frequently used models (namely, those belonging to the Gaussian affine class) tend to assign high probability to future scenarios where policy rates decrease further and become significantly negative, which contradicts the very existence of the ZLB. Because of these unrealistic features, several of the most popular term structure models produce flawed estimates of the expected future policy rates and the term premia embedded in the yield curve. A consensus has recently emerged that the so-called "shadow rate models", first introduced by Black (1995), are particularly apt

[1] The views expressed are those of the authors and do not necessary represent those of the Bank of Italy. Email addresses: `marcello.pericoli@bancaditalia.it` and `marco.taboga@bancaditalia.it`. We thank Giuseppe Grande, Jean-Paul Renne, Stefano Siviero, Gabriele Zinna and the participants in the Workhop on Low Inflation and its Implications for Monetary Policy for helpful comments.

to obviate the aforementioned drawbacks. These models are characterized by the existence of an unobservable shadow rate, which could be thought of as the short-term interest rate that would prevail if storing physical currency was impossible or not economically viable[2] (also other interpretations are possible; see below for a more detailed discussion). In such a framework, it can happen that the shadow rate remains well below zero for prolonged periods of time, and, as long as the shadow rate is negative, the actual short-term rate remains at the ZLB, which generates the persistence observed in the data[3].

Due to their high nonlinearity, shadow-rate models are particularly difficult to estimate and have been so far only estimated with approximate methods. A contribution of this paper is to propose an exact Bayesian method for their estimation (see Section 3 for a review of the existing methods).

We employ shadow rate models to study developments in euro and US dollar yield curves since the end of the '90s. Our estimates of the shadow rate confirm that there has been a "dramatic divergence" (Orphanides 2014) of monetary policies in the euro area and in the US over the past years. According to our estimates, between 2009 and 2013 the shadow rate was much lower in the US than in the euro area, while the opposite has happened since 2014. The divergence is expected to last well into the future: at the end of our sample (January 2015), the most likely date of the lift-off from the zero lower bound (i.e., the first increase in policy rates) was estimated to be around mid-2015 in the US and around 2020 in the euro area.

As far as risk premia are concerned, we find that the estimates provided by our shadow rate model can be significantly different from those provided by traditional models. For example, the Federal Reserve regularly provides updated estimates of the term premia embedded in the US government bond yield curve, obtained with a Gaussian affine model (Kim

[2]Nominal interest rates cannot go significantly below zero (more precisely, they cannot go below minus the cost of storing physical currency) because it is possible to make arbitrages between currency and short-term loans.

[3]Note that the short-term rate has an option-like behavior: it is equal to the shadow rate when the shadow rate is positive, but it remains at zero when the shadow rate becomes negative.

and Wright 2005). According to these estimates, term premia on long-term bonds have been on a declining trend since the '90s and were significantly negative in the period 2012-2013. According to our model, instead, term premia did not display any significant downward trend and remained positive until the end of 2014, when they turned slightly negative. We also find evidence of a strong positive co-movement between euro and US dollar term premia, in accordance with the evidence provided by other studies (e.g., Jotikasthira et al. 2015).

The interest in modelling the term structure at the ZLB was initially stimulated by the low level of interest rates observed on the Japanese government bond market, where short-term rates have hovered around 0.5 per cent since the '90s and slightly above 0 per cent since 2011. As a consequence, a number of studies focus on the Japanese market (see, e.g., Gorovoi and Linetsky 2004, Ueno et al. 2006, Kim and Singleton 2012, Christensen and Rudebusch 2014, Monfort et al. 2014). Recently, some studies have analyzed also the US market, where the Federal Funds rate target has been between 0 and 0.25 per cent since the end of 2008 (see, e.g., Hamilton and Wu 2011, Andreasen and Meldrum 2013, Bauer and Rudebusch 2013, Christensen and Rudebusch 2013, Krippner 2013a, Krippner 2013b, Priebsch 2013, Christensen et al. 2014, Kim and Priebsch 2014, Wu and Xia 2014). Finally, the extremely low levels of short-term rates observed in the euro area since the second half of 2012 has fostered studies also on this market (see, e.g., Renne 2012 and Lemke and Vladu 2014).

There are two main reasons why new term structure models have been proposed to study interest rates near the ZLB.

The first reason is that several popular models, namely those belonging to the class of Gaussian affine term structure models (GATSM; e.g., Hamilton and Wu 2012, Joslin et al. 2013) do not constrain interest rates to be positive. This fact has minor practical consequences when observed rates are far from the ZLB. However, when rates are close to zero, these models tend to assign high probability to future scenarios in which short-term rates go much below zero; since such scenarios cannot materialize in practice, due to the possible arbitrage between loans and physical cash, attaching a high probability to them determines a downward bias in the estimates of expected future rates and an upward bias in those of term premia. One way to enforce positivity is to drop the Gaussianity assumption and work with strictly positive processes such as CIR processes or AR gamma zero processes (Monfort et al. 2014). Another possibility is to abandon the affine specification of the short term rate and use a quadratic specification with appropriate restrictions; this approach leads to the well-known class of quadratic term structure models (QTSM) studied in Ahn et al. (2002), Leippold and Wu (2002) and Realdon (2006) among others. Another possibility is to use shadow rate models: as explained above, even if they feature a shadow rate that can become significantly negative, they constrain short-term rates to be positive.

The second reason why new models have been proposed is that traditional models are often unable to reproduce one of the main stylized facts that are observed when interest rates are near the ZLB, namely, that once the short term rate reaches the bound it tends to remain there for long periods of time. As suggested by the specification analysis carried out by Kim and Singleton (2012), not only shadow rate models can provide a satisfactory solution to this problem, but they can also outperform the other models in fitting realized excess returns and, implicitly, risk premia.

According to our empirical results, shadow rate models indeed provide a description of the dynamics of interest rates that has several realistic features. For example, as explained above, the differences between the dynamics of the estimated shadow rate in the euro area and in the US seem to reflect the different monetary policies adopted in the two blocks (e.g., a much earlier quantitative easing in the US, and the prospects of the Federal Reserve lifting off rates from the ZLB several quarters – if not years – before the ECB). Moreover, the estimated (most likely) lift-off date for the United States at the end of the sample coincides with the date that was frequently reported by the press as being the most likely at that time. All in all, estimates of the shadow rate seem a powerful tool to rigorously summarize, in a very parsimonious manner, the wealth of information that comes from the term structure of interest rates.

We do, however, find evidence of some shortcomings of shadow rate models. First, we find that, unless some constraints are imposed – with a modicum of subjectivity – on the parameters, the estimated models tend to fit the yield curves very well, but can give rise to unreasonable estimates of other quantities of interest: for example, unconstrained models tend to produce implied volatilities of the shadow rate that are an order of magnitude higher than the historical volatilities of the short-term rate, as well as estimated trajectories of the shadow rate that are characterized by implausibly large negative values (the latter drawback has been found in the majority of studies on shadow rate models). While the constraints we need to impose in order to rule out these anomalous behaviors seem quite mild, nonetheless they introduce some degree of subjectivity in the estimation process. Second, we find that the economic interpretation of the estimated level of the shadow rate, when it is negative, can be challenging. For example, even when some constraints are introduced, the median of the posterior distribution of the shadow rate in the euro area is about -6 per cent at the end of our sample[4]. It seems doubtful that such a large negative number can be interpreted as the interest rate that would be optimally set by the central bank in the absence of the ZLB. Instead, it seems more reasonable to interpret a negative value of the shadow rate as a measure of distance from the lift-off; in other words, a large negative number signals that the most likely date for the lift-off is far away in the future. In this sense, the shadow rate is more of a statistical device to parsimoniously describe the stance of monetary policy at the ZLB than a proper economic measure of the interest rate that would prevail in the absence of the ZLB. This interpretation is in line with the interpretation of Krippner (2013a), who also interprets the shadow rate as a measure of the monetary stance, after observing that its movements tend to be broadly consistent with the timing of unconventional monetary policy events.

The rest of the paper is organized as follows: Section 2 describes the model; Section 3

[4]In the absence of constraints, values as low as -50 per cent are obtained.

presents the estimation methodology; Section 4 discusses the data; Section 5 presents the empirical results; Section 6 concludes.

2 The model

In this section we describe the main building blocks of our shadow rate term structure model (SRTSM).

SRTSM were first proposed by Black (1995) and have been recently studied by, inter alia, Kim and Singleton (2012), Bauer and Rudebusch (2013), Krippner (2013a), Christensen and Rudebusch (2014), Ichiue and Ueno (2013), and Andreasen and Meldrum (2013).

In a SRTSM the short term rate r_t, that is, the interest rate applied to risk-free one-period loans, is equal to zero or to the shadow rate s_t, whichever is larger.

$$r_t = \max(0, s_t) \tag{1}$$

One possible interpretation is as follows: the shadow rate is the interest rate that would prevail in the absence of physical currency; however, in the presence of currency, nominal interest rates cannot go below zero because it is possible to make arbitrages between currency and short-term loans; therefore, the short-term rate has an option-like behavior: it is equal to the shadow rate when the shadow rate is positive, but it remains at zero when the shadow rate becomes negative.

Note that even if the short-term rate goes below zero (as, for example, short-term OIS rates in the euro area in 2014-2015), it cannot go below minus the cost of the physical storage of currency, which represents the cost of making an arbitrage between currency and risk-free loans. We assume in what follows that the cost of storage is negligible, and we treat any excursion of the observed value of r_t below zero as a pricing error.

The shadow rate is specified as an affine function of a $K \times 1$ vector X_t of latent factors

$$s_t = a + bX_t \tag{2}$$

where a is a scalar and b is a $1 \times K$ vector of factor loadings.

Note that in a standard affine term structure model (ATSM) equation (2) also holds, but $r_t = s_t$. The latter is the only difference between ATSM and SRTSM. In particular, there are no differences in the specification of the factor dynamics.

In what follows, we assume Gaussian dynamics. The pricing factors X_t follow a first order vector autoregression under the real-world measure P:

$$X_t = \mu + \rho X_{t-1} + \Sigma\varepsilon_t \tag{3}$$

where μ is a $K \times 1$ drift vector, ρ is a $K \times K$ autoregression matrix, Σ is a $K \times K$ volatility matrix and ε_t is a sequence of IID standard multivariate normal $K \times 1$ random vectors.

Under the usual assumptions on the functional form of the pricing kernel, the dynamics of X_t under the risk-neutral pricing measure Q are

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

where the drift vector $\overline{\mu}$ and the autoregression matrix $\overline{\rho}$ under Q are generally different from those under P, but the volatility matrix Σ remains unchanged and the sequence of errors η_t is also IID standard multivariate normal.

We impose the following restrictions on the parameters of the model:

1. $\Sigma = I$;

2. $\overline{\mu} = 0$;

3. $\overline{\rho}_{ii} \geq \overline{\rho}_{jj}$ if $i > j$;

4. $b \geq 0$;

5. $\overline{\rho}$ diagonal;

While restrictions 1-4 are necessary for identification, restriction 5 is over-identifying and could be replaced by the weaker requirement that $\overline{\rho}$ be an upper diagonal matrix satisfying the real Schur property (see Pericoli and Taboga 2012).

We assume discrete compounding, so that the price of a zero-coupon bond expiring in n periods is

$$P_t^n = \mathrm{E}^{\mathrm{Q}}\left[\left.\frac{1}{\prod_{j=0}^{n-1}(1+r_{t+j})}\right| X_t\right] \tag{5}$$

and its yield is

$$y_t^n = \left(\frac{1}{P_t^n}\right)^{1/n} - 1 \tag{6}$$

Let $n_1, \ldots, n_M$ be M maturities of interest. Assume they are in increasing order, so that n_M is the largest bond maturity.

As in Bauer and Rudebusch (2013), we discretize the support of the sequence $\{\eta_t\}$ in order to compute prices and yields (eq. 5 and 6). Because P_t^n depends only on the joint distribution of $\{\eta_t\}$, which is IID, we can focus on

$$\widetilde{\eta}_k = \eta_{t+k} \tag{7}$$

for $k = 1, \ldots, n_M$, independently of t. We extract at random 500 realizations of the sequence $\widetilde{\eta}_k$, we compute their antithetic sequences (by changing signs), and we use them to form an antithetic sample of $1,000$ realizations. Realizations of the sequence are denoted by $\{\widetilde{\eta}_k^s\}$ for $s = 1, \ldots, S$, where $S = 1,000$. Once extracted, $\widetilde{\eta}_k^s$ remains constant across maturities, time periods, and iterations of the estimation algorithm.

We then compute prices as

$$P_t^n = \frac{1}{S}\sum_{s=1}^{S} \frac{1}{\prod_{j=0}^{n-1}\left(1+\max\left(0, a + b\left(\overline{\rho}^j X_t + \sum_{\{k\in\mathbb{N}:k\leq j\}}\overline{\rho}^{j-k}\widetilde{\eta}_k^s\right)\right)\right)} \tag{8}$$

As demonstrated by Bauer and Rudebusch (2013) this discretization introduces negligible approximation errors (less than one basis point on average when $S = 1,000$) and can be considered, for all practical purposes, equivalent to the exact computation of P_t^n.

Denote by

$$y_t = \left[\begin{array}{cccc} y_t^{n_1} & y_t^{n_2} & \dots & y_t^{n_M} \end{array} \right]^\top \tag{9}$$

the vector of yields.

We assume that the yields are observed with error, so that the observed yields are

$$y_t^o = y_t + v_t \tag{10}$$

where v_t is a sequence of IID multivariate normal random vectors having zero mean and covariance matrix V. The pricing errors are assumed to be cross-sectionally independent, so that V is diagonal.

3 The estimation method

Different methods have been proposed to estimate shadow rate models, such as: i) lattices (Ichiue and Ueno 2007); ii) finite-difference methods (Kim and Singleton 2012); iii) parameter estimation with a companion affine model, and approximate extended Kalman filtering (Bauer and Rudebusch 2013); iv) option pricing approximation (Krippner 2013a and Christensen and Rudebusch 2014); v) ignoring Jensen inequality corrections so as to exploit analytical results on truncated normal distributions (Ichiue and Ueno 2013); vi) perturbations and Taylor series expansions around a deterministic steady state (Andreasen and Meldrum 2013); vi) approximations of forward rates (Wu and Xia 2014). While all of these methods rely on one or more approximations (whose impact on results is often difficult to assess), we use a Bayesian method that relies only on the discretization in eq. (8). As we already explained, the discretization consists in replacing the continuous support of the joint distribution of $\{\eta_t\}$ with a discrete one. First of all, this replacement has been proved to be completely innocuous (as it introduces numerical errors that are, to all practical purposes, negligible). Second, the discrete state space can be considered as the true state space of the model (remember that it is fixed across maturities, time periods, and iterations), so that, in fact, no approximation is introduced.

Let T denote the last observation in the sample.

Denote the matrix of observed yields by

$$Y^o = \left[\begin{array}{cccc} y_1^o & y_2^o & \dots & y_T^o \end{array} \right] \tag{11}$$

the matrix of theoretical yields by

$$Y = \left[\begin{array}{cccc} y_1 & y_2 & \dots & y_T \end{array} \right]$$

and the matrix of pricing factors by

$$X = \left[\begin{array}{cccc} x_1 & x_2 & \dots & x_T \end{array} \right] \tag{12}$$

We estimate quantities of interest by simulating from the posterior distribution of the parameters and of the latent states:

$$f\left(a, b, \mu, \rho, \overline{\rho}, V, X \left| Y^o \right.\right) \propto f\left(Y^o \left| a, b, \mu, \rho, \overline{\rho}, V, X \right.\right) f\left(a, b, \mu, \rho, \overline{\rho}, V, X\right) \tag{13}$$

Prices P_t^n and yields y_t^n depend only on the risk neutral dynamics (eq. 4) and on the functional relation between the factors and the short-term rate (eq. 1 and 2). As a consequence, the only quantities that are required to compute P_t^n and y_t^n are: 1) the diagonal elements of $\overline{\rho}$; 2) the coefficients a and b; 3) the factors X_t. Therefore,

$$f\left(Y^o \left| a, b, \mu, \rho, \overline{\rho}, V, X \right.\right) = f\left(Y^o \left| a, b, \overline{\rho}, V, X \right.\right) \tag{14}$$

Furthermore, we have that

$$
\begin{aligned}
f(a,b,\mu,\rho,\overline{\rho},V,X) &= f(X\,|a,b,\mu,\rho,\overline{\rho},V)\, f(a,b,\mu,\rho,\overline{\rho},V) \\
&= f(X\,|\mu,\rho)\, f(a,b,\mu,\rho,\overline{\rho},V)
\end{aligned}
\tag{15}
$$

where $f(X\,|\mu,\rho)$ is determined by the physical dynamics of the factors (eq. 3).

We assume that the prior on the parameters is uniform improper on the set of admissible values

$$
f(a,b,\mu,\rho,\overline{\rho},V) \propto 1_A(a,b,\mu,\rho,\overline{\rho},V) \tag{16}
$$

where 1_A is the indicator of the set A of admissible values for the parameters (to be specified below).

By putting together equations (13), (14), (15) and (16), we obtain the posterior density

$$
f(a,b,\mu,\rho,\overline{\rho},V,X\,|Y^o) \propto f(Y^o\,|a,b,\overline{\rho},V,X)\, f(X\,|\mu,\rho)\, 1_A(a,b,\mu,\rho,\overline{\rho},V) \tag{17}
$$

The three terms in the above product are straightforward to compute. The first two terms are Gaussian likelihoods. The first term, that is,

$$
f(Y^o\,|a,b,\overline{\rho},V,X) \tag{18}
$$

is the likelihood of the pricing errors v_t, which is straightforward to compute given that the theoretical yields can be computed analytically from equation (8). The second term, that is,

$$
f(X\,|\mu,\rho) \tag{19}
$$

is the likelihood of the VAR errors ε_t. Finally the third term is just an indicator that takes value 1 on the parameter space and 0 elsewhere.

We use a random walk Metropolis-Hastings algorithm with block structure (e.g., Bagasheva et al. 2008) to generate draws from the posterior distributions of the parameters. Each parameter (including the individual entries of X) is treated as a separate block and acceptance probabilities are computed as ratios of posterior densities, calculated according to equation (17). More details on the MCMC algorithm can be found in the Appendix.

4 The data

We use data on the term-structure of Overnight Indexed Swap (OIS) rates (in euros and US dollars). OIS are swap contracts where one counterparty receives a variable payment indexed to the interest rate on overnight unsecured interbank deposits between prime banks, and the other counterparty receives the fixed OIS rate. Because overnight interbank deposits between prime banks are considered virtually risk free, OIS rates are deemed a very good proxy for long-term risk free rates[5] (e.g., Morini 2009, Mercurio 2009, Ejsing et al. 2012, Taboga 2014). Given the timing of the payments of OIS contracts, a simple recursive calculation allows to extract a term structure of zero-coupon spot rates from the term-structure of OIS rates. We use the 3-month and 6-month maturities as well as all the yearly maturities from 1 to 10 years. Data, downloaded from Bloomberg, is available for all maturities since 2005 for euro OIS and since 2012 for US dollar OIS. We backfill the dataset back to 1999 with zero-coupon government bond rates (for euros we use German government bonds). The last observations refer to the end of January 2015, which is used in the estimation and treated as if it was the end of the first quarter of 2015. Figure 1 displays the time series of the OIS rates at selected maturities.

5 Results

For the Bayesian estimation of the model, we make the following choices:

[5] By risk free we mean free of credit and liquidity risk. Of course, long-term OIS rates are subject to interest rate risk.

1. we set the number of factors $K = 3$;
2. we use quarterly data (we have a total of 65 quarters in our sample);
3. we use all the maturities in our dataset (for a total of 12 maturities);

As a consequence, the vector of parameters to be estimated has 226 entries (1 parameter for the scalar a; 3 parameters for the vector b; 3 parameters for the vector μ; 9 parameters for the matrix ρ; 3 parameters for the diagonal matrix $\overline{\rho}$; $M = 12$ parameters for the diagonal matrix V; $K \times T = 3 \times 65 = 195$ parameters for the matrix of pricing factors X).

We perform 250,000 draws for each block (the first 50,000 are used as a burn-in sample and discarded), for a total of 56,500,000 iterations of the Markov chain. Raftery and Lewis' (1995) run length control diagnostic[6] indicates that the sample size is more than 10 times the minimum required size.

The feasible set for the parameters is specified in the Appendix.

At each draw of the parameter vector from the posterior distribution, we compute the entire trajectory of the shadow rate s_t for $t = 1 \ldots, T$, as well as simulated trajectories 30 years ahead for each date: $s_{t+1}^{(t)}, \ldots, s_{t+120}^{(t)}$ where the superscript (t) indicates a simulation started at time t. We use the simulated trajectories to compute risk-neutral yields, risk premia, and the distributions of the lift-off dates at the end of the sample (i.e., for $T = t$). The risk-neutral yields are obtained by using the P-measure instead of the Q-measure in equations (5) and (6). The risk premia are calculated as differences between observed and risk-neutral yields. The lift-off date is defined as the first date on which the shadow rate exceeds 25 basis points, and it can be interpreted as the first date in which the policy rate is raised from the zero lower bound. Thus, we obtain a posterior distribution not only over the parameters, but also over the trajectories of the shadow rates, of the risk premia and of the lift-off dates.

[6] The parameters of the diagnostic are set in such a way that the minimum required size allows to estimate the 2.5% quantile of the posterior distribution of each entry of the parameter vector with an error <1% with probability 95%.

Figure 2 displays the median and the first and last deciles of the posterior distribution of the shadow rate over the entire sample period. According to these estimates, there has been a significant divergence of monetary policies in the euro area and in the US over the past years. In particular, between 2009 and 2013 the shadow rate was much lower in the US than in the euro area (the posterior median is on average -2% in the US vs 0.4% in the euro area), while the opposite has happened since 2014 (-0.9% in the US vs -2.9% in the euro area on average).

Note that the shadow rate decreased to a level significantly below zero in the euro area between the second and the fourth quarter of 2012, but then quickly reverted to zero in the first half of 2013. This behavior might reflect the fact that the European Central Bank (ECB) undertook significant expansionary measures during 2012 (a rate cut, the institution of the Outright Monetary Transactions, a relaxation of collateral rules); however, since the beginning of 2013 the balance sheet of the ECB started to shrink significantly due to the early repayment of a considerable portion of the funds lent to banks through the 3-year Longer Term Refinancing Operations; this reduction might have been perceived by market participants as tantamount to a monetary restriction.

At the end of our sample (January 2015), the posterior median of the USD shadow rate, which had reached a trough of -3.5% in the first quarter of 2013, was slightly positive, at 0.1%. On the contrary, the posterior median of the euro shadow rate, which had hovered around zero until the beginning of 2014, literally collapsed afterwards, reaching 5.9% at the end of January 2015, after the details of the ECB's quantitative easing were announced.

As highlighted both in the Introduction and in the Conclusions, while it seems doubtful that a negative value of the shadow rate can literally be interpreted as the interest rate that would prevail in the absence of the ZLB, it is probably more correct to interpret it as a summary statistic of the stance of monetary policy that provides a measure of distance from the lift-off (i.e., the date of the first increase in policy rates). As a matter of fact, the stark differences between the euro area and the US are reflected in the posterior distribution of lift-off dates at the end of the sample (Figure 3): market participants expected - according to our estimates - that the most likely date for a lift-off from the ZLB would be around mid-2015 in the US[7] and around 2020 in the euro area.

As far as risk premia are concerned (Figures 4 and 5), we find that indeed the estimates provided by our shadow rate model can be significantly different from those provided by traditional models. For example, the Federal Reserve regularly provides updated estimates of the term premia embedded in the US government bond yield curve, obtained with a Gaussian affine model (Kim and Wright 2005). According to these estimates, term premia on long-term bonds have been on a declining trend since the '90s and were significantly negative in the period 2012-2013 (Figure 6). According to our model, instead, term premia did not display any significant downward trend and remained positive until the end of 2014, when they turned slightly negative. We also find evidence of a strong positive co-movement between euro and US dollar term premia (the correlation coefficient between the posterior median of the two premia is 0.57; see also Figure 6), in accordance with the evidence provided by other studies[8] (e.g. Jotikasthira et al. 2015).

6 Conclusions

We have proposed a shadow rate model of the term structure of interest rates and an exact Bayesian method for its estimation. The model has attractive theoretical properties that make it suitable, unlike several popular term structure models, to deal with the challenges that arise when policy rates are near the ZLB. The key element of the model is the explicit characterization of the so-called shadow rate, a fictitious interest rate that measures the

[7] According to the FOMC participants' assessments of appropriate monetary policy (as published in the December 2014 FOMC projection tables, the last available before the end of our sample, in January 2015), 15 out of 17 FOMC members expected that the lift-off would happen in 2015.

[8] The co-movement in long-term rates could be explained by common factors in inflation and inflation uncertainty. However, it is also likely to be a real phenomenon, due to commonalities in real term premia. The latter could be generated by the behavior of preferred habitat investors (pension funds, life insurers, domestic and foreign central banks) whose demand for long-term bonds affects all main international markets. Other factors such as common trends in interest rate risk and demographics can also be at play.

stance of monetary policy and has the following two properties: 1) it coincides with the policy (short-term) rate when the latter is positive; 2) it can become negative when the policy rate is stuck at the ZLB, and the more it is negative the further the first lift-off date is away in the future (the lift-off date is the date on which the central bank will abandon the ZLB and raise rates for the first time). It is by now well understood that, once the ZLB is reached, the monetary stance is altered by acting on long-term interest rates through forward guidance or through direct or indirect market interventions. The shadow rate provides a measure of the monetary stance at the ZLB by summarizing parsimoniously the information about the monetary stance that is embedded in the term structure of interest rates.

We have employed our shadow rate model to study developments in euro and US dollar yield curves since the end of the '90s. Our estimates of the shadow rate provide further evidence, as well as a quantification, of the "dramatic divergence" (Orphanides 2014) of monetary policies in the euro area and in the US over the past years. According to our estimates, between 2009 and 2013 the shadow rate was much lower in the US than in the euro area, while the opposite has happened since 2014. This reflects the different monetary policies adopted in the two blocks (e.g., a much earlier quantitative easing in the US, and the prospects of the Federal Reserve lifting off rates from the ZLB several quarters – if not years – before the ECB).

Besides summarizing the wealth of information that comes from the term structure of interest rates, the model also allows to estimate the risk premia embedded in long-term interest rates. For a number of technical reasons, these estimates should be more reliable than those provided by traditional models. Studies employing traditional models often find that the risk premia have been on a declining trend since the '90s and were significantly negative in recent years. There are reasons to suspect that this result is not genuine, but is due to some biases of traditional models and to their inability to adequately model interest rates near the ZLB. As a matter of fact, according to our model, risk premia did not display any significant downward trend and remained positive until the end of 2014, when they turned slightly negative. We have also found evidence of a strong positive co-movement between euro and US dollar term premia, in accordance with the evidence provided by other studies.

7 Appendix

7.1 The MCMC algorithm

This appendix explains how we generate a sample from the posterior distribution of the parameters by using a random-walk Metropolis Hastings algorithm with block structure.

Stack all parameters $a, b, \mu, \rho, \overline{\rho}, V, X$ in a vector θ, and denote its entries by $\theta_1, \ldots, \theta_N$ (each entry constitutes a block).

The vector of parameters is randomly drawn $J_B + J_K$ times and the first J_B draws are discarded (they are a so-called burn-in sample, also used for tuning the transition density of the chain). The last J_K draws are instead kept and constitute a sample of serially dependent draws (following a Markov chain) from the posterior distribution of θ. The value of θ at the j-th iteration of the Markov Chain is denoted by θ^j and its i-th entry by θ_i^j.

The posterior density of a generic draw θ^j, denoted by $f\left(\theta^j \,|Y^o\right)$, is known up to a constant of proportionality that does not depend on θ^j. Define a $N \times 1$ vector κ of standard deviations of the random-walk increments that will be adaptively adjusted to target an acceptance rate between 7.5 and 15 per cent; the starting value for κ is

$$\kappa_i = 0.0001 \quad i = 1, \ldots, N \tag{20}$$

The chain starts from an admissible value θ^0. The j-th iteration is made up of the following steps:

1. set $l = j - N\,[j/N]$ where $[j/N]$ denotes the integer part of j/N;

2. draw a random number z_j from a standard normal distribution;

3. build a new $N \times 1$ vector $\overline{\theta}$ such that $\overline{\theta}_i = \theta_i^{j-1}$ for $i \neq l$ and $\overline{\theta}_i = \theta_i^{j-1} + \kappa_i z_j$ for $i = l$;

4. compute the acceptance probability a_j as follows:

$$a_j = \min\left(1, \frac{f\left(\overline{\theta}\,|Y^o\right)}{f\left(\theta^j\,|Y^o\right)}\right) \tag{21}$$

5. draw a random number u_j from the uniform distribution on $[0, 1]$;

6. if $u_j \leq a_j$ then set $\theta^j = \overline{\theta}$; otherwise, set $\theta^j = \theta^{j-1}$;

7. if $j \leq J_B$, adjust k_l [9];

8. if $j = J_B + J_K$ end the algorithm, otherwise go back to step 1.

7.2 The admissible set

This section describes the specification of the admissible set.

We impose some restrictions beyond those that are necessary for identification. This is motivated by the observation that, unless some over-identifying constraints are imposed on the parameters, shadow rate models tend to produce implied volatilities of the shadow rate that are an order of magnitude higher than the historical volatilities of the short-term rate, as well as estimated trajectories of the shadow rate that are characterized by implausibly large negative values (the latter drawback has been found in the majority of studies on shadow rate models).

A vector of parameters belongs to the admissible set A (i.e., $1_A(a, b, \mu, \rho, \overline{\rho}, V) = 1$) if and only if:

- all the identification restrictions are satisfied;

[9] The adjustment is done as follows: at each iteration, if an exponentially weighted moving average (with forgetting factor equal to 0.99) of past acceptance indicators (1 in case of acceptance and 0 otherwise) is below 7.5 per cent for the parameter θ_i, we decrease κ_i by a factor of 0.99; if the same moving average is above 15 per cent, we increase κ_i by a factor of 1.01. This choice of parameters for adjusting κ, although admittedly arbitrary, maintained the acceptance rate broadly on target over a number of repetitions of the algorithm.

- all the eigenvalues of ρ are between 0 and 0.99 (this corresponds to an upper bound of approximately 17 years on the half-life of the shocks);

- all the diagonal elements of $\overline{\rho}$ are between 0 and 0.99 (as a consequence, we put an upper bound on the half-life of the shocks also under the pricing measure);

- D is such that all the standard deviations of the pricing errors are less than 20 basis points (in other words, we assume that on average all maturities are priced at a reasonable level of accuracy);

- the conditional standard deviation of the one-period-ahead forecast errors of the shadow rate, that is,

$$\sqrt{\text{Var}\left[s_{t+1}\left|X_t\right.\right]} = \|a\|$$

is between 25 and 100 basis points[10] (to broadly match the sample standard deviation of the first differences of the short term rate, which is equal to 0.42 for the euro and to 0.50 for the US dollar);

- the unconditional mean of the shadow rate under Q, which is equal to a, is between 1 and 6 per cent;

- the unconditional mean of the shadow rate under P, which is equal to

$$a + b\left(I - \rho\right)^{-1}\mu$$

is between 1 and 4 per cent;

- the unconditional mean of the shadow rate under Q is greater than the unconditional mean of the shadow rate under P (which implies that on average bond risk premia are positive), but the positive difference between the two means cannot exceed 200 basis

[10]Note that quarterly compounded rates need to be multiplied by 400 in order to be expressed in percentage points on an annual basis.

points (based on the empirical evidence that the expectations hypothesis is, at least unconditionally, difficult to reject and, therefore, large deviations from the expectations hypothesis are, unconditionally, deemed implausible).

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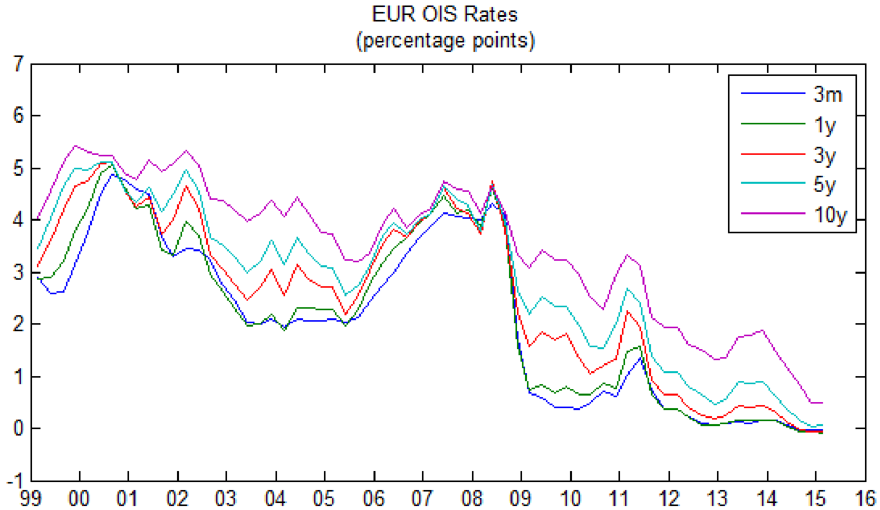
Figure 1 - OIS rates



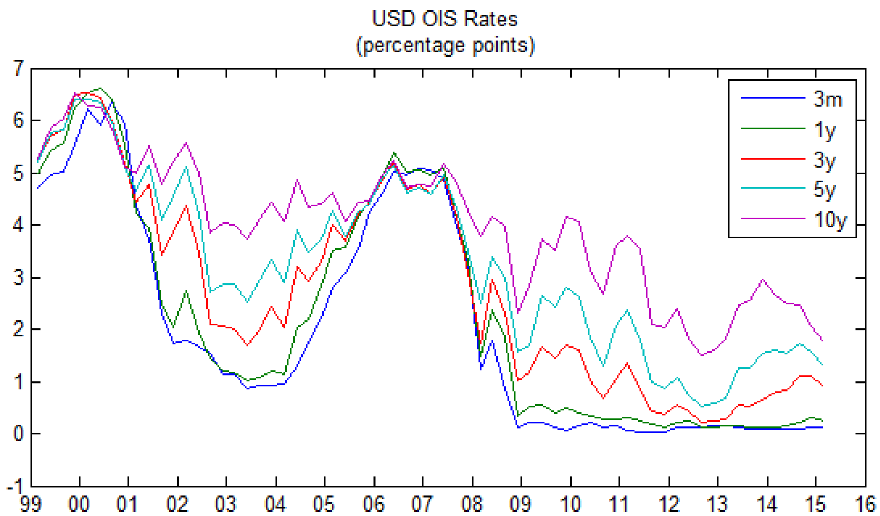


Figure 2 - Shadow rates

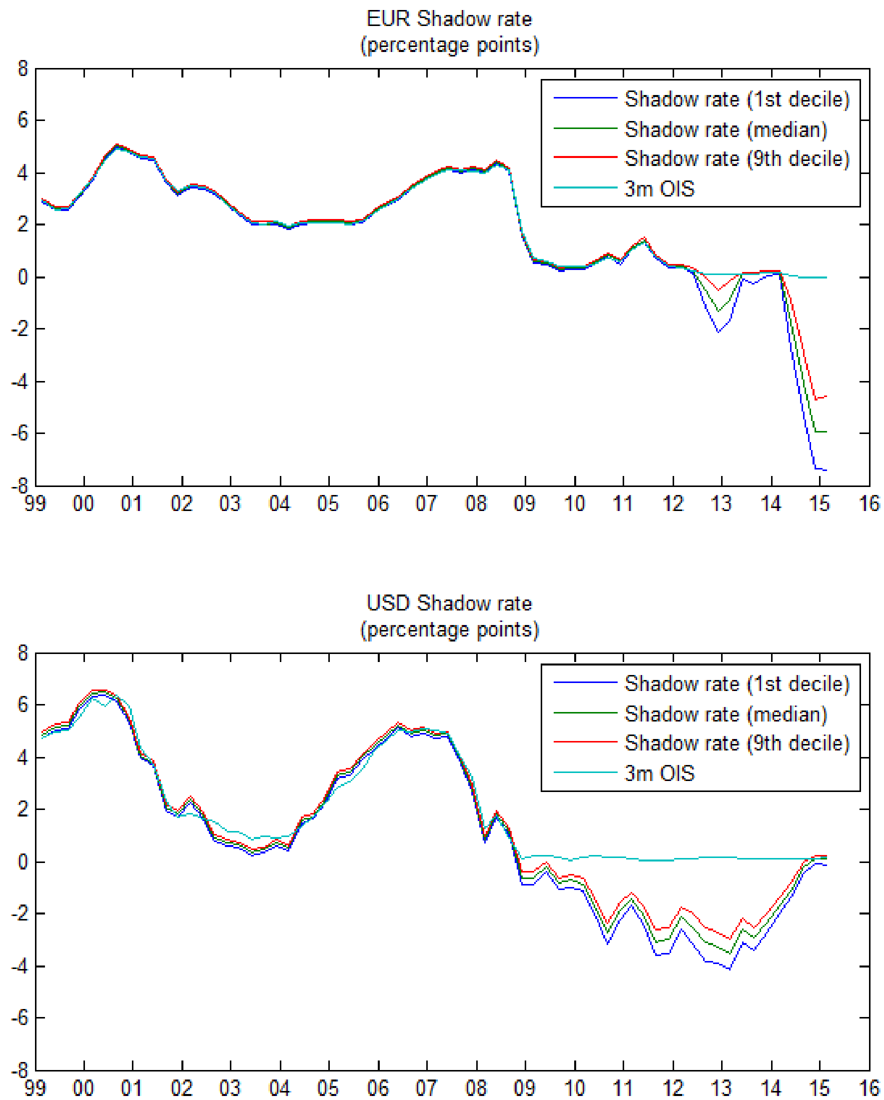


Figure 3 - Distribution of lift-off dates

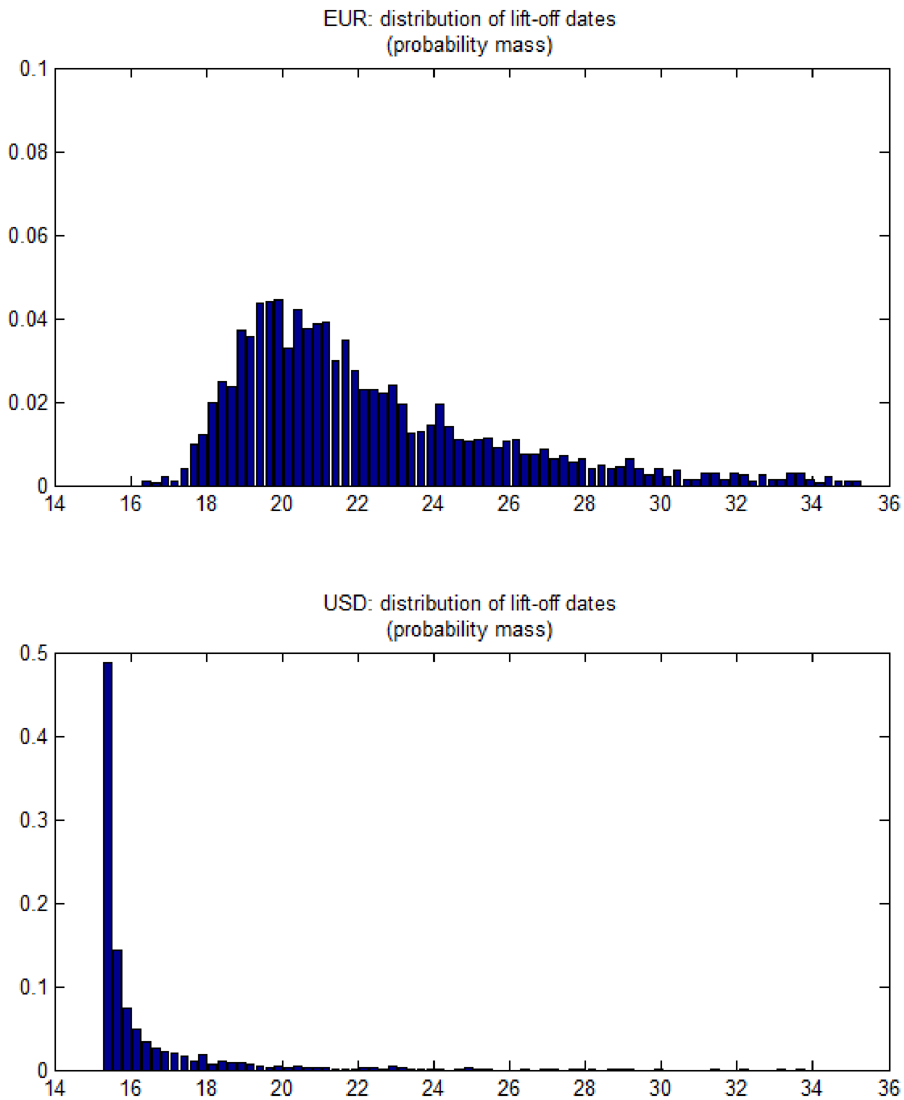


Figure 4 - 10-year risk premia

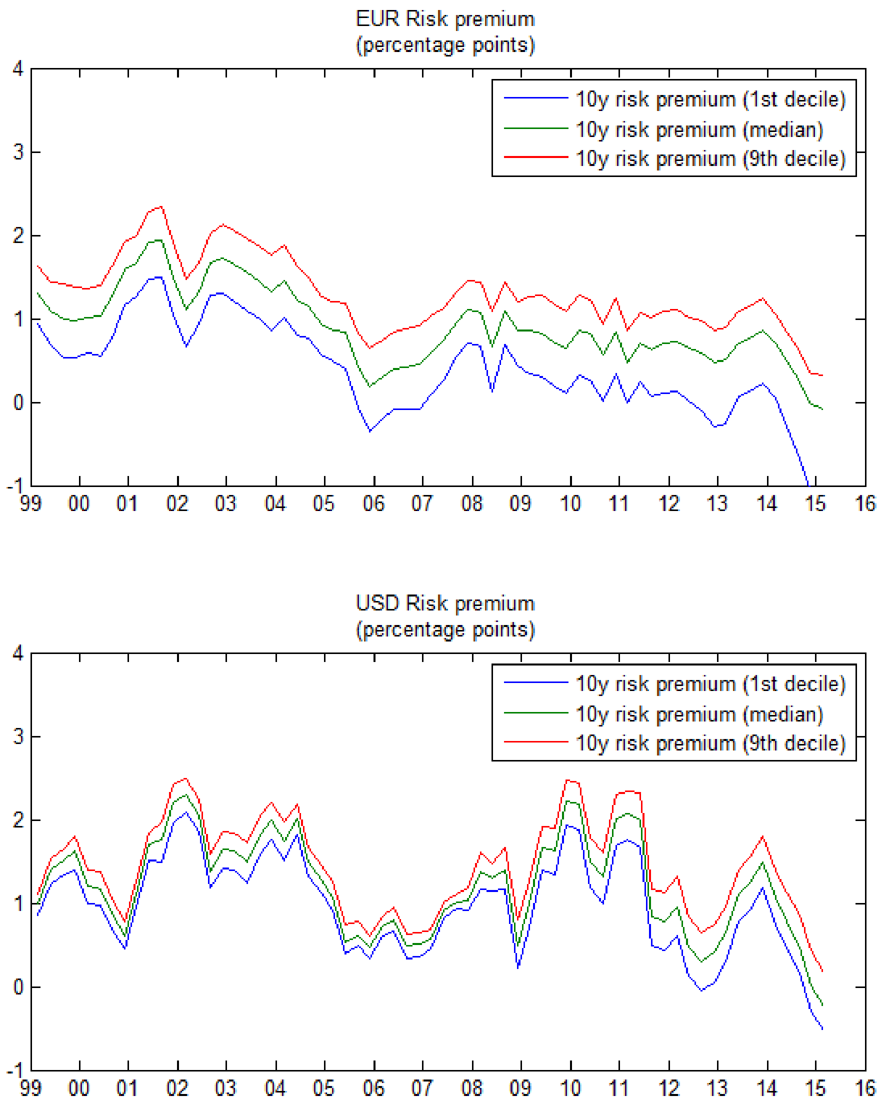
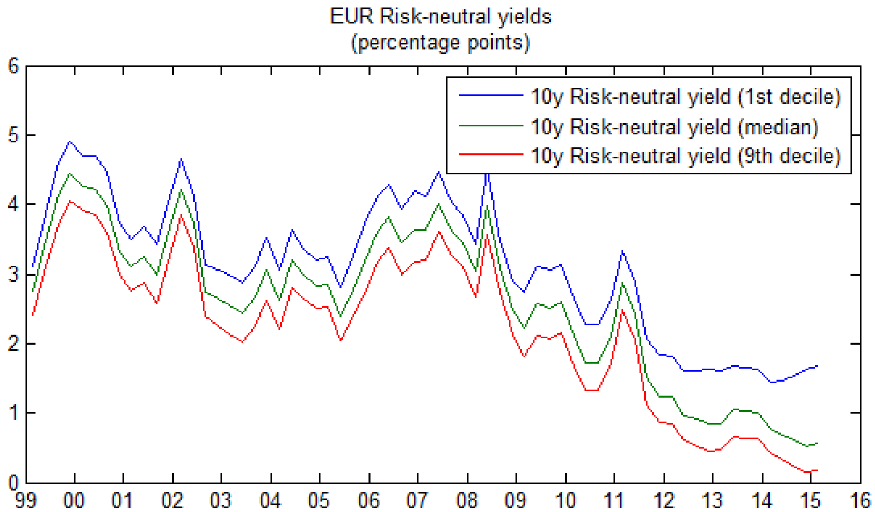


Figure 5 - 10-year risk-neutral yields



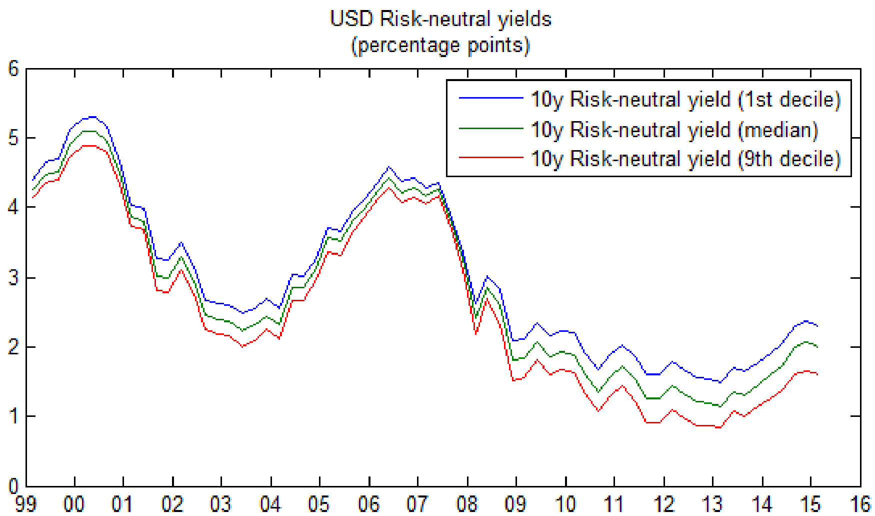
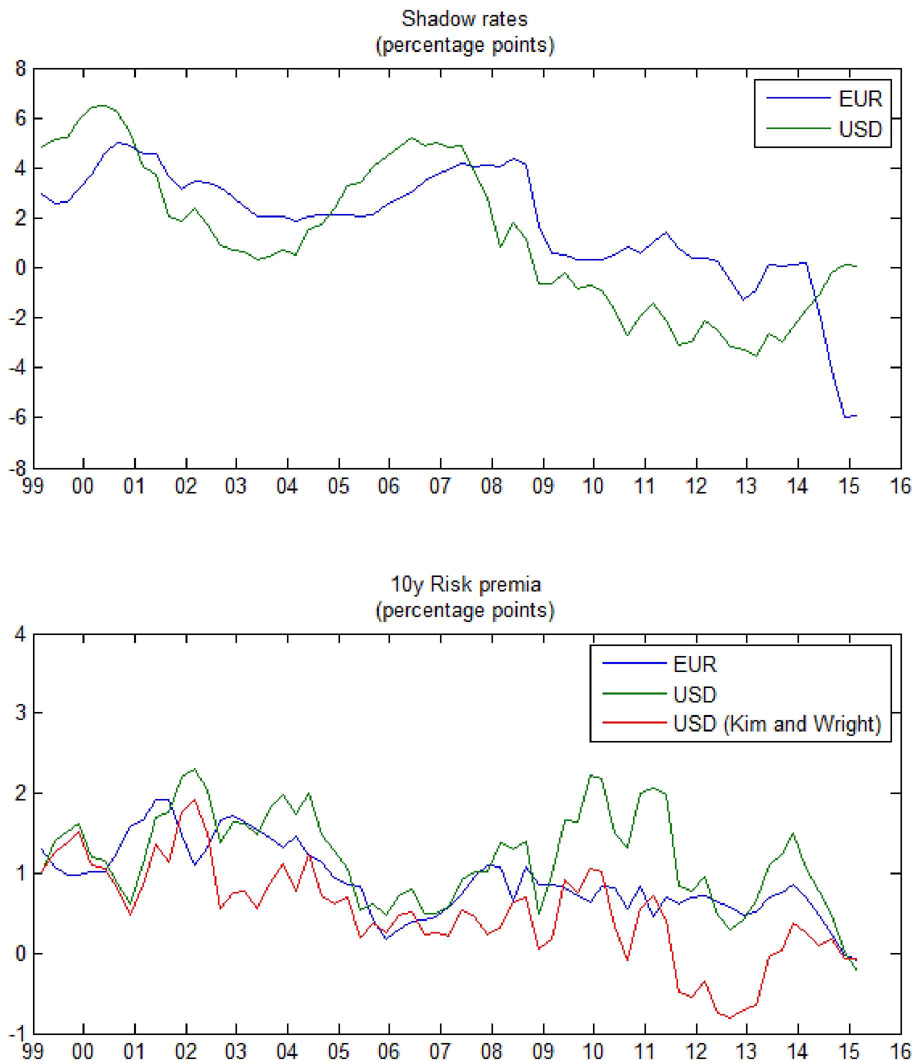


Figure 6 - Comparisons of shadow rates and risk premia



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