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Forecasting the Yield Curve with Linear Factor Models

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Abstract

In this work we compare the interest rate forecasting performance using a broad class of linear models. The models are estimated through a MCMC procedure with data from the US and Brazilian markets. We show that a simple parametric specification has the best predictive power, but it does not outperform the random walk. We also find that macroeconomic variables and no-arbitrage conditions have little effect to improve the out-of-sample fit, while a financial variable (stock index) increases the forecasting accuracy.

JEL classification: G1, E4, C5.

Keywords: Yield curve forecasting, macroeconomic variables, affine models.

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1 Introduction

Modeling the term structure of interest rates is a challenging task that has, from a practical perspective, at least three purposes. Firstly, with this tool one can price fixed-income instruments and manage the risk of bonds and derivatives. Secondly, it allows monitoring observed and unobserved economic variables such as the risk premium, default risk, inflation and real activity. Finally, it allows forecasting future interest rates. In this study, we address the latter issue, using a rich class of linear factor models.

Users of yield curve forecasts are numerous. Treasuries manage the emission and maintenance of the stock of public debt, which continuously demands an assessment of current and future interest rates. Investors must track their portfolios' performance against the opportunity cost of investing in low-risk bonds. Central banks react to expected inflation and economic activity by adjusting the short rate, thereby affecting the whole curve.

Term structure models can be classified in different ways. If restrictions on the evolution of the yields are imposed in order to avoid risk-free profit opportunities, then the model is known as arbitrage-free. Otherwise the model is said to be purely statistical. Arbitrage-free models contain some ingredients arising from equilibrium models and thus have strong economic appeal. Seminal works within this class are Vasicek (1977), Cox et al. (1985) and Heath et al. (1992), while Nelson and Siegel (1987) and Svensson (1994) are pioneer works in the class of statistical models. Moreover, term structure models may or may not directly include macroeconomic and financial factors driving the yield curve. Among others, we can cite the works of Ang and Piazzesi (2003), Diebold et al. (2005), Hördahl et al. (2006), and Ludvigson and Ng (2007), all of which used macroeconomic variables to model the term structure of interest rates. Finally, the relation between factors and interest rates may be linear or assume a more general specification. Examples of linear models are the class of affine models studied by Duffie and Kan (1996)[1], while Ahn et al. (2002) and Leippold and Wu (2002) constitute examples of non-linear models.

Although several works from macroeconomics, finance and econometrics have been devoted to term structure models, few of them have analyzed the out-of-sample forecasting performance. Predictability questions regarding yield curve are firstly studied by Fama and Bliss (1987), who investigated the relationship between forward and future spot rates. More recently, Duffee (2002) shows that the affine models produce poor US yields forecasts.

[1]An affine model is an arbitrage-free term structure model, such that the state process X is an affine diffusion, and the yields are also affine in X.

Diebold and Li (2006) propose a two-stage model based on the Nelson and Siegel (1987) framework to forecast the US term structure that presented better results than some competing models. Nevertheless, Almeida and Vicente (2008) show that the inclusion of no-arbitrage conditions in a latent model improves the out-of-sample fit. Ang and Piazzesi (2003) find that an affine model with macroeconomic variables outperforms the unrestricted VAR model containing the same observable factors. They also show that models with macro factors forecast better than models with only unobservable factors. In the same line, Hördahl et al. (2006) confirm that the forecasting performance of a model with observable factors is superior to models based on latent factors.

The papers mentioned above deal with models where the interest rates are linear functions of state factors (observable or latent). For a variety of reasons (the most important is the simplicity of implementation), linear models are nowadays the workhorse of yield curve modeling. However, to the best of our knowledge there is no study based on the same dataset that provides a full comparison of the out-of-sample performance of different specifications of linear models [2]. In this article we try to fill this gap in the finance literature. We analyze arbitrage-free and purely statistical models, with or without observable variables. In addition to testing the models with US data as usual, we also consider a database from an emerging country, Brazil.

All the models analyzed in this study present constant volatility[3]. Although stochastic volatility processes have some nice properties, the standard approach in the interest rate forecasting has been to use homocedastic models. Besides their parsimony, constant volatility models seems to be a natural choice when the aim is forecasting, since in this family there is no factor collecting information about the volatility process. Therefore, it is expected that the mean of the yields distribution can be better captured. Duffee (2002) tests the forecasting power of affine models and shows that the Gaussian specification outperforms non-constant volatility models, which suggest that this intuition is true. Furthermore, models with constant volatility do a good job of explaining some stylized facts (as shown, for instance, by Dai and Singleton, 2002 and Bikbov and Chernov, 2004).

We estimated the models using Monte Carlo Markov Chain (MCMC) (see Johannes and Polson, 2006), a Bayesian approach that does not require any maximization, only the repeated sampling of complete conditional

[2]Although some recent studies have addressed the forecasting performance of different linear interest rate models (besides the works cited above, we can also include Vicente and Tabak, 2008, and Möench, 2008), we believe that no one has implemented a comparative analysis as comprehensive as ours.

[3]Constant volatility models are also known as Gaussian models.

distributions. This method obtains distributions of all parameters (and, consequently, functions of parameters such as forecasts) conditional on the data. Thus, it is a tool that permits the measurement of the degree of uncertainty associated with the available information for estimating a given model.

Our main results can be summarized as follows. Firstly, linear models have poor forecasting power. When the benchmark is the random walk they do not do a good job[4]. Secondly, the class of parametric models estimated in a two-step process outperforms the other competitors. Thirdly, the inclusion of macroeconomic variables does not improve the out-of-sample fit, but the inclusion of a financial variable (a stock index) contributes positively to forecasting accuracy. Fourthly, arbitrage-free models, represented in this study by the affine family, exhibit low predictive power. Finally, we find that imposing zero mean to the in-sample error increases the forecasting ability of the models.

The rest of this article is organized as follows. In Section 2 we present the models. Section 3 details the estimation procedure. Section 4 describes the dataset used. In Section 5 we discuss the results of implementing the models and in Section 6 we offer our concluding remarks.

2 Models

There are a wide variety of interest rates models that can be used to forecast the yield curve. Both no-arbitrage models, which have microeconomic foundations and are estimated jointly, and the two-step model of Diebold and Li (2006), which aims only at statistical adherence, provide legitimate predictive tools. Our purpose is to compare the out-of-sample fit of the most common linear interest rate models.

In general, linear models can be specified through a state-space system. Let $Y_t = (Y_t^1, \ldots, Y_t^N)^\intercal$ be the vector of interest rates at time t and $X_t = (M_t, \theta_t)$ the state vector composed of p observable factors M_t and q latent factors θ_t. The general formulation for the linear model with constant volatility is:

$$\begin{aligned} Y_t &= A(\tau,\Psi) + B^M(\tau,\Psi)M_t + B^\theta(\tau,\Psi)\theta_t + \sigma u_t \qquad (1) \\ &= A(\tau,\Psi) + B(\tau,\Psi)X_t + \sigma u_t,\ u_t \sim N(0,\mathbf{I}_N) \ \text{ and} \end{aligned}$$

$$X_t = \mu + \Phi X_{t-1} + \Sigma\varepsilon_t,\ \varepsilon_t \sim N(0,\mathbf{I}_{p+q}), \qquad (2)$$

[4]In order to improve the predictive power we try some variations of the models. For instance, in the estimation procedure we use a moving window instead of a fixed window. However the results are just slightly better.

where $\sigma \in \mathbb{R}^{N\times N}$, $\mu \in \mathbb{R}^N$, and $\Phi, \Sigma \in \mathbb{R}^{(p+q)\times(p+q)}$. Each model is characterized by the functions $A(\tau,\Psi) \in \mathbb{R}^N$, $B^M(\tau,\Psi) \in \mathbb{R}^{N\times p}$ and $B^\theta(\tau,\Psi) \in \mathbb{R}^{N\times q}$, where τ is a vector of the time to maturities of yields and Ψ is a vector stacking the model parameters[5]. In the next subsections we describe the competitor forecasting models used in this work.

2.1 Joint models

In the class of joint models the observation and state equations (1 and 2, respectively) are estimated simultaneously. They include the affine family and dynamic versions of parametric models estimated in a single step.

2.1.1 Affine model

This is a standard model that precludes arbitrage-free profit opportunities. Following the approach of Ang and Piazzesi (2003), we derive the discrete-time version of the affine model (na)[6]. Basically we have to include the no-arbitrage condition in the general specification defined by (1) and (2).

The instantaneous short term rate is given by $r_t = \delta_0 + \left(\delta_1^M, \delta_1^\theta\right)^\intercal (M_t, \theta_t) = \delta_0 + \delta_1 X_t$. We assume the existence of a pricing measure $\mathbb{Q}$ under which discounted security prices are martingales with respect to the filtration $(\mathcal{F}_t)_{t\geq 0}$ ($\mathcal{F}_t$ represents the information available at time t). The connection between the pricing and the objective probability measures is given by an extended affine market price of risk (Cheridito et al., 2007), $\lambda_t = \lambda_0 + \lambda_1 X_t$. The price at time t of a zero-coupon bond that pays \$1 at maturity date $t + \tau_n$ is $p_t^n = \exp(\alpha_n + \beta_n^\intercal X_t)$, where

$$\begin{aligned} \beta_{n+1}^\intercal &= -\delta_1^\intercal(1 + \Phi^\star + \ldots + \Phi^{\star n}), \\ \alpha_{n+1} &= -\delta_0 + \alpha_n + \beta_n^\intercal \mu^\star + \frac{1}{2}\beta_n^\intercal \Sigma\Sigma^\intercal \beta_n, \end{aligned} \tag{3}$$

with initial conditions $\alpha_1 = -\delta_0$, $\beta_1 = -\delta_1$, and $\Phi^\star = \Phi - \Sigma\lambda_1$, $\mu^\star = \mu - \Sigma\lambda_0$. Then $Y_t^n = -\log p_t^n/\tau_n = A_n + B_n^\intercal X_t$, where $A_n = -\alpha_n/\tau_n$ and $B_n = -\beta_n/\tau_n$. Therefore $Y_t = A + BX_t$, where $A = (A_1, \ldots, A_N)$ and B is an $N \times N$ matrix with rows given by B_n. Adding the error term in this equation we obtain (1).

Dai and Singleton (2000) show that some parameter restrictions are necessary to identify the model, because there are combinations of the parameters and state variables that generate the same the yield curve. Consequently

[5]The vector Ψ is model dependent.

[6]For an extensive analysis of the affine family, see Duffie and Kan (1996).

there are multiple sub-identified models corresponding to the same data. However, several equivalent alternatives can be used to achieve the identification. In particular, we follow Matsumura (2008) and set $\Phi^\star_{\theta M} = 0$, $\Phi^\star_{\theta\theta} = 0$, $\delta_1^\theta = 1$[7]. Thus $\Phi^\star$ and $(\mathbf{I}_{p+q} + \Phi^\star + \ldots + \Phi^{\star n})$ are upper triangular.

It turns out that if we also impose $\delta_1^M = 0$, in which case the short rate follows an infinite no-discount forward looking Taylor rule (see Ang et al., 2007), then $B^M = 0$. Hence, setting $B^M = 0$ can be justified in an identified affine model as a choice of monetary policy consistent with a Taylor rule. Based on this remark we set $B^M = 0$ in all other models[8].

2.1.2 Quasi-affine model

Besides the usual affine model, we propose a model with the same specification for B, but in which A is determined so that the sample mean of the error term in (1) is zero. Since $B^M = 0$ and the stochastic process θ has zero mean (an identification condition of the affine model, see Matsumura, 2008), we have $A = \overline{Y}$, which simplifies the estimation procedure. Then (1) and (2) are replaced by

$$Y_t = \overline{Y} + B^\theta(\cdot)\theta_t + \sigma u_t,\ u_t \sim N(0, \mathbf{I}_N) \text{ and} \tag{4}$$

$$X_t = \Phi X_{t-1} + \Sigma\varepsilon_t,\ \varepsilon_t \sim N(0, \mathbf{I}_{p+q}), \tag{5}$$

where in (5) we fix $\mu = 0$ by subtracting the mean of M from X. The loading $B^\theta(\cdot)$ is the same as in the affine model. We call this model as quasi-affine (qa). It is more flexible and easier to estimate than the affine model, but relaxes the no-arbitrage condition.

2.1.3 Nelson-Siegel model

We analyzed two versions of the Nelson and Siegel (1987) model - henceforth NS model. The first one is the standard approach (sns) defined by:

$$Y_t = B^\theta(\gamma)\theta_t + \sigma u_t,\ u_t \sim N(0, \mathbf{I}_3) \text{ and} \tag{6}$$

$$X_t = \mu + \Phi X_{t-1} + \Sigma\varepsilon_t,\ \varepsilon_t \sim N(0, \mathbf{I}_{p+q}), \tag{7}$$

where

$$B_n^\theta = \left(1, \left(1 - \mathrm{e}^{-\gamma\tau_n}\right)/\gamma\tau_n, \left(1 - \mathrm{e}^{-\gamma\tau_n}\right)/\gamma\tau_n - \mathrm{e}^{-\gamma\tau_n}\right). \tag{8}$$

[7] Matrix $\Phi^\star$ can be split as $\begin{bmatrix} \Phi^\star_{MM}\Phi^\star_{M\theta} \\ \Phi^\star_{\theta M}\Phi^\star_{\theta\theta} \end{bmatrix}$.

[8] Other authors adopted this same restriction for B^M (see, for instance, Diebold et al., 2006).

In the second version, we add the loading A in the transition equation. As in the quasi-affine model, we set $A = \overline{Y}$ and $\mu = 0$:

$$Y_t = \overline{Y} + B^\theta(\gamma)\theta_t + \sigma u_t,\ u_t \sim N(0, \mathbf{I}_3), \tag{9}$$

$$X_t = \Phi X_{t-1} + \Sigma\varepsilon_t,\ \varepsilon_t \sim N(0, \mathbf{I}_{p+q}). \tag{10}$$

Note that the inclusion of the loading A in the standard version of the NS model allows us to exactly fit the long-term mean of the yield curve and ensures that the measurement errors u_t have zero mean. We call this version the extended NS model (ens).

2.1.4 Legendre model

The Legendre model (see Almeida et al., 1998) is very similar to the NS model. The only difference is the parametric form of the loadings. In the Legendre model (lg) they are a sequence of polynomials:

$$B_n^\theta = \left(1, x_n, \frac{1}{2}(3x_n^2 - 1), \frac{1}{2}(5x_n^3 - 3x_n)\right), \tag{11}$$

where $x_n = 2\tau_n/\ell - 1$ and ℓ is the longest maturity in the bond market.

2.1.5 Common factor model

Litterman and Scheickman (1991) show the yield variations can be summarized by three independent movements. In other words, since the yields of different maturities are highly correlated, one can reduce the dimension of the interest rate space without losing significant information. Based on this fact, we proposed a common factor (cf) model (see West, 1997) to forecast the yield curve. In the common factor model there are only identification restrictions. We do not impose any economic or parameterization condition. The vector space X is composed of latent factors and its dimension must be less than N:

$$Y_t = BX_t + \sigma u_t,\ u_t \sim N(0, \mathbf{I}_N), \tag{12}$$

with $X_t = \theta_t$ and $q < N$. In the empirical forecasting exercise presented in Section 5 we estimated the cf model using two latent factors, that is, $q = 2$.

2.2 Conditional models

In the class of conditional models, we use a two-step estimation procedure. First, we estimate the observation equation. Next, we estimate the state equation.

2.2.1 Diebold-Li model

The Diebold and Li (2006) model - henceforth DL model - is very similar to the standard NS model. The interest rates follow the same dynamics given by (6):

$$Y_t = B^{\theta}\theta_t + \sigma u_t,\ u_t \sim N(0, \mathbf{I}_N). \tag{13}$$

However, while we estimate the NS model in a single step using MCMC (see Section 3), DL propose a simpler procedure. First, they set the lambda parameter at $\gamma = 0.0609$, to maximize the curvature of the term structure (the third loading on (6)) at 30 months, and estimated θ_t for all t by ordinary least squares:

$$\theta_t = \text{argmin} \sum_t \left(Y_t - B^{\theta}\theta_t\right)^2. \tag{14}$$

In this work, the value of γ is also kept fixed. However, we use the procedure proposed by Almeida et al. (2009) to choose the γ value to be adopted. The idea is to search for a γ under which the model generates its best in-sample fit.

The next step is to assume that the latent factors follow an AR (dla) or VAR (dlv) process, which is used to forecast the θ's and consequently the yield curve:

$$\widehat{\theta}_t = \mu + \Phi\widehat{\theta}_{t-1} + \Sigma\varepsilon_t,\ \varepsilon_t \sim N(0, \mathbf{I}_{p+q}). \tag{15}$$

In the version with macro factors, we have:

$$\widehat{\theta}_t = \mu^{\theta} + \Phi^{\theta M}M_{t-1} + \Phi^{\theta\theta}\widehat{\theta}_{t-1} + \Sigma^{\theta M}\varepsilon_t^M + \Sigma^{\theta\theta}\varepsilon_t^{\theta}. \tag{16}$$

The same two-step procedure can be used in the Legendre model. That is, the dynamics of θ in the Legendre model can be set as an AR (lga) ou VAR (lgv) process.

2.2.2 AR and VAR models

In order to compare the interest rate models described previously with traditional econometric techniques, we also consider AR and VAR models.

The AR model (ar) uses a yield only approach in which the dynamic of the interest rate with time to maturity n is given by

$$Y_t^n = \mu + \Phi_n Y_{t-1}^n + \sigma_n u_t^n. \tag{17}$$

Since the yields are highly correlated, instead of the standard VAR model, we use a simpler version (var) in which the explanatory variables have a lower dimension,

$$Y_t^n = \mu + \Phi_n Z_{t-1} + \sigma_n u_t^n. \tag{18}$$

In the yields-only version we set $Z_t = (Y_t^1, Y_t^N)$ and in the macro/finance version we set $Z_t = (Y_t^1, Y_t^N, M_t)$, where N denotes the longest maturity.

3 Inference

The joint models are estimated via the MCMC method, while the DL models (NS and Legendre versions) are estimated by the simpler procedure described in Section 2.2.1. The AR and VAR models are estimated using ordinary least squares. The inclusion of macro variables makes the inference task more difficult due to the optimization problems in high dimension spaces and non-linearity in the parameters. This fact motivated us to use the MCMC algorithm, a Bayesian approach less vulnerable to these issues than the traditional maximum likelihood technique. General references about MCMC are Robert and Casella (2004) and Gamerman and Lopes (2006). For the specific case of financial econometrics, the work of Johannes and Polson (2006) is very useful.

MCMC is a method to obtain the joint distribution $f(\Psi, \theta|M, Y)$ of the parameters and latent variables conditional on observed data. Although $f(\Psi, \theta|M, Y)$ is generally unknown and extremely complex, the Clifford-Hammersley theorem guarantees that if some technical conditions are satisfied, then it can be uniquely characterized by the lower dimensional distributions $f(\Psi|M, Y, \theta)$ and $f(\theta|M, Y, \Psi)$. These distributions, in turn, can be characterized by even lower dimensional distributions. For instance, if the set of parameters is divided into subsets, $\Psi = (\Psi_1, \ldots, \Psi_k)$, then the distributions $f(\Psi_i|\Psi_{-i}, M, Y, \theta)$ determine $f(\Psi|M, Y, \theta)$. Using Gibbs sampling or the Metropolis algorithms, the full conditional distribution can be recovered from lower dimensional ones, avoiding tricky non-linear optimizations.

The Bayesian approach provides a posterior distribution of the time-series path of X_t. Consequently, we can easily make forecasts of the yield curve. The advantages of the MCMC method encouraged us to use it. Other recent studies that deal with the yield curve modeling using macro variables adopt the MCMC estimation procedure (see, for instance, Ang et al., 2007). The details of the implementation of the MCMC for each model presented in Section 2 can be found in Matsumura (2008).

4 Data

We use data on US zero coupon bond yields with maturities of 1, 3, 6, 12, 24, 36, 60, 84 and 120 months from January 1987 to March 2009. The sampling period begins a few months before Alan Greenspan succeeded Paul Volcker as Fed chairman in the summer 1987, and two months before the crash of the New York Stock Exchange. The appointment of Alan Greenspan as Fed chairman is considered by several authors a change in US monetary policy (see, for instance, Bernanke and Woodford, 2006)[9]. The macro variables are of two types: an inflation measure, represented by the Consumer Price Index (CPI), and a real activity measure, represented by the output gap. All these data were taken from the Fed database and collected at a monthly frequency[10]. The finance variable is represented by the log of the Dow Jones Industrial Average index provided by Bloomberg. We split this database into two parts. The first, composed of 197 monthly observations from January 1987 to June 2003, is the in-sample period in which the estimations of the models are made. The second, from July 2003 to March 2009, is the out-of-sample period in which the forecasting power of the models is evaluated. Figure 1 shows the evolution of the zero coupon bond yields with maturities of 1, 12, 60, and 120 months. The yields are decreasing over time, varying from 10% at the beginning of the sample to 1% at the end of the sample. Figure 2 plots the time series of the US observable factors. Note that the output gap and the Dow Jones fell at the end of the sample as a consequence of the subprime crisis.

Besides the US market, we also analyze the forecasting performance of the linear models using data from the Brazilian economy. This allows us to test the models in an emerging country where idiosyncrasies, such as a short yield curve and imperfect market, are present. Brazil is one of the most important emerging countries having the largest equity and bond markets in Latin American. The starting point in our Brazilian sample is January 1999, when Brazil adopted the floating exchange rate regime after a devaluation of the local currency. The in-sample period ends in March 2006, and the forecasting exercise uses 36 months of data afterwards. Brazilian spot yields with maturities of 1, 2, 3, 6, 9, 12, 18, 24 and 36 months are extracted from the `ID x Pre` fixed-for-floating rate swap, an instrument traded in the BM&F Bovespa[11]. Similar to the US, the Brazilian macro variables are the Comprehensive Consumer Price Index (IPCA) and the industrial output

[9]Since economic agents can anticipate the changes in Fed's Board, we decided to start the analysis few months before the summer 1987.

[10]The data are available at the website
http://www.federalreserve.gov/econresdata/releases/statisticsdata.htm.

[11]BM&F Bovespa is the main Brazilian derivatives exchange and one of the world's largest according to the Futures Industry Association's report (see Burghardt and Acworth, 2008). For more information about the `ID x Pre` swap, see the BM&F Bovespa website, www.bmf.com.br/portal/Home2/portal_english.asp.

gap. The IPCA is the main consumer price index in Brazil. We constructed the output gap ourselves by modeling industrial production as the sum of a trend and a cyclical component using the Hodrick and Prescott (1997) filter. The finance variable is the log of the Ibovespa, which is the main Brazilian stock market index. The IPCA, industrial production, and the Ibovespa were obtained from the Institute of Applied Economic Research (IPEA) website, http://www.ipea.gov.br. Figure 3 shows the evolution of the zero coupon bond yields with maturities of 1, 12, and 36 months. At the beginning and end of the sample, yields are increasing and the term structure is upward sloping (long-term rates higher than short-term rates). However, between 2004 and 2007 the shape of the term structure changes to downward sloping. Figure 4 shows the time series of the Brazilian observable factors. Note that apart from the 2002 electoral period, Brazilian inflation was at roughly the same level as American inflation. As in the US, real activity and stock returns plummeted at the end of the sample due to the subprime crises.

5 Results

In this section we analyze the predictive power of the models presented in Section 2 using US and Brazilian data. To compare the out-of-sample forecasting performance of the models, we choose the random walk (RW) as the benchmark. If the processes under study have high persistency, RW frequently adheres well to the data, sometimes better than more sophisticated models. Closely related with RW is the Theil-U (TU) statistic, defined by:

$$TU(n) = \left(\sum_{t_{\text{out}}} (Y_{t+h}^n - \widehat{Y}_{t+h|t}^n)^2 / \sum_{t_{\text{out}}} (Y_{t+h}^n - Y_t^n)^2 \right)^{\frac{1}{2}},$$

where a hat indicates forecast. In other words, the TU statistic is just the ratio between the root mean squared errors (RMSE) of a particular model and the RMSE of the RW. We consider two forecasting horizons. For the US curve we use h equals 1 and 12 months. The Brazilian curve is shorter, therefore we set h equal to 1 and 6 months.

Since there are several models, instead of reporting the TU for each maturity, we summarized the results through three accuracy measures based on the TU statistic:

- The t criterion is the number of maturities such that the expected value of TU is less than one (model better than RW). Note that this expression can be computed since MCMC provides sample distributions of any function of the parameters. In the class of conditional models the t criterion is the number of maturities such that the TU is less than one.

- The s criterion is number of maturities such that $E(\text{TU})+1.65\sigma(\text{TU}) < 1$. Note that the criterion s amounts to a statistically significant t. The s criterion is not defined for the conditional models.

- The d criterion is number of maturities such that the Diebold and Mariano (1995) statistics is higher than 1.65 (indicating a significance at a 90% confidence level).

Table 1 shows the t, s and d criteria for 1- and 12-month ahead forecasts of the US yields. In the macro version we use the output gap and the inflation index as the observable variables. In the financial version we take the stock index as the observable variable. For the na model, the MCMC procedure only converges with one latent factor or with one latent plus one financial variable. For the qa model we obtain convergence using two latent factors (with or without observable factors)[12]. In general, none of the models forecast the yield curve well. Certainly, the subprime crisis is one of the reasons for this weak forecasting performance. The models are estimated under market conditions very different from the nervous situation that appears at the end of the out-of-sample period. None of the models consistently outperforms the RW. The parametric models (NS, Legendre and common factors) estimated jointly clearly present the best predictions among the models tested. The inclusion of observable macroeconomic variables seems to have little effect on the out-of-sample fit. The same is true for the no-arbitrage condition, since the affine model shows poor forecasting power. However, the financial variable (Dow Jones Index) contributes positively to the out-of-sample fit, particularly in the long horizon forecast (12 months). Table 2 presents the t, s and d criteria for 1- and 6-month ahead forecast of the Brazilian yields[13]. The performance of the models in the Brazilian market is worse than in the US market. Some reasons, such as the small sample, market imperfections and liquidity problems can explain this fact. Only the conditional models have forecasting power similar to the RW.

Comparing our findings with other works, we have some interesting conclusions. Firstly, we agree with Vicente and Tabak (2008), who show that exponential parametric models present better predictive ability than affine

[12] The convergence of the chains are monitored by the Gelman and Rubin (1992) diagnostics. In the interest of saving space, we do not report the Gelman-Rubin statistic. However, they are available upon request.

[13] The na model does not converge with Brazilian data. Thus we do not report results of this model for Brazil.

Gaussian models. However, this does not represent a final answer about the inclusion of no-arbitrage conditions. As pointed by Almeida and Vicente (2008), to precisely address this question it is necessary to compare models in which the only difference among them is the no-arbitrage restriction. However, Filipovic (1999) shows that there is no-arbitrage-free version of the NS model. Secondly, we do not manage to reproduce the results of Ang and Piazzesi (2003), since the incorporation of macro factors does not significantly improve the forecasting power of the models. As the dataset used by them is different from ours, this suggests that the sample period considered affects the results. Thirdly, we confirm the findings of Duffee (2002), who provides evidence that the affine class is not an appropriate tool to forecast interest rates. We also present an empirical illustration of the theoretical result obtained by Joslin et al. (2010). They show that within Gaussian models, enforcing no-arbitrage has no effect on out-of-sample forecasts of the yield curve. Another important conclusion of our study is that complex models do not necessarily do a good job of improving the out-of-sample fit of the yield curve. The NS model is very simple, but it is still the model that best predicts interest rates. Finally, the estimation in one step via the MCMC procedure seems to contribute positively to the forecasting accuracy because the joint models outperform the conditional ones.

6 Conclusion

We studied different classes of linear term structure models in order to assess the comparative advantages concerning out-of-sample forecasts. The method used to estimate the main models is the MCMC procedure. The MCMC technique is a Bayesian approach that avoids some problems usually observed in likelihood methods. We analyzed two different economies: a developed market, represented by the US, and an emerging market, represented by Brazil. In general, the models have poor predictive power. The parametric models have the best forecasting accuracy. However they do not consistently outperform the random walk. Our results indicate that the inclusion of macroeconomic variables and the no-arbitrage restriction does not improve the out-of-sample fit. On the other hand, the financial variable seems to add important information to capture the yield dynamics. In the Brazilian market the forecasting errors are much larger than those observed for the US market. In this market, we note a slight superiority of the parametric models estimated in two-step.

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Model	Yields only						Macro						Finance					
	$t1$	$t6$	$d1$	$d6$	$s1$	$s6$	$t1$	$t6$	$d1$	$d6$	$s1$	$s6$	$t1$	$t6$	$d1$	$d6$	$s1$	$s6$
na	0	0	0	0	0	0	-	-	-	-	-	-	0	0	0	0	0	0
qa	0	0	0	0	0	0	0	0	0	0	0	0	0	0	1	0	0	0
cf	0	1	2	4	0	2	0	0	1	6	0	2	0	0	2	5	0	1
ens	0	0	2	4	0	3	0	2	3	4	0	4	0	0	3	4	1	3
sns	0	0	0	1	0	0	0	3	0	4	0	4	0	0	0	6	0	5
lg	0	0	2	4	2	3	0	2	3	5	0	4	0	0	3	4	2	3
var	0	0	1	4	-	-	1	0	3	1	-	-	0	0	1	8	-	-
dlv	0	0	1	4	-	-	0	0	4	3	-	-	0	0	2	6	-	-
lgv	0	1	2	1	-	-	1	0	4	0	-	-	0	0	3	4	-	-
ar	0	0	3	2	-	-	-	-	-	-	-	-	-	-	-	-	-	-
dla	0	0	2	0	-	-	-	-	-	-	-	-	-	-	-	-	-	-
lga	0	0	2	1	-	-	-	-	-	-	-	-	-	-	-	-	-	-

Table 1: Results of the US yield forecasting.

This table presents the t, s, and d criteria for 1- and 12-month ahead forecasting of the US yields with maturities of 1, 3, 6, 12, 24, 36, 60, 84 and 120 months. The t criterion represents the number of maturities such that $E(TU) < 1$. The s criterion is the number of maturities such that $E(TU) + 1.65\sigma(TU) < 1$. The d criterion is the number of times such that the Diebold and Mariano (1995) statistic is higher than 1.65 across the maturities.

Model	Yields only						Macro						Finance					
	$t1$	$t6$	$d1$	$d6$	$s1$	$s6$	$t1$	$t6$	$d1$	$d6$	$s1$	$s6$	$t1$	$t6$	$d1$	$d6$	$s1$	$s6$
na	-	-	-	-	-	-	-	-	-	-	-	-	-	-	-	-	-	-
qa	0	0	1	0	1	0	-	-	-	-	-	-	0	0	2	1	1	1
cf	0	0	3	0	0	0	0	0	0	0	0	0	0	0	1	0	0	0
ens	2	1	2	1	0	0	0	0	1	0	0	0	1	0	2	0	0	0
sns	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0
lg	2	0	2	0	2	0	0	0	1	0	0	0	2	0	2	0	2	0
var	2	0	3	0	-	-	0	0	0	0	-	-	3	0	4	0	-	-
dlv	2	0	3	0	-	-	0	0	1	0	-	-	3	0	4	0	-	-
lgv	2	0	3	0	-	-	0	0	1	0	-	-	3	0	4	3	-	-
ar	0	0	0	0	-	-	-	-	-	-	-	-	-	-	-	-	-	-
dla	0	0	0	0	-	-	-	-	-	-	-	-	-	-	-	-	-	-
lga	0	0	2	0	-	-	-	-	-	-	-	-	-	-	-	-	-	-

Table 2: Results of the Brazilian yield forecasting.

This table presents the t, s, and d criteria for 1- and 6-month ahead forecasting of the Brazilian yields with maturities of 1, 2, 3, 6, 9, 12, 18, 24 and 36 months. The t criterion represents the number of maturities such that $E(TU) < 1$. The s criterion is the number of maturities such that $E(TU) + 1.65\sigma(TU) < 1$. The d criterion is the number of times such that the Diebold and Mariano (1995) statistic is higher than 1.65 across the maturities.

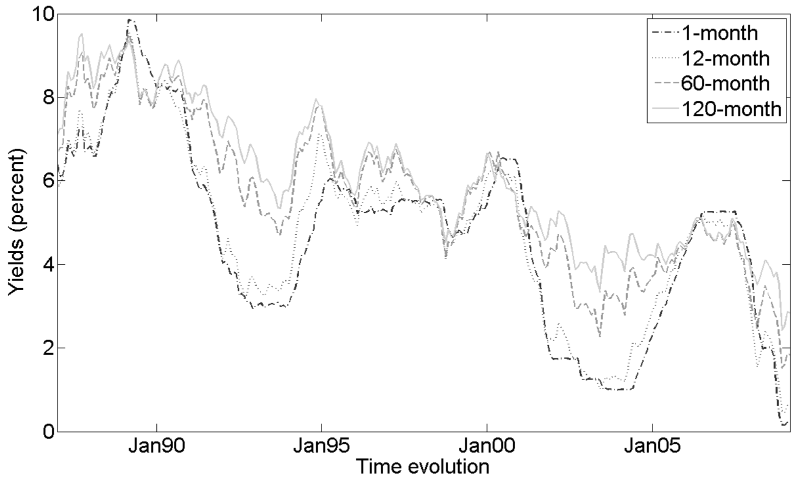

Figure 1: US zero coupon bond yields.
This figure contains the time series of US (annualized) monthly zero coupon bond yields with maturities of 1 month, 12 months, 60 months and 120 months between January 1987 and March 2009.

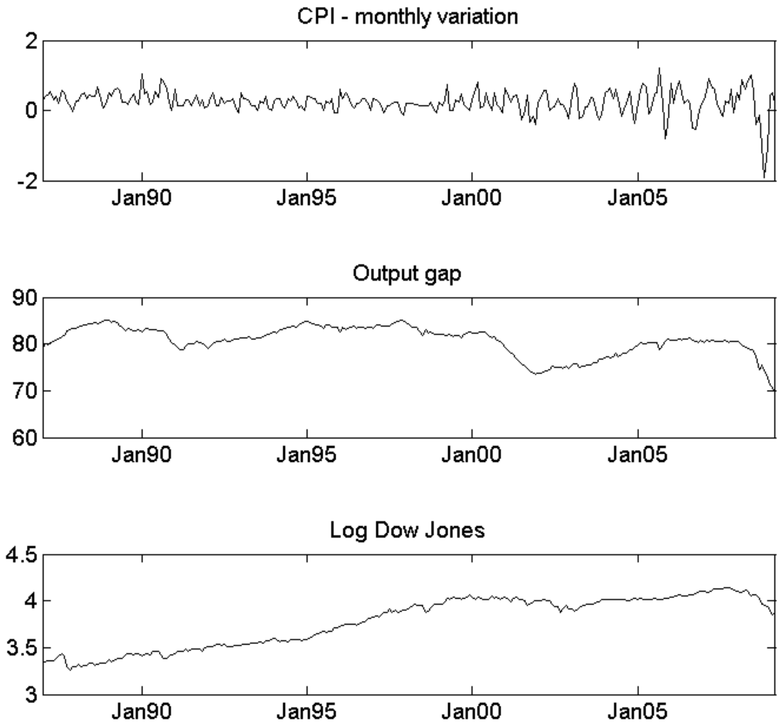

Figure 2: US observable factors.

This figure contains the time series of US observable factors between January 1987 and March 2009. The top panel shows the evolution of the monthly variation of the CPI; the central panel depicts the evolution of the output gap; and the bottom panel shows the log of the Dow Jones Index.

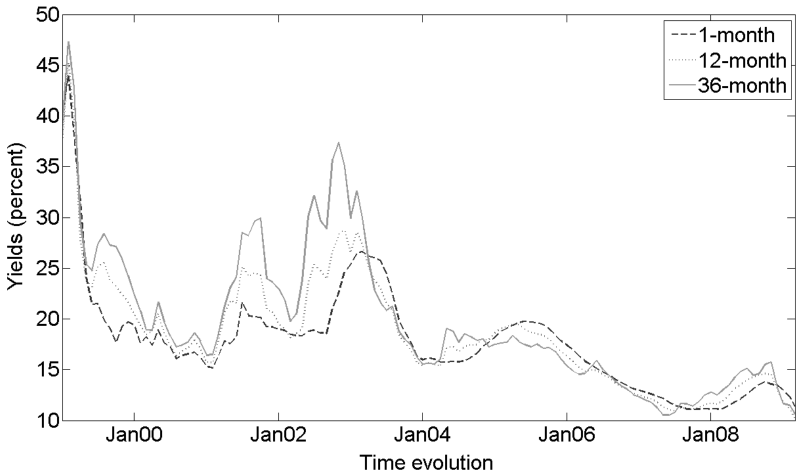

Figure 3: Brazilian zero coupon bond yields.
This figure contains the time series of Brazilian (annualized) monthly zero coupon bond yields with maturities of 1 month, 12 months, and 36 months between January 1999 and March 2009.

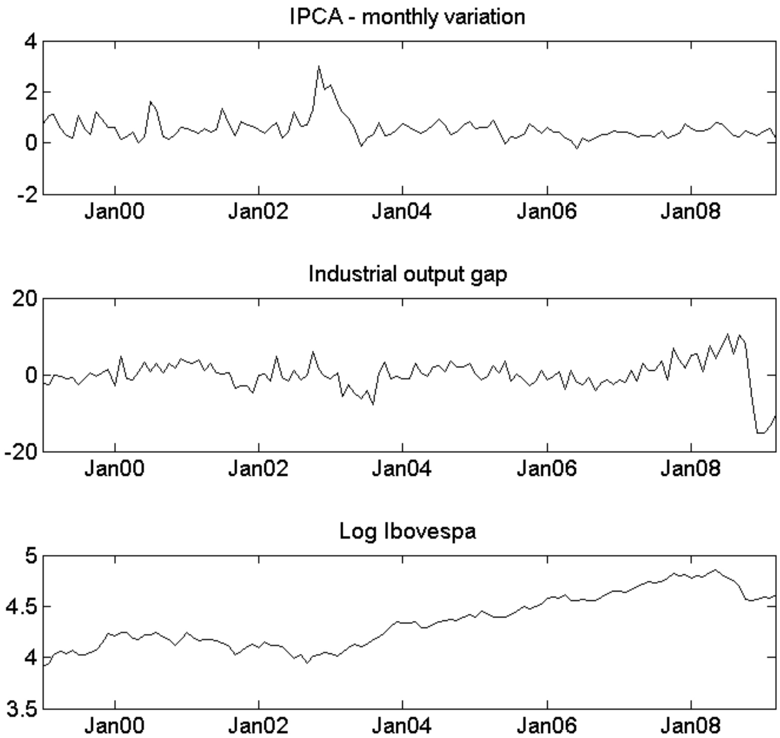

Figure 4: Brazilian observable factors.

This figure contains the time series of Brazilian observable factors between January 1999 and March 2009. The top panel shows the evolution of the monthly variation of the IPCA; the central panel depicts the evolution of the industrial output gap; and the bottom panel shows the log of the Ibovespa.

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