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## Yield curve factors, term structure volatility, and bond risk premia

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# Yield Curve Factors, Term Structure Volatility, and Bond Risk Premia

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# Yield Curve Factors, Term Structure Volatility, and Bond Risk Premia \*

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## Abstract

We introduce a Nelson-Siegel type interest rate term structure model with the underlying yield factors following autoregressive processes revealing time-varying stochastic volatility. The factor volatilities capture risk inherent to the term structure and are associated with the time-varying uncertainty of the yield curve's level, slope and curvature. Estimating the model based on U.S. government bond yields applying Markov chain Monte Carlo techniques we find that the yield factors and factor volatilities follow highly persistent processes. Using the extracted factors to explain one-year-ahead bond excess returns we observe that the slope and curvature yield factors contain the same explanatory power as the return-forecasting factor recently proposed by Cochrane and Piazzesi (2005). Moreover, we identify

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slope and curvature risk as important additional determinants of future excess returns. Finally, we illustrate that the yield and volatility factors are closely connected to variables reflecting macroeconomic activity, inflation, monetary policy and employment growth. It is shown that the extracted yield curve components have long-term prediction power for macroeconomic fundamentals.

*Key words:* Term Structure Modelling; Yield Curve Risk; Stochastic Volatility; Factor Models; Macroeconomic Fundamentals

*JEL classification:* C5, E4, G1

## 1 Introduction

Much research in financial economics has been devoted to the modelling and forecasting of interest rates and the term structure thereof. Popular theoretical approaches to term structure modelling are equilibrium models as proposed by Vasicek (1977), Cox, Ingersoll, and Ross (1985), Duffie and Kan (1996), Dai and Singleton (2002) or Duffee (2002) and no-arbitrage models in the line of Hull and White (1990) or Heath, Jarrow, and Morton (1992). Though these approaches are successful in capturing the term structure over the cross-section of maturities, they typically reveal a rather limited performance when they are used to dynamically predict future interest rates. Diebold and Li (2006) employ the Nelson and Siegel (1987) exponential components framework to dynamically model the yield curve in terms of three underlying factors associated with its level, slope, and curvature. They document that it provides significantly better out-of-sample forecasts than equilibrium factor models and competing time series approaches based on forward rates, yield levels or yield changes. Diebold, Rudebusch, and Aruoba (2006) illustrate causalities between the Nelson-Siegel factors and macroeconomic fundamentals which can be exploited for predictions of the latter. Alternatively, Cochrane and Piazzesi (2005) use a single "tent-shaped" linear combination of forward rates to predict one-year bond excess returns. They show that the so-called return forecasting factor significantly outperforms the predictive power of the level, slope and curvature factors stemming from the first three principal components extracted from the bond return covariance matrix (see Litterman and Scheinkman (1991)). Cochrane and Piazzesi argue that the return-forecasting factor contains information which is not easily captured by level, slope and curvature term structure factors.

Building on this literature, the objective of this paper is twofold: Firstly, we aim to empirically close the gap between the factor based term structure models proposed by Diebold and Li (2006) and Cochrane and Piazzesi (2005) and to study the relationships

between Nelson-Siegel yield curve factors and the Cochrane-Piazzesi return-forecasting factor. Secondly, we extend the dynamic Nelson-Siegel framework in order to capture time-varying interest rate risk. We study the importance of the latter for explaining future excess returns and link it to macroeconomic fundamentals. Specifically, we address the following three research questions: (i) Do Nelson-Siegel yield factors and the Cochrane-Piazzesi factor have the same explanatory power for future bond excess returns or do they capture different pieces of yield curve information? (ii) To which extent reveal the yield curve factors time-varying volatility and gives the latter rise to risk premia in future bond excess returns? (iii) How are the factor volatilities linked to macroeconomic fundamentals and what are their short-run and long-run relations?

Stochastic volatility in the Nelson-Siegel factors capture uncertainty in the yield curve shape. Empirical evidence suggests that changes in bond return premia over time are related to time-varying changes in the riskiness of bonds. Using GARCH-in-Mean models this is confirmed by Engle, Ng, and Rothschild (1990) and Engle and Ng (1993) showing that the return premium of a Treasury bill is driven by the time-varying risk premium of an equally weighted bill (market) portfolio which depends on its own conditional excess return variance. Whereas Engle and Ng (1993) analyze the influences of a changing bill market volatility on the shape of the yield curve, we focus on the importance of these effects for the predictability of bond return premia. This task is performed by extending the dynamic version of the Nelson and Siegel (1987) model proposed by Diebold and Li (2006) to allow the level, slope and curvature factors revealing time-varying volatility. Expressing the model in a state space form, we assume that the (unobservable) yield factors are driven by autoregressive processes with stochastic volatility. The stochastic factor volatilities reflect the bond market volatility in terms of the risk inherent to the shape of the yield curve and constitute a parsimonious alternative to a high-dimensional GARCH or SV model for the yields themselves. Hence, while Engle and Ng (1993) approximate the bill market volatility in terms of the conditional variance of a bill market portfolio's excess return, we extract volatility components from the yield curve itself. This allows us to link the approach by Engle and Ng (1993) with those by Diebold and Li (2006) and Cochrane and Piazzesi (2005).

Following Cochrane and Piazzesi (2005) we use monthly unsmoothed Fama-Bliss zero yields covering a period from January 1964 through December 2003 with maturities of up to five years. This allows us to directly compare our findings with those of Cochrane and Piazzesi (2005). The yield factors and factor volatilities are extracted from the data using Markov chain Monte Carlo (MCMC) methods and are used in

rolling window regressions of one-year-ahead excess returns.

Based on our empirical study, we can summarize the following main findings: (i) The slope and curvature yield factors describe the time-variation in future one-year-ahead bond excess returns with an  $R^2$  of about 36 percent clearly rejecting the expectations hypothesis of the term structure of interest rates. Hence, whereas Cochrane and Piazzesi (2005) conclude that their "tent-shaped" return forecasting factor has more prediction power than the level, slope and curvature factors stemming from principal components based on the yield covariance matrix, we find that this conclusion does not hold when the corresponding factors are extracted from a dynamic Nelson-Siegel model. Rather, we find that particularly the slope and curvature factor reveal a similar explanatory power as the Cochrane-Piazzesi return forecasting factor. (ii) We find strong evidence for persistent stochastic volatility dynamics in the Nelson-Siegel factors. It turns out that risks inherent to the shape of the yield curve as represented by the extracted slope and curvature volatility have explanatory power for future yearly bond excess returns beyond Cochrane and Piazzesi's return-forecasting factor. In particular, including the volatility factors in rolling window regressions increases the (adjusted)  $R^2$  from 36 percent to up to 50 percent. (iii) Our results provide evidence that the factor volatilities' explanatory power for future excess returns arises because of two effects. Firstly, it stems from a risk premium due to the uncertainty in the yield curvature. Secondly, we observe a converse effect arising from a negative relation between the slope volatility and expected excess returns. (iv) It turns out that both yield factors and factor volatilities are closely linked to macroeconomic fundamentals, such as capacity utilization, industrial production, inflation, employment growth as well as the federal funds rate. Prediction error variance decompositions show evidence for significant long-run effects of macroeconomic variables on term structure movements and volatilities thereof. Converse relations are found as well and reveal a particular importance of the curvature volatility. (v) Finally, we observe that the factor volatilities are clearly increased during economic recession periods.

The remainder of the paper is structured as follows. In Section 2, we describe the dynamic Nelson and Siegel (1987) model as put forward by Diebold and Li (2006) and Diebold, Rudebusch, and Aruoba (2006) and discuss the proposed extension allowing for stochastic volatility processes in the yield factors. Section 3 presents the data and illustrates the estimation of the model using MCMC techniques. Empirical results from regressions of one-year excess bond returns on the extracted yield factors are shown in Section 4. Section 5 gives the corresponding results when factor volatilities are used as regressors. In Section 6, the dynamic interdependencies between yield factors, factor

volatilities and macroeconomic fundamentals are investigated. Finally, Section 7 gives the conclusions.

## 2 A Dynamic Nelson-Siegel Model with Stochastic Volatility

Let  $p_t^{(n)}$  denote the log price of an  $n$ -year zero-coupon bond at time  $t$  with  $t = 1, \dots, T$  denoting monthly periods and  $n = 1, \dots, N$  denoting the maturities. Then, the yearly log yield of an  $n$ -year bond is given by  $y_t^{(n)} := -\frac{1}{n}p_t^{(n)}$ . The one-year forward rate at time  $t$  for loans between time  $t+12(n-1)$  and  $t+12n$  is given by  $f_t^{(n)} := p_t^{(n-1)} - p_t^{(n)} = ny_t^{(n)} - (n-1)y_t^{(n-1)}$ . In the following we focus on one-year returns observed on a monthly basis. Then, the log holding-period return from buying an  $n$ -year bond at time  $t-12$  and selling it as an  $(n-1)$ -year bond at time  $t$  is defined by  $r_t^{(n)} := p_t^{(n-1)} - p_{t-12}^{(n)}$ . Finally, we define excess log returns by  $z_t^{(n)} := r_t^{(n)} - y_{t-12}^{(1)}$ .

Nelson and Siegel (1987) propose modeling the forward rate curve in terms of a constant plus a Laguerre polynomial function as given by

$$f_t^{(n)} = \beta_{1t} + \beta_{2t}e^{-\lambda_t n} + \beta_{3t}\lambda_t e^{-\lambda_t n}.$$

Small (large) values of  $\lambda_t$  produce slow (fast) decays and better fit the curve at long (short) maturities. Though the Nelson-Siegel model is neither an equilibrium model nor a no-arbitrage model it can be heuristically motivated by the expectations hypothesis of interest rates. As stressed by Nelson and Siegel (1987), Laguerre polynomials belong to a class of functions which are associated with solutions to differential equations. In this context, forward rates can be interpreted as solutions to a differential equation underlying the spot rate.

The corresponding yield curve is given by

$$y_t^{(n)} = \beta_{1t} + \beta_{2t} \left( \frac{1 - e^{-\lambda_t n}}{\lambda_t n} \right) + \beta_{3t} \left( \frac{1 - e^{-\lambda_t n}}{\lambda_t n} - e^{-\lambda_t n} \right).$$

Diebold and Li (2006) interpret the parameters  $\beta_{1t}$ ,  $\beta_{2t}$  and  $\beta_{3t}$  as three latent dynamic factors with loadings 1,  $(1 - e^{-\lambda_t n})/\lambda_t n$ , and  $\{(1 - e^{-\lambda_t n})/\lambda_t n\} - e^{-\lambda_t n}$ , respectively. Then,  $\beta_{1t}$  represents a long-term factor whose loading is constant for all maturities. In contrast, the loading of  $\beta_{2t}$  starts at one and decays monotonically and quickly to zero. Consequently,  $\beta_{2t}$  may be viewed as a short-term factor. Finally,  $\beta_{3t}$  is a medium-term factor with a loading starting at zero, increasing and decaying to zero in the limit. Since it can be shown that  $y_t^\infty = \beta_{1t}$ ,  $y_t^\infty - y_t^0 = -\beta_{2t}$ , and  $y_t^0 = \beta_{1t} + \beta_{2t}$  it is

naturally to associate the long-term factor  $\beta_{1t}$  with the level of the yield curve, whereas  $\beta_{2t}$  and  $\beta_{3t}$  capture its slope and curvature, respectively. Figure 1<sup>1</sup> illustrates plots of the Nelson-Siegel factor loadings with fixed  $\lambda = 0.045$  stemming from our estimation results below.

Denoting the yield factors in the sequel by  $L_t := \beta_{1t}$ ,  $S_t := \beta_{2t}$  and  $C_t := \beta_{3t}$ , we can represent the model in state-space form

$$y_t = A f_t + \varepsilon_t, \quad (1)$$

where  $f_t := (L_t, S_t, C_t)'$  denotes the  $(3 \times 1)$  vector of latent factors,  $y_t := (y_t^{(1)}, y_t^{(2)}, \dots, y_t^{(N)})'$  is the  $(N \times 1)$  vector of yields and

$$A := \begin{pmatrix} 1 & \frac{1-e^{-\lambda \cdot 1}}{\lambda \cdot 1} & \frac{1-e^{-\lambda \cdot 1}}{\lambda \cdot 1} - e^{-\lambda \cdot 1} \\ 1 & \frac{1-e^{-\lambda \cdot 2}}{\lambda \cdot 2} & \frac{1-e^{-\lambda \cdot 2}}{\lambda \cdot 2} - e^{-\lambda \cdot 1} \\ \vdots & \vdots & \vdots \\ 1 & \frac{1-e^{-\lambda \cdot N}}{\lambda \cdot N} & \frac{1-e^{-\lambda \cdot N}}{\lambda \cdot N} - e^{-\lambda \cdot N} \end{pmatrix}$$

represents the  $(N \times 3)$  matrix of factor loadings. Finally, for the  $(N \times 1)$  vector of error terms  $\varepsilon_t$  we assume

$$\varepsilon_t := (\varepsilon_t^{(1)}, \varepsilon_t^{(2)}, \dots, \varepsilon_t^{(N)}) \sim \text{i.i.d. } N(0, \Sigma)$$

with

$$\Sigma = \text{diag} \left\{ (\sigma^{(1)})^2, (\sigma^{(2)})^2, \dots, (\sigma^{(N)})^2 \right\}. \quad (2)$$

Note that we assume the decaying factor  $\lambda_t = \lambda$  to be constant over time. This is in accordance with Diebold and Li (2006) and the common finding that time variations in  $\lambda_t$  have only a negligible impact on the model's fit and prediction power.<sup>2</sup>

Following Diebold and Li (2006), the latent dynamic yield factors are assumed to follow a first order vector autoregressive (VAR) process,

$$f_t = \mu + \Phi f_{t-1} + \eta_t, \quad (3)$$

where  $\Phi$  is a  $(3 \times 3)$  parameter matrix,  $\mu$  denotes a  $(3 \times 1)$  parameter vector, and the  $(3 \times 1)$  vector  $\eta_t$  is assumed to be independent from  $\varepsilon_t$  with

$$\eta_t \sim \text{i.i.d. } N(0, H_t). \quad (4)$$

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<sup>1</sup>All figures and tables are shown in the appendix.

<sup>2</sup>This is also confirmed by own analyzes. Actually, we also allowed  $\lambda_t$  to be time-varying but indeed found that this extra flexibility is not important for the model's goodness-of-fit.



Diebold and Li (2006) assume the conditional variances to be constant over time, i.e.  $H_t = H$ . This enables a straightforward two-step estimation of the model. In particular, fixing  $\lambda$  to a predetermined value, the latent factors  $L_t$ ,  $S_t$ , and  $C_t$  can be estimated period-by-period using ordinary least squares. Then, in a second step, the estimated factors can be used in a VAR model as represented by (3).

Given the objective of our study, we aim to extend the model to capture time-varying interest rate risk. This allows us to link the factor based approaches by Diebold and Li (2006), Diebold, Rudebusch, and Aruoba (2006), and Cochrane and Piazzesi (2005) to the work by Engle, Ng, and Rothschild (1990) and Engle and Ng (1993) relating bond market risk to bond return premia. Therefore, we propose specifying the covariance matrix  $H_t$  in terms of a (multivariate) SV process of the form

$$\text{vech}(\ln H_t) = \mu_h + \Phi_h \text{vech}(\ln H_{t-1}) + \xi_t, \quad (5)$$

where  $\text{vech}(\cdot)$  denotes the vech-operator stacking the distinct elements of the covariance matrix,  $\mu_h$  is a  $(6 \times 1)$  dimensional parameter vector and  $\Phi_h$  is a  $(6 \times 6)$  dimensional parameter matrix. The error terms  $\xi_t$  are assumed to be independent from  $\eta_t$  and  $\varepsilon_t$  and are normally distributed with covariance matrix  $\Sigma_h$  capturing the "covariance of covariance",

$$\xi_t \sim \text{i.i.d. } N(0, \Sigma_h). \quad (6)$$

However, fully parameterizing the matrices  $\Phi$ ,  $H_t$  and  $\Phi_h$  leads to a complicate model which is difficult to estimate and is typically over-parameterized in order to parsimoniously capture interest rate dynamics and associated risks. Hence, to overcome the computational burden and curse of dimensionality, we propose restricting the model to a diagonal specification with

$$\Phi = \text{diag}(\phi^L, \phi^S, \phi^C), \quad (7)$$

$$H_t = \text{diag}(h_t^L, h_t^S, h_t^C), \quad (8)$$

$$\Phi_h = \text{diag}(\phi_h^L, \phi_h^S, \phi_h^C), \quad (9)$$

where  $\text{diag}(\cdot)$  captures the diagonal elements of a (symmetric) matrix in a corresponding vector. As shown in the empirical analysis below, these restrictions are well supported

by the data.<sup>3</sup> Then, the latent factor structure can be expressed by

$$\begin{pmatrix} L_t \\ S_t \\ C_t \end{pmatrix} = \begin{pmatrix} \mu^L \\ \mu^S \\ \mu^C \end{pmatrix} + \begin{pmatrix} \phi^L & 0 & 0 \\ 0 & \phi^S & 0 \\ 0 & 0 & \phi^C \end{pmatrix} \begin{pmatrix} L_{t-1} \\ S_{t-1} \\ C_{t-1} \end{pmatrix} + \eta_t, \quad (10)$$

where  $\eta_t \sim \text{i.i.d. } N(0, H_t)$  with

$$\text{diag}(\ln H_t) = \begin{pmatrix} \ln(h_t^L) \\ \ln(h_t^S) \\ \ln(h_t^C) \end{pmatrix} = \begin{pmatrix} \mu_h^L \\ \mu_h^S \\ \mu_h^C \end{pmatrix} + \begin{pmatrix} \phi_h^L & 0 & 0 \\ 0 & \phi_h^S & 0 \\ 0 & 0 & \phi_h^C \end{pmatrix} \begin{pmatrix} \ln(h_{t-1}^L) \\ \ln(h_{t-1}^S) \\ \ln(h_{t-1}^C) \end{pmatrix} + \begin{pmatrix} \xi_t^L \\ \xi_t^S \\ \xi_t^C \end{pmatrix}. \quad (11)$$

We refer  $h_t^L$ ,  $h_t^S$  and  $h_t^C$  to as so-called "factor volatilities" capturing the time-varying uncertainty in the level, slope and curvature of the yield curve. The level volatility  $h_t^L$  corresponds to a component which is common to the time-varying variances of all yields. It is associated with underlying latent (e.g. macroeconomic) information driving the uncertainty in the overall level of interest rates. Indeed, it can be seen as a model implied proxy of the bond market volatility which is captured by Engle, Ng, and Rothschild (1990) in terms of the conditional excess return variance of an equally weighted bill market portfolio. Correspondingly,  $h_t^S$  is associated with risk inherent to the slope of the yield curve. It reflects the riskiness in yield spreads, and thus time-variations in the risk premium which investors require to hold long bonds instead of short bonds. Finally,  $h_t^C$  captures uncertainties associated with the curvature of the yield curve, which can vary between convex, linear and concave forms. Obviously, such variations mainly stem from time-varying volatility in bonds with mid-term maturities.

An alternative way to capture time-varying volatility in the term structure of interest rates would be to allow  $\Sigma$  itself to be time-varying. However, this would result in  $N$ -dimensional MGARCH or SV models which are not very tractable if the cross-sectional dimension  $N$  is high. Therefore, we see our approach as a parsimonious alternative to capture interest rate risk. Note that the slope and curvature factors can be interpreted as particular (linear) combinations of yields associated with factor portfolios mimicking the steepness and convexity of the yield curve.<sup>4</sup> Then, the corresponding slope and curvature volatilities are associated with the volatilities of

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<sup>3</sup>Note that we also estimated models with non-zero off-diagonal elements in  $\Phi$  and in a time-invariant matrix  $H$  and found that most off-diagonal parameters are indeed statistically insignificant.

<sup>4</sup>This interpretation is also reflected in the linear combinations of yields which are typically used to empirically approximate the underlying yield curve factors. In particular, level, slope and curvature are often approximated by  $\frac{1}{3}(y_t^{(1)} + y_t^{(3)} + y_t^{(5)})$ ,  $y_t^{(5)} - y_t^{(1)}$ , and  $2y_t^{(3)} - y_t^{(5)} - y_t^{(1)}$ . See also Section 3.3.

the underlying factor portfolios. In this sense, they capture time-variations in yields' variances and covariances driving the yield curve shape.

Using this structure and the imposed normality assumptions, the unconditional moments of the yields are straightforwardly given by  $E[y_t] = A E[f_t]$  and  $\text{Var}[y_t] = A \text{Var}[f_t] A$ , where the moments of the  $i$ -th element are given by

$$E[f_t^{(i)}] = \mu^i (1 - \phi^i)^{-1}, \quad (12)$$

$$\text{Var}[f_t^{(i)}] = \frac{1}{1 - \phi^{i2}} \left[ \frac{\mu_h^i}{1 - \phi_h^i} + \frac{(\sigma_h^i)^2}{2(1 - \phi_h^{i2})} \right], \quad (13)$$

$$\text{Corr}[f_t^{(i)}, f_{t-k}^{(i)}] = \phi^{ik}, \quad k > 0, \quad (14)$$

where  $i \in \{L, S, C\}$ .

Accordingly, the correlation structure in higher-order moments of de-measured yield factors corresponds to that of a basic SV model and can be approximated by

$$\text{Corr}[a_t^{ip}, a_{t-k}^{ip}] \approx C(p, (\sigma_h^i)^2) \phi_h^{ik}, \quad k > 0, \quad (15)$$

where  $a_t^i := \ln |\eta_t^i| = \ln |f_t^{(i)} - \mu - \phi^i f_{t-1}^{(i)}|$  and

$$C(p, (\sigma_h^i)^2) = \frac{\mathcal{A}(p, (\sigma_h^i)^2) - 1}{\mathcal{A}(p, (\sigma_h^i)^2) \mathcal{B}(p) - 1}, \quad (16)$$

$$\mathcal{B}(p) = \sqrt{\pi} \Gamma\left(p + \frac{1}{2}\right) \Gamma\left(\frac{p}{2} + \frac{1}{2}\right)^{-2}, \quad (17)$$

$$\mathcal{A}(p, (\sigma_h^i)^2) = \exp\left(p^2 \frac{(\sigma_h^i)^2}{1 - \phi_h^{i2}}\right). \quad (18)$$

### 3 Extracting Yield Curve Factors and Factor Volatilities

#### 3.1 Data

In order to make our results comparable to those by Cochrane and Piazzesi (2005), we use the same set of monthly unsmoothed Fama-Bliss zero-coupon yields covering a period from January 1964 through December 2003 with maturities ranging between one and five years. The data is available from the Center for Research in Security Prices (CRSP) and is constructed using the method of Fama and Bliss (1987) based on end-of-month data of U.S. taxable, non-callable bonds for annual maturities up to five years. Here, each month a term structure of one-day continuously compounded forward rates is calculated from available maturities up to one year. To extend beyond

a year, Fama and Bliss (1987) use the assumption that the daily forward rate for the interval between successive maturities is the relevant discount rate for each day in the interval. This allows to compute the term structure based on a step-function in which one-day forward rates are the same between successive maturities. Then, the resulting forward rates are aggregated to generate end-of-month term structures of yields for annual maturities up to five years. Summary statistics of the data are given in Panel A of Table 2.

### 3.2 MCMC Based Inference

The diagonal model specified above corresponds to a three-level latent hierarchical model with six latent processes.

Let  $\Theta$  denote the collection of the model parameters. Moreover, let  $F_t := (L_t, S_t, C_t)$  and  $V_t := (h_t^L, h_t^S, h_t^C)$ . Then, the likelihood function of the model is given by

$$p(\Theta|Y) = \int_{F_1} \int_{F_2} \cdots \int_{F_T} p(Y|\Theta, F_1, F_2, \dots, F_T) p(F_1, F_2, \dots, F_T|\Theta) dF_1 dF_2 \cdots dF_T,$$

where  $p(Y|\Theta, F_1, F_2, \dots, F_T)$  denotes the (conditional) density of the data  $Y$  given the parameters  $\Theta$  and the latent factors and reflects the imposed structure as given by (1) and (2). Furthermore,  $p(F_1, F_2, \dots, F_T|\Theta)$  denotes the (conditional) joint density of the latent factors, given the model parameters  $\Theta$  and is determined by (3). Since the factors are unobservable, they have to be integrated out resulting in a  $(3 \cdot T)$ -dimensional integral. Obviously,  $p(F_1, F_2, \dots, F_T|\Theta)$  depends on a further set of unknown components as represented by the volatility components  $V_1, \dots, V_T$ . It is computed as

$$\begin{aligned} p(F_1, F_2, \dots, F_T|\Theta) &= \int_{V_1} \int_{V_2} \cdots \int_{V_T} p(F_1, F_2, \dots, F_T|\Theta, V_1, V_2, \dots, V_T) \\ &\quad \times p(V_1, V_2, \dots, V_T|\Theta) dV_1 dV_2 \cdots dV_T, \end{aligned}$$

where  $p(V_1, V_2, \dots, V_T|\Theta)$  denotes the joint density of the volatility factors as determined by (5). This likelihood function cannot be computed analytically in closed form and requires numerical approximation techniques. We propose estimating the model using Markov chain Monte Carlo (MCMC) based inference. Consequently, we consider  $\Omega := \{\Theta, F_1, \dots, F_T, V_1, \dots, V_T\}$  to be a random vector whose posterior distribution

$p(\Omega|Y)$  can be arranged according to

$$\begin{aligned}
p(\Omega|Y) &= p(F_1, F_2, \dots, F_T, V_1, V_2, \dots, V_T, \Theta|Y) \\
&\propto p(Y|F_1, F_2, \dots, F_T, V_1, V_2, \dots, V_T, \Theta) \\
&\times p(F_1, F_2, \dots, F_T|V_1, V_2, \dots, V_T, \Theta) \\
&\times p(V_1, V_2, \dots, V_T|\Theta) \\
&\times p(\Theta).
\end{aligned} \tag{19}$$

By specifying the prior distributions  $p(\Theta)$ <sup>5</sup>, we utilize Gibbs and Metropolis-Hastings samplers to simulate the posterior distribution,  $p(\Omega|Y)$ . Then, both parameter and factor estimates are obtained by taking the sample averages of the corresponding MCMC samples. The estimated latent yield factors and volatility factors will be used in the following subsections as predictors for future excess returns. In order to make our results comparable to those by Cochrane and Piazzesi (2005), we use the estimates of smoothed factors  $E[f_t|y^T]$  with  $y^t = (y_1, \dots, y_T)$  denoting all observations up to time  $T$ .<sup>6</sup>

### 3.3 MCMC Estimation Results

We start our analysis by estimating the model with constant volatility factors corresponding to the specification proposed by Diebold and Li (2006).<sup>7</sup> The estimation results are given in Panel A of Table 1. The dynamics of  $L_t$ ,  $S_t$  and  $C_t$  are very persistent with estimated autoregressive coefficients of 0.98, 0.96 and 0.91, respectively. In particular, whereas the level of interest rates close to a unit root, the persistence of the spread component is lower but still relatively high. This finding is in strong accordance with the literature.

The model implied unconditional mean of the level factor, given by  $\mu^L/(1 - \phi^L)$ , equals 7.96 which is close to its empirical mean of 7.12. Correspondingly, the mean value of the slope factor equals  $-1.96$  reflecting that during the sample period the yield curve has been on average upward sloped.<sup>8</sup> Finally, the mean of the curvature

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<sup>5</sup>For specific details on the choice of the prior distributions, see Appendix A.

<sup>6</sup>Using instead filtered factors  $E[f_t|y^{t-1}]$  would allow us to evaluate the out-of-sample prediction performance of the model in line with Diebold and Li (2006). We keep this issue for further research.

<sup>7</sup>Exploiting the linearity of the model, it could be alternatively estimated using quasi maximum likelihood based on the Kalman filter, see e.g. Harvey (1990). However, to keep our econometric approach consistent, we estimate all specifications in this paper using MCMC techniques.

<sup>8</sup>Recall that we define the slope as the difference between short yields and long yields.

factor is  $-0.28$  but not significantly different from zero. Hence, on average we do not observe a strong curvature in the yield curve. The estimated decay factor  $\lambda$  equals  $0.055$ , implying the curvature loading  $(1 - \exp(-\lambda n))/(\lambda n - \exp(-\lambda n))$  to be maximized for a maturity of  $2.72$  years. The last column in Table 1 reports the Geweke (1992) Z-scores which are used to test the convergence of the Markov chains generated from the MCMC algorithm.<sup>9</sup> It turns out that all Markov chains have been properly converged. The descriptive statistics shown in Panel B of Table 2 indicate that the dynamic Nelson-Siegel model captures a substantial part of the dynamics in the yields confirming the findings by Diebold and Li (2006). Nevertheless, it turns out that there are still autocorrelations in the residuals as well as squared residuals indicating that there are neglected dynamics in the first and second moments of the process.

Figure 2 plots the resulting estimated Nelson-Siegel factors and their corresponding empirical approximations. We observe that the estimated slope factor is basically perfectly correlated with its empirical counterpart yielding a correlation of  $-0.99$ . The corresponding correlations for the level and curvature factors are  $0.90$  and  $0.59$ . This indicates that level and slope factors can be easily approximated by their corresponding empirical counterparts whereas approximations of the curvature factor tend to be rather difficult.

Panel B of Table 1 shows the parameter estimates of the model with stochastic volatility components, given by equations (1) - (3) and (5). The estimated decay parameter equals  $0.045$  implying that the curvature loading is maximized at a maturity of  $3.33$  years. The estimates of the yield factor parameters are close to those of the constant volatility model. The estimated dynamic parameters in the volatility components are  $0.977$ ,  $0.964$  and  $0.933$  for the level, slope and curvature volatilities, respectively. Hence, as for the yield curve factors we also find a high persistence in the stochastic volatility processes. This is particularly true for the level and slope volatility.

## 4 Explaining Bond Excess Returns Using Yield Factors

### 4.1 Nelson-Siegel Factors as Predictors

In this section, we examine the explanatory power of the extracted Nelson-Siegel factors for future bond excess returns. Following Cochrane and Piazzesi (2005), we regress monthly one-year-ahead bond excess returns with maturities of two up to five years on

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<sup>9</sup>For details, see Appendix A.

the estimated level, slope and curvature factors, i.e.,

$$z_t^{(n)} = c + \beta_L L_{t-12} + \beta_S S_{t-12} + \beta_C C_{t-12} + \varepsilon_t^{(n)}, \quad n = 2, 3, 4, 5. \quad (20)$$

Panel A of Table 3 reports the estimation results based on alternative regressions. Two caveats should be taken into account. Firstly, because of the overlapping windows, the errors  $\varepsilon_t^{(n)}$  are per construction strongly autocorrelated. This requires using estimators of the parameter covariances which are robust against serial correlation (and potential heteroscedasticity). In accordance with Cochrane and Piazzesi (2005) we apply the classical heteroscedasticity and autocorrelation consistent (HAC) estimators proposed by Hansen and Hodrick (1980) given by

$$Cov[\hat{b}] = E[x_t x_t']^{-1} \left[ \sum_{j=-k}^k E[x_t x_{t-j}' \varepsilon_{t+1} \varepsilon_{t+1-j}] \right] E[x_t x_t']^{-1} \quad (21)$$

and the (Bartlett) kernel estimator proposed by Newey and West (1987) given by

$$Cov[\hat{b}] = E[x_t x_t']^{-1} \left[ \sum_{j=-k}^k \frac{k-|j|}{k} E[x_t x_{t-j}' \varepsilon_{t+1} \varepsilon_{t+1-j}] \right] E[x_t x_t']^{-1}, \quad (22)$$

where  $x_t$  denotes the vector of regressors and  $j$  denotes the order of lag truncation.

Secondly, high persistence in the yield factors used as regressors might cause spurious effects affecting the  $R^2$ . Accordingly, we evaluate the goodness-of-fit based on  $R^2$ -values only in combination with Newey-West and Hansen-Hodrick adjusted tests for joint significance. Moreover, it turns out that information criteria such as the Bayes Information Criterion confirm the evaluation results based on  $R^2$  numbers and joint significance test.<sup>10</sup> Finally, as shown below, we observe that the predictability arises typically from those factors which reveal the lowest persistence. This confirms the robustness of our results.

It is shown that the level factor is virtually insignificant and has no explanatory power for future bond excess returns. Neglecting the latter in the regression reduces the  $R^2$  values<sup>11</sup> and Hansen Hodrick HAC  $\chi^2$ -statistics for joint significance only slightly. This result is mostly true for maturities longer than two years. These results consistent with Duffee (2002) showing based on the essentially affine term structure model that the level factor is irrelevant for bond excess returns. Though the Nelson-Siegel framework is different from Duffee's model, the extracted yield factors behave in a quite similar way.

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<sup>10</sup>For sake of brevity, these measures are not shown in the paper.

<sup>11</sup>Throughout the paper, the  $R^2$  refers to the coefficient of determination, adjusted by the number of regressors.

As stressed by Diebold, Piazzesi, and Rudebusch (2005) the loadings in (1) are quite close to those estimated from the three factor essentially affine model. In contrast, the coefficients for the slope and curvature factor are highly significant. We find that future excess returns decrease with the slope (defined as short minus long) and increase with the curvature. This result is consistent with, for instance, Fama and Bliss (1987) and Campbell and Shiller (1991). The positive coefficient for the curvature factor indicates that future excess returns are expected to be higher the more hump-shaped, i.e. convex or concave the current yield curve. Hence, a major factor driving future excess returns is the yield spread between mid-term and short-term bonds.

Including all yield factors leads to a  $R^2$  of up to 36 percent, revealing basically the same explanatory power as in Cochrane and Piazzesi (2005) using their “tent-shaped” return-forecasting factor. The corresponding  $\chi^2$ -values are clearly well above the five percent critical value indicating that Nelson-Siegel factors jointly do contain significant information for future bond excess returns. Obviously, the explanatory power arises mainly from the slope and curvature factors which are statistically significant for all individual bonds. Omitting both factors from the regressions clearly reduces the  $R^2$  and  $\chi^2$ -values. This is particularly true for longer maturities and the curvature factor which turns out to be most important for the prediction of future excess returns.

## 4.2 The Cochrane-Piazzesi Return-Forecasting Factor

Cochrane and Piazzesi (2005) propose forecasting bond excess returns with the so called return-forecasting factor,  $\vartheta_t$ , defined as a linear combination

$$\vartheta_t = \gamma' f_t \quad (23)$$

of five forward rates  $f_t = (1, y_t^{(1)}, f_t^{(2)}, \dots, f_t^{(5)})$  with weights  $\gamma = (\gamma^{(0)}, \gamma^{(1)}, \dots, \gamma^{(5)})$ . The weights are estimated by running a (restricted) regression of average (across maturity) excess returns on the forward rates,

$$\frac{1}{4} \sum_{n=2}^5 z_t^{(n)} = \gamma^{(0)} + \gamma^{(1)} y_{t-12}^{(1)} + \gamma^{(2)} f_{t-12}^{(2)} + \dots + \gamma^{(5)} f_{t-12}^{(5)} + u_t. \quad (24)$$

Then, the return forecasting regression for individual bond excess returns is given by

$$z_t^{(n)} = b^{(n)} \vartheta_{t-12} + \varepsilon_t^{(n)}, \quad n = 2, 3, 4, 5, \quad (25)$$

with regression coefficients  $b^{(n)}$  and the restriction  $\frac{1}{4} \sum_{n=2}^5 b^{(n)} = 1$ .

Cochrane and Piazzesi (2005) show that  $\vartheta_t$  contain information for future excess returns which are not captured by yield factors as represented by the first three principal components of the covariance matrix of yields. As suggested by Litterman and



Scheinkman (1991), the latter serve as empirical proxies for the level, slope and curvature movements of the term structure. Panel A in Table 4 reports the  $R^2$  values and Hansen-Hodrick HAC  $\chi^2$ -statistics based on regressions where  $z_t^n$  is regressed on (i) the principal components (PC's), (ii) the return-forecasting factor,  $\vartheta_t$ , and (iii) the Nelson-Siegel yield curve factors. It turns out that both the return-forecasting factor and the Nelson-Siegel factors have effectively the same explanatory power with an  $R^2 \approx 0.37$  implied by  $\vartheta_t$  and an  $R^2 \approx 0.36$  implied by the Nelson-Siegel factors. This results is widely confirmed by the  $\chi^2$ -statistics which are similar for longer maturities. Running regressions of the return-forecasting factor on the estimated Nelson-Siegel factors yields<sup>12</sup>:

$$\begin{aligned} \vartheta_t &= 0.025 \cdot L_t - 0.629 \cdot S_t + 0.901 \cdot C_t, & R^2 &= 0.704, \\ & \quad (0.819) \quad (-6.968) \quad (9.359) & & [186.4] \\ \vartheta_t &= -0.676 \cdot S_t + 0.902 \cdot C_t, & R^2 &= 0.699. \\ & \quad (-9.235) \quad (8.333) & & [137.8] \end{aligned}$$

Hence, we observe that around 70 percent variation in the return-forecasting factor is explained by the slope and the curvature factor, whereas the level factor  $L_t$  is statistically insignificant. This result illustrates that the return-forecasting factor and the Nelson-Siegel slope and curvature factors are closely related and capture similar information for future excess returns.

However, the close correspondence between the return-forecasting factor and yield curve factors does not hold if the latter are constructed from principal components of the covariance matrix. Principal component factors reveal a significantly lower explanatory power with an  $R^2$  of approximately 0.25 and clearly reduced  $\chi^2$ -statistics. Figure 3 shows the Nelson-Siegel curvature loading  $((1 - e^{-\lambda_t n})/\lambda_t n) - e^{-\lambda_t n}$ , the return-forecasting factor loading  $\gamma^{(n)}$ , and the loading of the third PC factor. We observe that both the return-forecasting factor and the Nelson-Siegel curvature factor are curved at the long end of the yield curve, whereas the PC curvature is curved only at the short end. Cochrane and Piazzesi (2005) argue that in order to capture relevant information about future bond excess returns contained in the four-year to five-year yield spread, the factor loading should be curved at long end.<sup>13</sup> This might explain why the Nelson-Siegel yield factors outperform the PC yield factors and why the former have similar forecasting power as the return-forecasting factor. It also stresses the importance of the curvature factor.

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<sup>12</sup>In parentheses we show the robust t-statistics, where the brackets contain the Hansen-Hodrick HAC  $\chi^2$ -test statistics on joint significance.

<sup>13</sup>Note that the return-forecasting factor is only 'tent-shaped' when it is estimated from forward rates. If it is estimated from yields, it is curved at long end.

In order to test the joint performance of both the Nelson-Siegel yield factors as well as the Cochrane-Piazzesi forecasting factor we run the regression

$$z_t^{(n)} = \beta_L L_{t-12} + \beta_S S_{t-12} + \beta_C C_{t-12} + \varphi \vartheta_t + \varepsilon_t^{(n)}. \quad (26)$$

Panel B of Table 4 gives the corresponding estimation results. We observe that the joint inclusion of Nelson-Siegel factors and the return-forecasting factor further increases the explanatory power compared to the cases where both sets of factors are included individually. This is evident by a clear increase of the  $\chi^2$ -statistics and a slight increase of the  $R^2$  values. Given the close correspondence between both types of factors, it is not surprising that the inclusion of the return-forecasting factor reduces the significance of the individual Nelson-Siegel factors. This is particularly true for the slope factor where the curvature factor still provides significant information beyond the return-forecasting factor. Hence, we can conclude that even though both types of factors individually provide similar explanatory power they partly capture different pieces of information which can be exploited for the prediction of bond excess returns. These results complement the findings by Cochrane and Piazzesi (2005). Hence, we can conclude that the Nelson-Siegel factors do not contain significant explanatory power on bond excess returns beyond the return-forecasting factor, and vice versa.

## 5 Explaining Bond Excess Returns Using Factor Volatilities

As stressed above, the extracted factor volatilities can be heuristically interpreted as the volatilities of factor portfolios representing the level, steepness and convexity of the yield curve. A crucial question is whether riskiness in the yield curve is reflected in expected bond excess returns and give rise to a risk premium. By including the factor volatilities  $h_t^L$ ,  $h_t^S$  and  $h_t^C$  we run the regression

$$z_t^{(n)} = c + \alpha_L h_{t-12}^L + \alpha_S h_{t-12}^S + \alpha_C h_{t-12}^C + \varepsilon_t^{(n)}. \quad (27)$$

As shown in Panel B of Table 3 the volatility factors do contain significant information for future bond excess returns. Including all volatility components yields an  $R^2$  of up to 18 percent with all factors being jointly significant. The main explanatory power comes from the slope and curvature factor volatility, but not from the level volatility. Most interestingly, the impact of the slope volatility on future excess returns is negative. Hence, increasing uncertainty regarding the slope of the yield curve decreases future bond return premia. In other words, if the yield curve slope turns out to be stable,

positive excess returns become more likely. This result is in contrast to the hypothesis of a positive risk premium and is rather in line with a "stability compensation". In contrast, we find that future bond excess returns increase with the curvature volatility. As discussed above, the latter reflects the time-varying uncertainty regarding the convexity or concavity of the yield curve, respectively, and is dominantly driven by the riskiness of mid-term bonds. The positive impact on future excess returns is clearly in line with the hypothesis of a risk premium. Hence, our results provide evidence that the riskiness regarding yield curve slope and yield curve convexity work in opposite directions: Investors are compensated for taking risk regarding medium-term maturities and avoiding risk regarding long-term maturities. Hence, future excess returns are expected to be highest if spreads between long-term and short-term bonds are high and stable but the yield curve convexity is uncertain.<sup>14</sup>

A crucial question is whether the factor volatilities have still explanatory power if we control for the yield curve factors themselves. The corresponding estimation results are given in Panel C of Table 3. Interestingly, it turns out that the use of both Nelson-Siegel factors and factor volatilities yields an  $R^2$  of about 50 percent. In other words, the inclusion of volatility factors in addition to yield factors shifts the  $R^2$  from 36 percent to up to 50 percent. This indicates that factor volatilities have significant explanatory power for future excess returns even when we account for yield curve factors. These results are also strongly supported by a significant increase of the  $\chi^2$ -statistics showing that this additional prediction power mainly stems from the slope and curvature volatility.

The regression results shown in Panel C of Table 4 show that the volatility factors have also explanatory power beyond the Cochrane-Piazzesi return-forecasting factor. Actually, the  $R^2$  increases from 36 percent to up to 42 percent if the volatility factors are added to the Cochrane-Piazzesi return-forecasting factor. This implies that the volatility factors do contain significant information on bond excess returns which is neither subsumed by yield curve factors nor by the return-forecasting factor.

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<sup>14</sup>A potential explanation for the predictive power of volatility components for future excess returns could be that we predict log returns instead of simple returns. However, redoing the whole analysis based on simple returns even enforces our results and indicates that our findings are not due to a predictable volatility components in the mean of log returns.

## 6 Yield Factors, Factor Volatilities, and Macroeconomic Fundamentals

In this section, we aim to analyze in which sense yield factors and factor volatilities serve as mimicking portfolios for underlying macroeconomic fundamentals. Correspondingly, we extend our analysis by five macroeconomic variables: the inflation rate (INF), measured by monthly relative changes of the consumer price index, the manufacturing capacity utilization (CU), the federal funds rate (FFR), the employment growth rate (EMP) as well as industrial production (IP). The choice of the variables is motivated by the results by Diebold, Rudebusch, and Aruoba (2006) who identify manufacturing capacity utilization, the federal funds rate as well as annual price inflation as the minimum set of important variables driving the term structure of interest rates. By augmenting the set of variables by the employment growth rate as well as industrial production we aim parsimoniously capturing the (relative) level of industrial utilization, labor market activity, inflation as well as the monetary policy as key underlying macroeconomic fundamentals characterizing the state of the economy.

To analyze the contemporaneous correlations between yield factors and macroeconomic fundamentals we regress the yield factors on the contemporaneous (monthly) macroeconomic variables, i.e.

$$F_t = \mu + \beta_1 INF_t + \beta_2 IP_t + \beta_3 FFR_t + \beta_4 EMP_t + \beta_5 CU_t + \varepsilon_t, \quad (28)$$

where  $F_t := \{L_t, S_t, C_t, h_t^L, h_t^S, h_t^C\}$ . The results reported by Table 5 show that the federal funds rate and capacity utilization are significant determinants of the level and slope factor and explain a substantial part in variations of the latter.<sup>15</sup> The positive signs for FFR and negative signs for CU signs of the coefficients are economically plausible and in line with theory. While the level and slope factor are closely connected to monetary policy and macroeconomic activity, only a small fraction of variations in the yield curve curvature can be explained by the latter. Including all macroeconomic variables yields an  $R^2$  of just 14%. This result illustrates that it is rather difficult to explain the shape of the yield curve by observable variables.

The results in Table 5 show that not only the yield curve factors themselves but

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<sup>15</sup>As above, one might argue that the  $R^2$  values should be treated with caution since some of the regressors, such as CU and IP, are relatively persistent and might cause spurious correlation effects. However, robust tests on joint significance of the regressors yield the same conclusions. Moreover, the low explanatory power of the curvature factor regression indicates that spurious effects cannot be the major reason for high  $R^2$ 's.

also their volatilities are significantly driven by underlying macroeconomic dynamics. It turns out that periods of high inflation and capacity utilization are generally accompanied by a lower volatility in interest rate levels. This might be partly explained by monetary policy interventions. Moreover, we find evidence for leverage effects in the sense of higher (lower) level and slope volatilities in periods of higher (lower) federal fund rate levels. This confirms the results by Engle, Ng, and Rothschild (1990) and Engle and Ng (1993) of (positive) GARCH-in-Mean effects.<sup>16</sup> Whereas the curvature factor is not easily explained by observable macroeconomic variables, this is not true for the corresponding volatility. Actually, we observe that particularly the federal funds rate, the employment growth rate as well as capacity utilization are significant determinants of the time-varying uncertainty in the yield curve shape yielding an  $R^2$  of about 0.48. It turns out that periods of a high federal funds rate, low capacity utilization and negative employment growth induce higher variations in medium-term bonds and thus the term structure convexity. Overall, we can summarize that factor volatilities are even closer connected to observable macroeconomic variables than the factors themselves.

Table 6 shows the estimation results of a VAR(1) model of monthly yield factors and macroeconomic variables,

$$F_t = \mu + A F_{t-12} + \varepsilon_t, \quad (29)$$

where  $F_t := \{L_t, S_t, C_t, INF_t, IP_t, FFR_t, EMP_t, CU_t\}$ . We can summarize the following results: Firstly, the yield factors primarily depend on their own lags but not on those of the other factors confirming the diagonal specification of  $\Phi$  in (7). Secondly, we observe that the yield factors are not (short-term) predictable based on macroeconomic fundamentals. This is particularly true for the level and the curvature factor whereas for the curvature factor slight dependencies from lagged inflation rates, federal fund rates, and employment growth rates are observable. Thirdly, it turns out that level and slope factors have significant short-term prediction power for nearly all macroeconomic variables. In particular, rising interest rate levels and yield spreads predict increases in industrial production, federal funds rates, the growth rate of employment as well as the capacity utilization. In contrast, the term structure curvature contains no information for one-month-ahead macroeconomic variables. Overall these results widely confirm those by Diebold, Rudebusch, and Aruoba (2006).

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<sup>16</sup>In preliminary studies we found evidence for SV-in-Mean effects for the level factor. Given the close relation between the federal funds rate and the level of interest rates this effect is now obviously reflected in the present regressions. The results are not shown here but are available upon request from the authors.

In order to study the dynamic interdependencies between macroeconomic variables and the factor volatilities we estimate (29) with  $F_t := \{h_t^L, h_t^S, h_t^C, INF_t, IP_t, FFR_t, EMP_t, CU_t\}$ . The results shown in Table 7 indicate that most of the (short-term) dynamics are driven by process-own dependencies confirming also the assumption of a diagonal structure of  $H_t$  in (8). It turns out that the level volatility is dominantly predicted by past level and slope volatilities but not based on macroeconomic variables. Similar relations are also observed for the slope volatility where the latter also significantly (positively) depends on the lagged federal funds rate. In contrast, the curvature volatility depends solely on its own history. Hence, we can conclude that in the short run term structure volatilities are not predictable based on macroeconomic factors. Conversely we observe a weak predictability of the level volatility for future macroeconomic fundamentals. In particular, higher level volatilities predict increases in industrial production, employment growth rates as well as decreasing inflation rates and manufacturing capacity utilizations. Similar effects on inflation rates and capacity utilization is observed for the slope volatility. Interestingly, the strongest impact on future macroeconomic variables stems from the curvature factor which has significant prediction power for all macroeconomic factors. This finding illustrates again the importance of term structure curvature risk confirming our results above.

Long-term relations between the individual variables are analyzed based on prediction error variance decompositions (see e.g. Hamilton (1994)) implied by the VAR estimates discussed above. The corresponding plots are shown in Figures 5 to 12. We observe that not only in the short run but also in the long run macroeconomic variables virtually do not contribute to the prediction error variances in yield curve levels and curvatures. Only for the yield curve slope, particularly capacity utilization and industrial production can explain about 25% in prediction error variances after 100 months. Conversely, we observe significantly higher long-run forecasting ability of yield term factors for macroeconomic fundamentals. This is particularly apparent for the federal funds rate whose prediction error variance is dominated by the level and slope factor (by nearly 80%). For CU, EMP and IP we observe that yield curve factors - predominantly level and slope - can explain around 40% in long-run prediction error variances. Hence, in line with Diebold, Rudebusch, and Aruoba (2006) we can conclude that level factors serve as long-run predictors of future industrial utilization, employment growth and short-term monetary policy. A notable exception is the inflation rate which is not predictable based on yield curve factors, neither over the short run nor the long run.

Figures 9 to 12 show the corresponding variance decompositions implied by the

VAR estimates of  $F_t := \{h_t^L, h_t^S, h_t^C, INF_t, IP_t, FFR_t, EMP_t, CU_t\}$ . It is evident that macroeconomic fundamentals explain a major part in long-term prediction error variances of level and slope volatilities. Particularly capacity utilization and industrial production explain approximately 50% and 40% in long-term prediction error variances of the level volatility and slope volatility, respectively. In contrast, long-term curvature volatility is significantly more difficult to predict based on macroeconomic variables. Here, we find that the latter explain less than 20% in prediction error variances. Overall, we can conclude that the production intensity is an important long-term predictor of the riskiness in interest levels and slopes but to less extent in the curvature. Vice versa, we observe a more important role of the curvature volatility for the prediction of future macroeconomic activity. This is again particularly true for the capacity utilization, employment growth and (to less extent) industrial production whose prediction error variation after 100 months is significantly influenced by the current curvature volatility. In contrast, virtually no long-run explanatory power of level and slope volatilities for future macroeconomic variables can be identified. Hence, we observe that particularly the uncertainty with respect to the shape of the yield curve has long-term consequences for production activity and employment growth.

Further insights into the role of the extracted factor volatilities can be gained by Figure 4 which plots the latter over the sample period. It turns out that the slope volatility peaks in April 1974, April 1980 and March 2001 corresponding to three major economic recession periods in the U.S. as identified by the National Bureau of Economic Research (NBER). Viewing the slope factor as a short-term factor, its high fluctuations in these periods might be attributed to monetary policy reflected in short-term yields during economic recessions. The same pattern is observed for the curvature volatility capturing mainly the uncertainty in medium-term yields and significantly peaking during all recessions periods. Hence, we observe that interest rate risk during economic recessions is dominantly reflected in the shape of the yield curve but not in the overall level.

## 7 Conclusions

We propose a dynamic Nelson-Siegel type yield curve factor model, where the factors themselves reveal stochastic volatility. By estimating the model using MCMC techniques we extract both the Nelson-Siegel factors as well as their volatility components and use them to predict bond return premia and to relate them to underlying macroeconomic variables. This approach allows us to link the approaches by Diebold and Li (2006), Diebold, Rudebusch, and Aruoba (2006), Cochrane and Piazzesi (2005) on

factor-based term structure modeling with the GARCH-in-Mean models by Engle, Ng, and Rothschild (1990) and Engle and Ng (1993) capturing interest rate risk premia.

We can summarize the following main findings: (i) We find that the slope and curvature factors extracted from the dynamic Nelson-Siegel model describe time-variations in future yearly bond excess returns with an  $R^2$  of up to 36 percent. (ii) The Nelson-Siegel yield factors have basically the same explanatory power as the return-forecasting factor proposed by Cochrane and Piazzesi (2005). This result arises mainly because of a close similarity between the loadings of the “tent-shaped” return-forecasting factor and that of the slope and curvature Nelson-Siegel factor. This forecasting performance is not achieved when principal components are used as predictors. (iii) We show that the time-varying volatility associated with the level, slope and curvature factors have significant explanatory power for future excess returns beyond the factors themselves. Including the extracted factor volatilities in rolling window regressions increases the (adjusted)  $R^2$  to approximately 50 percent. It turns out that the explanatory power in the volatility factors mainly stem from the risk inherent to the yield curve’s slope and curvature. (iv) We document that riskiness regarding the yield curve shape (convexity) but not the riskiness regarding the slope induces a positive risk premium in excess returns. Actually, we find that slope uncertainties decrease future bond return premia revealing a compensation for stability in term structure slopes. (v) Yield term factors and - to an even larger extent - factor volatilities are closely connected to key macroeconomic variables reflecting capacity and production utilization, employment growth, inflation and monetary policy. (vi) We observe that macroeconomic variables have more long-run predictability for term structure volatilities than for the term structure itself. It turns out that capacity utilization and industrial production are important long-term predictors for risk inherent to the level and slope of the yield curve. Conversely, we observe that yield factors have significant forecasting ability for capacity utilization, employment growth and industrial production but only negligible impacts on the volatilities thereof. Nevertheless, we identify an important role of the curvature volatility for long-term predictions of macroeconomic variables. These results provide hints that risk inherent to the shape of the yield curve is relevant and seems to be effectively captured by a stochastic volatility component in the curvature factor.



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## A MCMC based Bayesian Inference

Let  $\Omega$  collect all model parameters including the latent variables, and let  $Y$  denote the observed data. Applying Clifford-Hammersley's theorem (see Hammersley and Clifford (1971), Besag (1974)), the posterior distribution

$$p(\Omega|Y) \propto p(Y|\Omega)p(\Omega) \quad (30)$$

can be broken up into a complete set of conditional posteriors,  $p(\Omega_i|\Omega_{-i}, Y)$ ,  $i = 1, \dots, N$ , where  $p(\Omega)$  denotes the prior distribution of  $\Omega$ ,  $N$  is the number of blocks,  $\Omega_i$  denotes the  $i$ -th block and  $\Omega_{-i}$  denotes all the elements of  $\Omega$  excluding  $\Omega_i$ . Then, the elements  $\Omega_i$  can be sampled according to the following Markov chain:

- Initialize  $\Omega^{(0)}$ .
- For  $i = 1, \dots, G$ :
  1. draw  $\Omega_1^{(i)}$  from  $p(\Omega_1|\Omega_2^{(i-1)}, \Omega_3^{(i-1)}, \dots, \Omega_N^{(i-1)}, Y)$ ,
  2. draw  $\Omega_2^{(i)}$  from  $p(\Omega_2|\Omega_1^{(i)}, \Omega_3^{(i-1)}, \dots, \Omega_N^{(i-1)}, Y)$ ,
  - $\vdots$
  - N. draw  $\Omega_N^{(i)}$  from  $p(\Omega_N|\Omega_1^{(i)}, \Omega_2^{(i)}, \dots, \Omega_{N-1}^{(i)}, Y)$ ,

where  $G$  is the number of MCMC iterations. In dependence of the form of the conditional posteriors we employ Gibbs or Metropolis-Hastings samplers as implemented in the software package BUGS (see Spiegelhalter, Thomas, Best, and Gilks (1996)). The procedure works well but is relatively inefficient in the given context. In order to guarantee a proper convergence of the Markov chain we run 2,500,000 MCMC iterations with a burn-in period of 500,000 iterations.<sup>17</sup>

All model parameters are assumed to be a priori independent and are distributed as follows:

- $\Sigma$  is the variance-covariance matrix with zero off-diagonal elements of equation (1). We assume that each of its elements follows an Inverse-Gamma(2.5,0.025) distribution with mean of 0.167 and standard deviation 0.024.
- For  $\lambda$  we assign a uniform distribution on the interval  $[0, 1]$ .
- For the persistent parameters of the yield curve factors  $\phi^i$ ,  $i = L, S, C$ , we assume their transformations  $(\phi^i + 1)/2$  to follow a beta distribution with parameters 20 and 1.5 implying a mean of 0.86 and a standard deviation of 0.11.

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<sup>17</sup>More efficient estimation algorithms for the model are on the future research agenda but are beyond the scope of the current paper.

- $\mu^i$ ,  $i = L, S, C$  in (3) are assumed to be independently normally distributed with mean 0 and variance 10.

- $h^i$ ,  $i = L, S, C$  in (3) are assumed to follow an Inverse-Gamma(2.5,0.025) distribution.

- For  $\phi_h^i$ ,  $i = L, S, C$  in (5), we assume their transformations  $(\phi^i + 1)/2$  to follow a beta distribution with parameters 20 and 1.5 implying a mean of 0.86 and a standard deviation of 0.11.

- $\mu_h^i$ ,  $i = L, S, C$  in (5) are assumed to be independently normally distributed with mean 0 and variance 10.

- $\sigma_h^i$ ,  $i = L, S, C$  in (5) are assumed to follow an Inverse-Gamma(2.5,0.025) distribution.

- $d^i$ ,  $i = L, S, C$  are assumed to be normally distributed with mean 0 and variance 10.

To test for the convergence of the generated Markov chain, we use the Z-score by Geweke (1992). Let  $\{\Omega^{(i)}\}_{i=1}^G$  denote the generated Markov chain with

$$\bar{\Omega}_1 = \frac{1}{G_1} \sum_{i=1}^{G_1} \Omega^{(i)}, \quad \bar{\Omega}_2 = \frac{1}{G_2} \sum_{i=p^*}^G \Omega^{(i)}, \quad p^* = G - G_2 + 1, \quad (31)$$

and let  $\hat{S}^1(0)$  and  $\hat{S}^2(0)$  denote consistent spectral density estimates (evaluated at zero) for  $\{\Omega^{(i)}\}_{i=1}^{G_1}$  and  $\{\Omega^{(i)}\}_{i=p^*}^G$ , respectively. If the sequence  $\{\Omega^{(i)}\}_{i=1}^G$  is stationary, then as  $G \rightarrow \infty$ ,

$$(\bar{\Omega}_1 - \bar{\Omega}_2)/[G_1^{-1}\hat{S}^1(0) + G_2^{-1}\hat{S}^2(0)] \xrightarrow{d} N(0, 1) \quad (32)$$

given the ratios  $G_1/G$  and  $G_2/G$  are fixed, and  $(G_1 + G_2)/G < 1$ . Geweke (1992) suggests using  $G_1 = 0.1G$  and  $G_2 = 0.5G$ .

Table 1: MCMC estimation results for dynamic Nelson-Siegel models. Based on monthly observations of unsmoothed U.S. Fama-Bliss zero coupon yields from January 1964 to December 2003 with maturities of one to five years. 480 observations.

A. Model without volatility factors										
	$\mu^L$	$\mu^S$	$\mu^C$	$\phi^L$	$\phi^S$	$\phi^C$	$h^L$	$h^S$	$h^C$	$\lambda$
Mean	0.028	-0.053	-0.028	0.994	0.956	0.906	0.307	0.597	0.757	0.055
SD	0.014	0.030	0.036	0.003	0.013	0.023	0.022	0.024	0.060	0.006
95% CI, lower	0.007	-0.116	-0.099	0.987	0.929	0.860	0.266	0.552	0.637	0.046
95% CI, upper	0.062	-0.001	0.042	0.998	0.980	0.947	0.351	0.645	0.871	0.069
Z-score	-0.171	1.198	1.615	0.196	1.352	-0.507	-1.737	-0.381	1.418	0.172

B. Model with stochastic volatility factors																
	$\mu^L$	$\mu^S$	$\mu^C$	$\phi^L$	$\phi^S$	$\phi^C$	$\mu_h^L$	$\mu_h^S$	$\mu_h^C$	$\phi_h^L$	$\phi_h^S$	$\phi_h^C$	$(\sigma_h^L)^2$	$(\sigma_h^S)^2$	$(\sigma_h^C)^2$	$\lambda$
Mean	0.023	-0.008	-0.025	0.994	0.981	0.879	-0.082	-0.083	-0.038	0.977	0.964	0.933	0.246	0.343	0.269	0.045
SD	0.011	0.008	0.034	0.002	0.006	0.024	0.053	0.037	0.023	0.016	0.017	0.034	0.078	0.070	0.083	0.002
95% CI, lower	0.005	-0.028	-0.092	0.988	0.966	0.830	-0.214	-0.169	-0.096	0.936	0.923	0.851	0.134	0.225	0.042	0.042
95% CI, upper	0.049	0.004	0.042	0.998	0.993	0.925	-0.017	-0.024	-0.005	0.996	0.992	0.987	0.439	0.498	0.048	0.048
Z-score	-1.647	-0.902	-1.085	1.615	0.878	1.463	-0.594	0.772	-1.196	-0.486	0.532	-1.383	0.378	-0.493	0.454	0.568

"95% CI" denotes the 95% credibility interval of the posterior distribution. The Z-score statistic is the Geweke (1992) test statistic for the convergence of MCMC samples, see Appendix A.

Table 2: Summary statistics of the data and model residuals

A. Zero yields from January 1964 to December 2003																							
	Autocorrelation of residuals												Autocorrelation of squared residuals										
	Mean	SD	1	2	3	4	5	6	7	8	9	10	1	2	3	4	5	6	7	8	9	10	
1	6.520	2.719	0.975	0.945	0.919	0.895	0.876	0.854	0.832	0.818	0.800	0.780	0.963	0.920	0.889	0.865	0.852	0.829	0.808	0.802	0.786	0.769	
2	6.741	2.633	0.979	0.953	0.929	0.908	0.890	0.870	0.851	0.834	0.816	0.798	0.973	0.938	0.911	0.890	0.874	0.854	0.835	0.823	0.808	0.792	
3	6.914	2.538	0.980	0.957	0.936	0.917	0.900	0.882	0.863	0.846	0.829	0.811	0.977	0.948	0.926	0.908	0.893	0.875	0.855	0.840	0.824	0.808	
4	7.049	2.484	0.980	0.959	0.941	0.922	0.906	0.889	0.870	0.854	0.836	0.818	0.978	0.953	0.934	0.917	0.903	0.885	0.864	0.849	0.831	0.814	
5	7.127	2.439	0.982	0.963	0.945	0.928	0.913	0.897	0.880	0.864	0.848	0.832	0.981	0.958	0.940	0.924	0.911	0.894	0.875	0.861	0.846	0.831	

B. Residuals of the model with constant volatility factors																							
	Autocorrelation of residuals												Autocorrelation of squared residuals										
	Mean	SD	MAE	1	2	3	4	5	6	7	8	9	10	1	2	3	4	5	6	7	8	9	10
1	0.002	0.049	0.0355	-0.114	-0.069	0.169	-0.077	0.026	0.055	-0.142	0.088	0.040	-0.012	0.223	0.135	0.158	0.179	0.129	0.042	0.137	0.150	0.202	0.145
2	-0.006	0.066	0.0487	0.360	0.257	0.316	0.130	0.073	0.076	0.086	0.105	0.198	0.147	0.221	0.214	0.078	0.241	0.191	0.116	0.141	0.115	0.252	0.208
3	-0.001	0.046	0.0307	0.114	0.073	0.194	0.038	0.036	0.171	-0.005	-0.008	0.091	-0.007	0.140	0.050	0.248	0.126	0.036	0.022	0.118	0.013	0.046	0.035
4	0.011	0.066	0.0463	0.283	0.157	0.386	0.159	0.124	0.188	0.017	0.084	0.006	-0.089	0.196	0.142	0.329	0.163	0.078	0.149	0.057	0.089	0.181	0.246
5	-0.003	0.036	0.0267	0.049	0.042	0.042	0.005	0.002	0.069	0.015	-0.026	0.096	-0.107	0.288	0.208	0.236	0.144	0.171	0.155	0.105	0.241	0.121	0.156

C. Residuals of the model with stochastic volatility factors																							
	Autocorrelation of residuals												Autocorrelation of squared residuals										
	Mean	SD	MAE	1	2	3	4	5	6	7	8	9	10	1	2	3	4	5	6	7	8	9	10
1	0.001	0.035	0.027	-0.108	-0.100	0.089	0.006	0.021	0.031	-0.115	0.007	0.088	0.020	0.065	0.019	0.056	0.032	0.009	0.000	0.010	0.065	0.134	0.033
2	-0.009	0.071	0.053	0.346	0.249	0.327	0.129	0.085	0.088	0.076	0.104	0.196	0.139	0.198	0.204	0.064	0.176	0.207	0.078	0.134	0.121	0.220	0.193
3	-0.003	0.046	0.032	0.141	0.086	0.182	0.036	0.023	0.164	-0.021	-0.004	0.085	-0.010	0.155	0.063	0.226	0.131	0.045	0.019	0.119	0.020	0.063	0.016
4	0.009	0.065	0.045	0.301	0.182	0.375	0.162	0.127	0.174	0.020	0.090	0.005	-0.090	0.203	0.167	0.329	0.172	0.079	0.159	0.052	0.100	0.183	0.255
5	-0.005	0.037	0.029	0.087	0.052	0.081	0.038	0.018	0.033	-0.003	-0.014	0.075	-0.105	0.139	0.151	0.124	0.103	0.012	0.009	-0.001	0.091	0.086	0.081

The individual rows are associated with the corresponding maturities of the underlying data. The columns give the corresponding lags. MAE denotes the mean absolute error defined as  $\text{MAE} = \frac{1}{T} \sum_{t=1}^T |y_t^{(n)} - \hat{y}_t^{(n)}|$ .

Table 3: Monthly regressions of one-year-ahead bond excess returns on Nelson-Siegel yield factors and factor volatilities. Yield factors and factor volatilities extracted from monthly observations of unsmoothed U.S. Fama-Bliss zero coupon yields from January 1964 to December 2003 with maturities of one to five years.

A. Regression: $z_t^{(n)} = c + \beta_L L_{t-12} + \beta_S S_{t-12} + \beta_C C_{t-12} + \varepsilon_t^{(n)}$																									
n	All yield factors							No level factor						No slope factor						No curvature factor					
	c	$\beta_L$	$\beta_S$	$\beta_C$	Adj. $R^2$	HH	NW	c	$\beta_S$	$\beta_C$	Adj. $R^2$	HH	NW	c	$\beta_L$	$\beta_C$	Adj. $R^2$	HH	NW	c	$\beta_L$	$\beta_S$	Adj. $R^2$	HH	NW
2	-0.92 (-1.31)	0.14 (1.54)	-0.25 (-2.37)	0.51 (4.41)	0.31	43.8	49.9	0.18 (0.61)	-0.26 (-2.09)	0.57 (4.10)	0.29	34.4	39.8	-0.68 (-0.73)	0.15 (1.22)	0.53 (3.80)	0.25	25.3	29.2	-1.78 (-1.88)	0.24 (1.84)	-0.27 (-2.21)	0.17	12.2	15.3
3	-0.92 (-0.70)	0.15 (0.82)	-0.51 (-2.48)	1.01 (4.62)	0.32	41.3	46.9	0.19 (0.36)	-0.52 (-2.19)	1.07 (4.36)	0.32	36.7	42.5	-0.44 (-0.25)	0.17 (0.71)	1.06 (4.01)	0.24	23.4	26.7	-2.63 (-1.46)	0.35 (1.35)	-0.57 (-2.31)	0.15	9.8	12.5
4	-1.40 (-0.77)	0.18 (0.74)	-0.79 (-2.78)	1.42 (4.66)	0.35	40.6	47.4	0.02 (0.03)	-0.81 (-2.45)	1.50 (4.46)	0.34	37.0	43.8	-0.65 (-0.26)	0.23 (0.64)	1.49 (3.92)	0.25	23.1	26.5	-3.80 (-1.53)	0.48 (1.31)	-0.87 (-2.57)	0.17	10.9	13.9
5	-1.86 (-0.84)	0.21 (0.69)	-1.01 (-2.93)	1.78 (4.95)	0.36	44.6	51.6	-0.26 (-0.34)	-1.02 (-2.57)	1.87 (4.76)	0.36	40.8	47.9	-0.92 (-0.31)	0.26 (0.60)	1.88 (4.13)	0.25	25.3	28.6	-4.87 (-1.62)	0.57 (1.30)	-1.10 (-2.69)	0.18	11.5	14.7
B. Regression: $z_t^{(n)} = c + \alpha_L h_{t-12}^L + \alpha_S h_{t-12}^S + \alpha_C h_{t-12}^C + \varepsilon_t^{(n)}$																									
n	All volatility factors							No level volatility factor						No slope volatility factor						No curvature volatility factor					
	c	$\alpha_L$	$\alpha_S$	$\beta_C$	Adj. $R^2$	HH	NW	c	$\alpha_S$	$\alpha_C$	Adj. $R^2$	HH	NW	c	$\alpha_L$	$\alpha_C$	Adj. $R^2$	HH	NW	c	$\alpha_L$	$\alpha_S$	Adj. $R^2$	HH	NW
2	-1.37 (-2.30)	3.50 (1.25)	-2.95 (-2.18)	3.11 (4.02)	0.18	22.1	22.3	-1.22 (-2.00)	-2.21 (-1.9)	3.52 (3.76)	0.14	13.9	14.4	-0.81 (-1.08)	0.84 (0.35)	1.46 (1.44)	0.06	2.18	2.77	0.14 (0.24)	4.69 (1.58)	-1.59 (-1.15)	0.07	2.54	3.18
3	-2.25 (-2.15)	5.42 (1.04)	-5.52 (-2.35)	5.60 (3.93)	0.16	20.6	21.7	-2.02 (-1.91)	-4.37 (-2.23)	6.23 (3.98)	0.14	15.9	16.7	-1.19 (-0.85)	0.44 (0.09)	2.50 (1.25)	0.04	1.53	1.93	0.48 (0.44)	7.56 (1.37)	-3.08 (-1.26)	0.06	2.16	2.71
4	-2.91 (-2.09)	7.37 (1.01)	-7.95 (-2.51)	7.54 (3.77)	0.16	20.6	21.5	-2.59 (-1.80)	-6.39 (-2.41)	8.40 (3.99)	0.11	16.0	16.9	-1.38 (-0.71)	0.21 (0.03)	3.07 (1.06)	0.03	1.11	1.39	0.77 (0.533)	10.26 (1.34)	-4.67 (-1.39)	0.07	2.32	2.92
5	-3.59 (-2.15)	8.95 (1.01)	-9.72 (-2.49)	8.95 (3.65)	0.16	19.3	20.2	-3.21 (-1.85)	-7.82 (-2.41)	9.99 (3.90)	0.13	15.3	16.2	-1.73 (-0.73)	0.19 (0.02)	3.49 (0.98)	0.02	0.95	1.19	0.77 (0.44)	12.38 (1.33)	-5.82 (-1.42)	0.07	2.35	2.95
C. Regression: $z_t^{(n)} = c + \beta_L L_{t-12} + \beta_S S_{t-12} + \beta_C C_{t-12} + \alpha_L h_{t-12}^L + \alpha_S h_{t-12}^S + \alpha_C h_{t-12}^C + \varepsilon_t^{(n)}$																									
n	All factors																								
	c	$\beta_L$	$\beta_S$	$\beta_C$	$\alpha_L$	$\alpha_S$	$\alpha_C$	Adj. $R^2$	HH	NW															
2	-4.51 (-5.82)	0.56 (2.73)	0.01 (0.10)	0.53 (6.56)	-6.30 (-1.87)	-3.13 (-2.90)	4.65 (4.88)	0.47	114.8	120.4															
3	-7.31 (-4.88)	0.87 (2.37)	-0.04 (-0.20)	1.03 (6.67)	-11.1 (-1.86)	-5.54 (-2.91)	8.29 (4.57)	0.46	105.9	109.4															
4	-10.4 (-5.04)	1.33 (2.87)	-0.11 (-0.40)	1.45 (6.81)	-17.2 (-2.26)	-7.74 (-3.13)	10.9 (4.32)	0.49	112.2	117.0															
5	-12.7 (-5.09)	1.65 (3.11)	-0.17 (-0.49)	1.80 (7.22)	-21.5 (-2.45)	-9.31 (-3.05)	13.1 (4.13)	0.50	118.8	122.8															

$z_t^{(n)}$  denotes the  $n$ -year one-year ahead bond excess return.  $L_t$ ,  $S_t$  and  $C_t$  denote the estimated level, slope and curvature factors, respectively. Their corresponding volatility factors are  $h_t^L$ ,  $h_t^S$  and  $h_t^C$ . Both yield curve factors and volatility factors are extracted from model (1), (3) and (5). HH and NW are  $\chi^2$  statistics for joint significance tests using Hansen-Hodrick and Newey-West corrections, respectively. The 5-percent critical values for  $\chi^2(2)$ ,  $\chi^2(3)$  and  $\chi^2(6)$  are 5.99, 7.82 and 12.59. The robust  $t$ -statistics based on HH corrections are reported in parentheses “( )”.

Table 4: Monthly regressions of one-year-ahead bond excess returns on PCA factors, the Cochrane-Piazzesi forecasting factor and Nelson-Siegel yield factors. Yield factors extracted from monthly observations of unsmoothed U.S. Fama-Bliss zero coupon yields from January 1964 to December 2003 with maturities of one to five years.

A.									
n	PCA factors			Return-forecasting factor			Nelson-Siegel factors		
	Adj. $R^2$	HH	$p$ -value	Adj. $R^2$	HH	$p$ -value	Adj. $R^2$	HH	$p$ -value
2	0.204	36.090	0.000	0.309	65.241	0.000	0.313	49.937	0.000
3	0.212	35.984	0.000	0.336	60.443	0.000	0.321	46.971	0.000
4	0.241	34.375	0.000	0.370	55.896	0.000	0.348	47.469	0.000
5	0.247	35.338	0.000	0.343	46.686	0.000	0.361	51.607	0.000

B. Forecasting regressions: $z_t^{(n)} = \beta_L L_{t-12} + \beta_S S_{t-12} + \beta_C C_{t-12} + \varphi \vartheta_t + \varepsilon_t^{(n)}$									
n	$\beta_L$	$\beta_S$	$\beta_C$	$\varphi$	Adj. $R^2$	HH	NW		
2	0.028 (0.917)	-0.046 (-0.530)	0.283 (2.024)	0.297 (3.760)	0.336	104.524	99.876		
3	0.022 (0.402)	-0.137 (-0.795)	0.533 (2.172)	0.563 (3.862)	0.358	90.910	89.637		
4	0.000 (0.007)	-0.245 (-1.016)	0.717 (2.173)	0.830 (4.167)	0.390	81.581	85.855		
5	-0.030 (-0.356)	-0.472 (-1.618)	1.130 (2.782)	0.786 (3.110)	0.381	67.406	73.079		

C. Forecasting regressions: $z_t^{(n)} = \alpha_L h_{t-12}^L + \alpha_S h_{t-12}^S + \alpha_C h_{t-12}^C + \varphi \vartheta_{t-12} + \varepsilon_t^{(n)}$									
n	$\alpha_L$	$\alpha_S$	$\alpha_C$	$\varphi$	Adj. $R^2$	HH	NW		
2	-1.039 (-0.552)	-0.956 (-0.993)	1.135 (2.114)	0.442 (6.730)	0.343	113.181	121.356		
3	-3.451 (-1.025)	-1.849 (-1.117)	2.394 (2.394)	0.870 (6.747)	0.381	122.700	131.127		
4	-5.411 (-1.202)	-2.758 (-1.254)	3.361 (2.569)	1.278 (6.967)	0.421	135.324	145.272		
5	-6.198 (-1.094)	-3.514 (-1.290)	3.795 (2.364)	1.509 (6.524)	0.393	109.265	121.594		

$z_t^{(n)}$  denotes the one-year-ahead bond excess return of  $n$ -year bonds.  $L_t$ ,  $S_t$  and  $C_t$  denote the estimated level, slope and curvature factors, respectively. Their corresponding volatility factors are  $h_t^L$ ,  $h_t^S$  and  $h_t^C$ . Both yield curve factors and volatility factors are extracted from model (1), (3) and (5).  $\vartheta_t$  denotes the return-forecasting factor of Cochrane and Piazzesi (2005). HH and NW are  $\chi^2$  statistics for joint significance tests using Hansen-Hodrick and Newey-West corrections, respectively. The 5-percent critical value of  $\chi^2(4)$  is 9.49.



Table 5: Linear regressions of monthly yield factors  $L_t$ ,  $S_t$ ,  $C_t$ , and factor volatilities  $h_t^L$ ,  $h_t^S$ ,  $h_t^C$ , on log changes of the consumer price index (INF), capacity utilization (CU), employment growth rate (EMP), the federal funds rate (FFR) and industrial production (IP). Yield factors and factor volatilities extracted from monthly observations of unsmoothed U.S. Fama-Bliss zero coupon yields from January 1964 to December 2003 with maturities of one to five years. Robust standard errors in parantheses.

	CONST	INF	CU	EMPLOY	FFR	IP	$R^2$
$L$	27.781 (3.355)	-0.258 (0.160)	-0.298 (0.044)	27.120 (17.756)	0.495 (0.035)	5.324 (4.787)	0.77
$S$	-29.404 (3.302)	0.159 (0.184)	0.322 (0.044)	-13.601 (19.192)	0.357 (0.029)	-7.225 (4.749)	0.69
$C$	10.248 (4.302)	-0.697 (0.378)	-0.147 (0.050)	38.976 (18.406)	0.104 (0.091)	-1.060 (6.277)	0.12
$h^L$	0.905 (0.206)	-0.040 (0.014)	-0.012 (0.003)	0.041 (0.827)	0.035 (0.002)	0.465 (0.285)	0.76
$h^S$	-0.058 (0.461)	-0.029 (0.037)	0.002 (0.006)	-3.207 (1.929)	0.071 (0.008)	-1.202 (0.846)	0.63
$h^C$	1.879 (0.530)	0.063 (0.038)	-0.014 (0.006)	-7.449 (2.316)	0.029 (0.010)	1.371 (0.767)	0.48

Table 6: VAR(1) estimates of the monthly yield factors  $L_t$ ,  $S_t$ ,  $C_t$ , log changes of the consumer price index (INF), capacity utilization (CU), employment growth rate (EMP), the federal funds rate (FFR) and industrial production (IP). Yield factors and factor volatilities extracted from monthly observations of unsmoothed U.S. Fama-Bliss zero coupon yields from January 1964 to December 2003 with maturities of one to five years. Robust standard errors in parantheses.

	$L_t$	$S_t$	$C_t$	$INF_t$	$CU_t$	$FFR_t$	$IP_t$	$EMPLOY_t$
$L_{t-1}$	0.966 (0.028)	0.085 (0.085)	-0.096 (0.072)	-0.011 (0.0264)	0.311 (0.059)	0.508 (0.133)	0.004 (0.001)	0.001 (0.000)
$S_{t-1}$	-0.010 (0.023)	0.982 (0.069)	-0.078 (0.063)	0.005 (0.023)	0.231 (0.055)	0.466 (0.127)	0.003 (0.001)	0.001 (0.000)
$C_{t-1}$	-0.011 (0.009)	0.020 (0.016)	0.894 (0.023)	-0.018 (0.012)	-0.013 (0.020)	-0.031 (0.0213)	0.000 (0.000)	0.000 (0.000)
$INF_{t-1}$	0.046 (0.046)	-0.026 (0.108)	-0.188 (0.113)	0.211 (0.072)	0.091 (0.077)	0.083 (0.078)	0.001 (0.002)	0.000 (0.000)
$CU_{t-1}$	-0.004 (0.004)	0.026 (0.013)	-0.018 (0.016)	0.001 (0.006)	0.973 (0.015)	0.006 (0.012)	0.000 (0.000)	0.000 (0.000)
$FFR_{t-1}$	0.023 (0.020)	-0.066 (0.074)	0.098 (0.059)	0.007 (0.022)	-0.278 (0.050)	0.572 (0.114)	-0.004 (0.001)	-0.001 (0.000)
$IP_{t-1}$	-0.222 (0.465)	-0.978 (1.272)	-1.359 (2.014)	0.949 (0.609)	4.190 (1.752)	1.466 (1.198)	1.017 (0.029)	0.039 (0.008)
$EMPLOY_{t-1}$	0.975 (1.463)	3.674 (3.991)	8.901 (5.210)	2.381 (1.768)	-2.675 (3.702)	-4.390 (4.483)	-0.168 (0.074)	0.908 (0.022)
$CONST$	0.412 (0.355)	-2.428 (1.072)	1.225 (1.410)	-0.138 (0.501)	1.945 (1.249)	-0.812 (0.977)	0.006 (0.021)	-0.001 (0.006)

Table 7: VAR(1) estimates of the monthly factor volatilities  $h_t^L$ ,  $h_t^S$ ,  $h_t^C$ , log changes of the consumer price index (INF), capacity utilization (CU), employment growth rate (EMP), the federal funds rate (FFR) and industrial production (IP). Yield factors and factor volatilities extracted from monthly observations of unsmoothed U.S. Fama-Bliss zero coupon yields from January 1964 to December 2003 with maturities of one to five years. Robust standard errors in parantheses.

	$h_t^L$	$h_t^S$	$h_t^C$	$INF_t$	$CU_t$	$FFR_t$	$IP_t$	$EMPLOY_t$
$h_{t-1}^L$	0.981 (0.010)	-0.116 (0.033)	-0.058 (0.038)	-0.491 (0.246)	-4.829 (3.014)	1.138 (0.732)	0.019 (0.012)	0.005 (0.003)
$h_{t-1}^S$	0.014 (0.003)	0.987 (0.028)	0.019 (0.014)	-0.145 (0.070)	-2.614 (1.106)	-0.055 (0.416)	0.003 (0.005)	0.001 (0.002)
$h_{t-1}^C$	0.007 (0.003)	0.019 (0.014)	1.003 (0.016)	0.212 (0.102)	5.321 (1.469)	0.458 (0.286)	-0.011 (0.006)	-0.004 (0.002)
$INF_{t-1}$	0.001 (0.001)	0.016 (0.010)	0.004 (0.005)	0.190 (0.075)	0.888 (0.339)	0.121 (0.087)	0.002 (0.002)	0.001 (0.001)
$CU_{t-1}$	0.000 (0.000)	0.001 (0.001)	0.000 (0.000)	0.007 (0.006)	0.586 (0.114)	0.013 (0.015)	0.000 (0.000)	0.000 (0.000)
$FFR_{t-1}$	0.000 (0.000)	0.005 (0.002)	0.001 (0.002)	0.023 (0.011)	0.361 (0.156)	0.945 (0.026)	-0.001 (0.000)	0.000 (0.000)
$IP_{t-1}$	0.017 (0.020)	-0.064 (0.083)	-0.069 (0.098)	0.634 (0.620)	-31.098 (11.181)	-1.487 (1.568)	1.025 (0.027)	0.044 (0.008)
$EMPLOY_{t-1}$	0.095 (0.056)	0.274 (0.219)	0.325 (0.283)	2.436 (1.708)	143.696 (36.878)	9.511 (3.923)	-0.162 (0.076)	0.904 (0.024)
$CONST$	-0.027 (0.014)	-0.074 (0.058)	-0.016 (0.073)	-0.751 (0.511)	25.893 (8.430)	-1.460 (1.315)	0.030 (0.027)	0.010 (0.008)

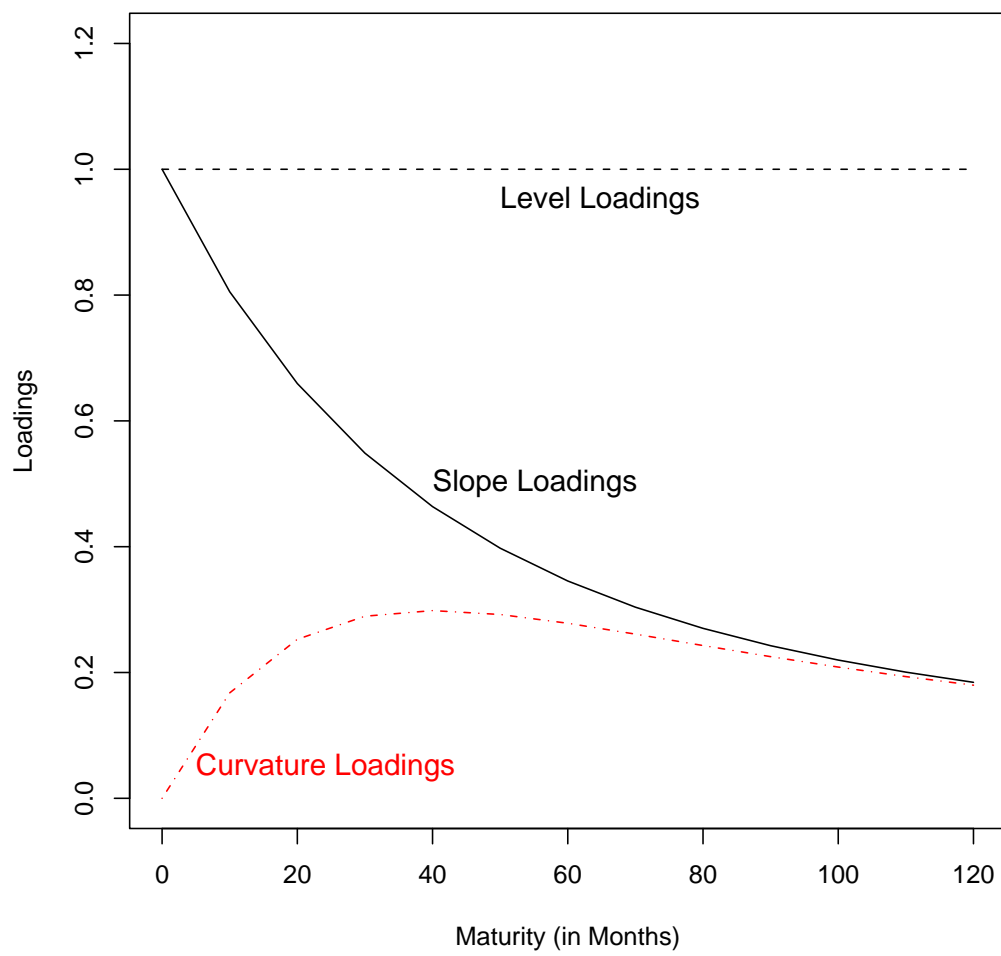


Figure 1: Plot of the Nelson-Siegel factor loadings.  $\lambda = 0.045$ .



Figure 2: The estimated yield factors (solid lines) and their empirical approximation (dotted lines).

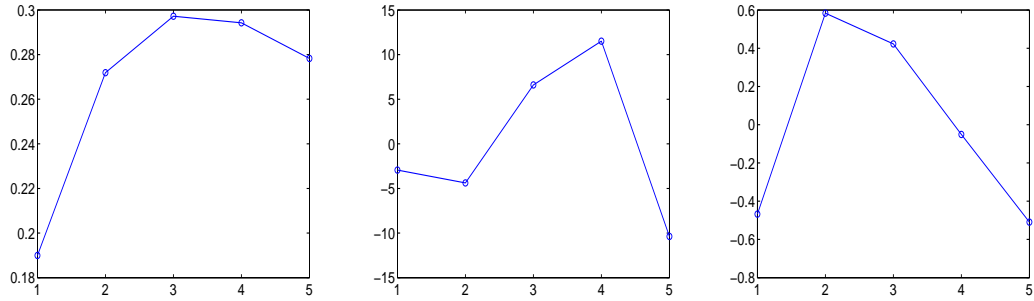


Figure 3: Loadings on the Nelson-Siegel curvature factor (left),  $\lambda = 0.045$ , the return-forecasting factor (middle) and the PC curvature factor (right).

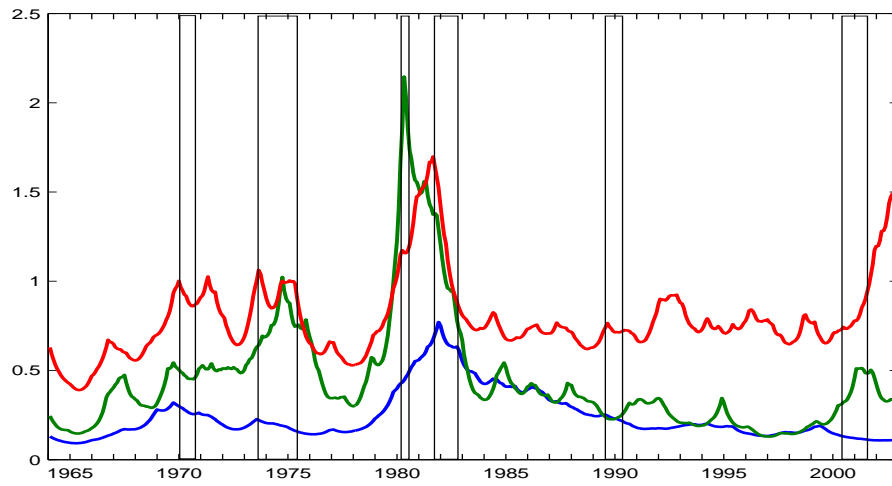


Figure 4: The estimated level volatility factor (blue line, top), the slope volatility factor (green line, middle) and curvature volatility factor (red line, bottom).

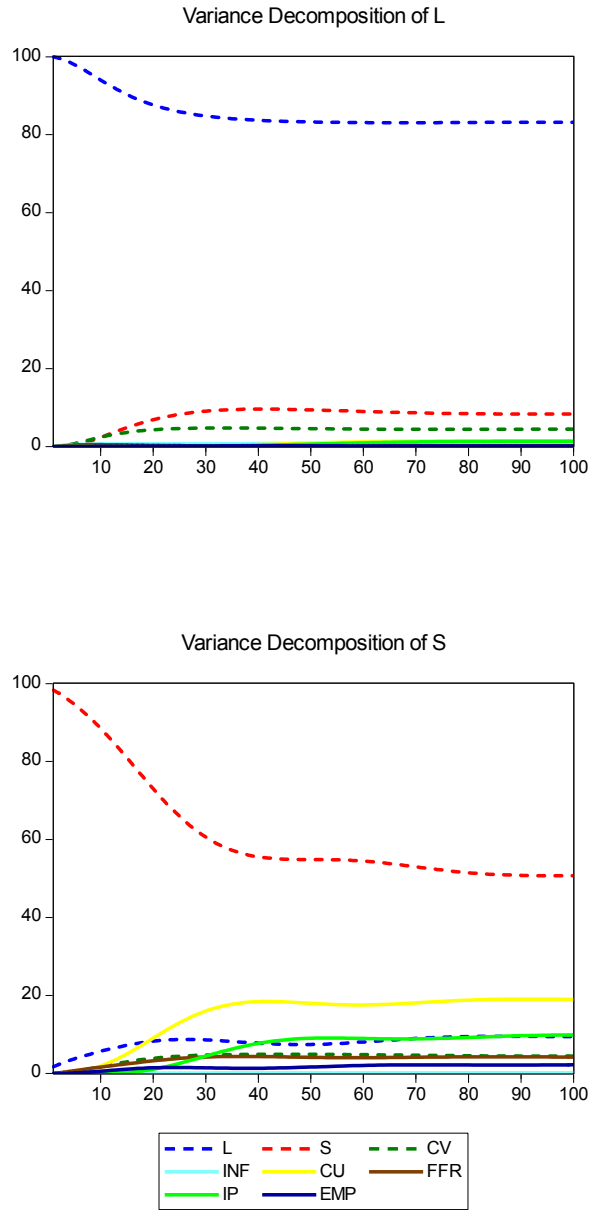


Figure 5: Prediction error decompositions of the level and slope factor. Based on a VAR(1) model of yield factors and macro factors using a Cholesky decomposition of the covariance.

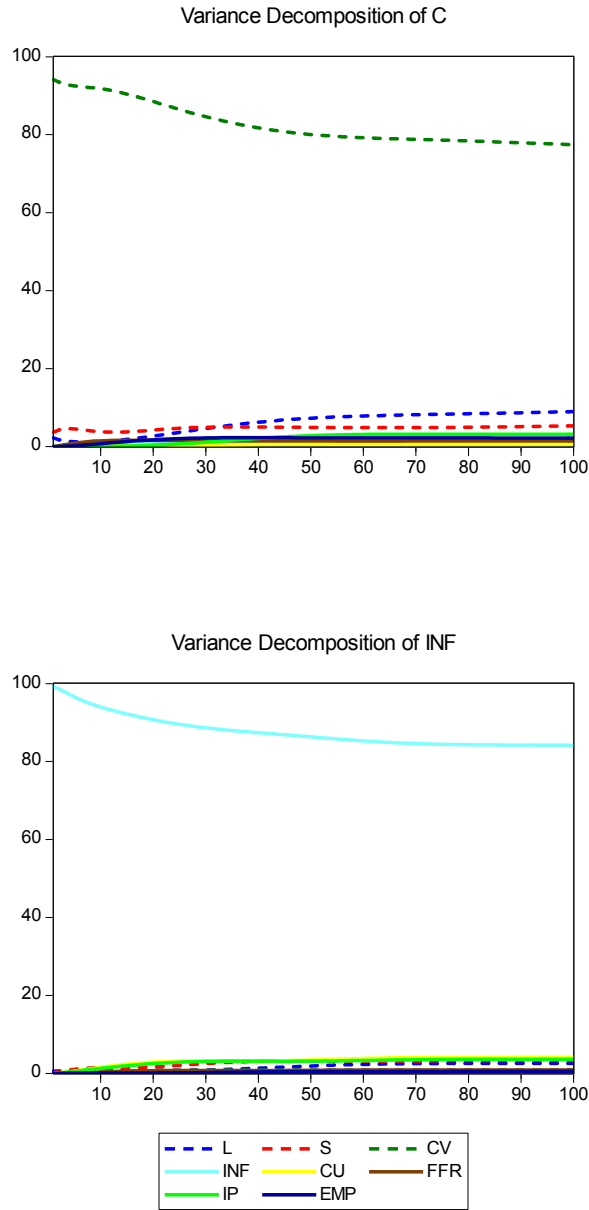


Figure 6: Prediction error decomposition of the curvature factor and DCPI factor. Based on a VAR(1) model of yield factors and macro factors using a Cholesky decomposition of the covariance.

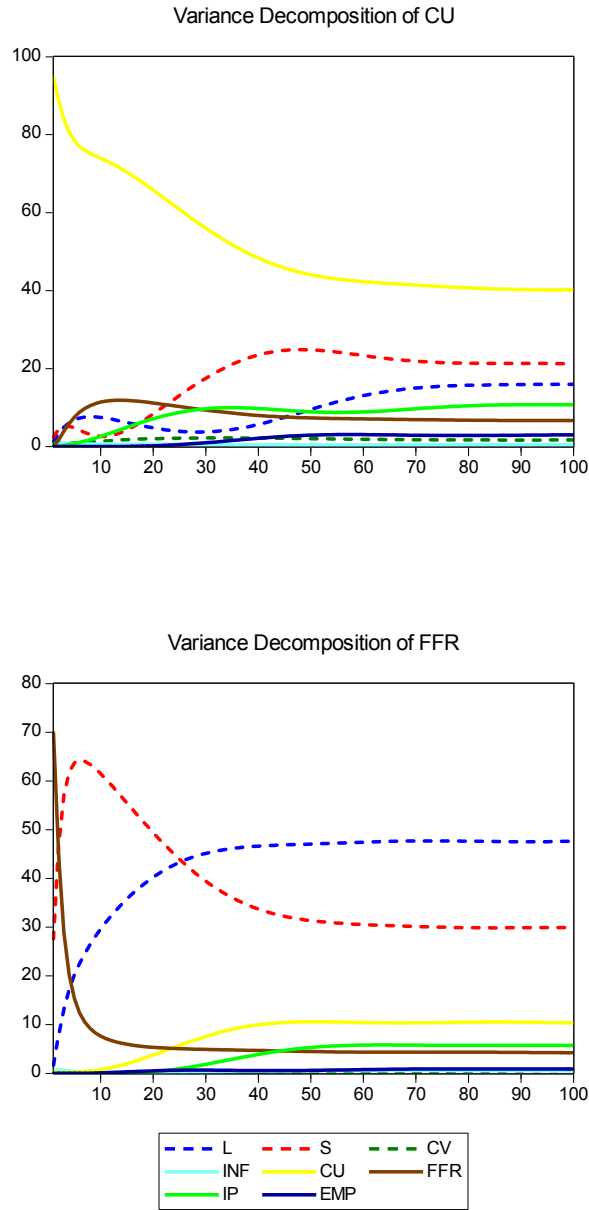


Figure 7: Prediction error decomposition of capacity utilization and of the federal funds rate. Based on a VAR(1) model of yield factors and macro factors using a Cholesky decomposition of the covariance.



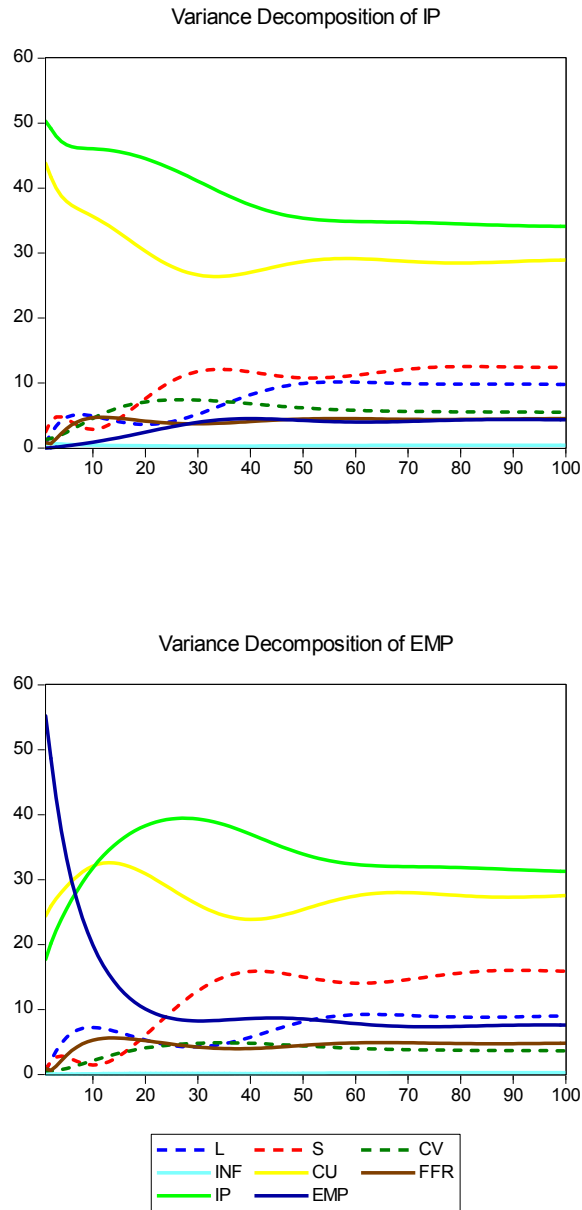


Figure 8: Prediction error decomposition of industrial production and of the employment growth rate. Based on a VAR(1) model of yield factors and macro factors using a Cholesky decomposition of the covariance.

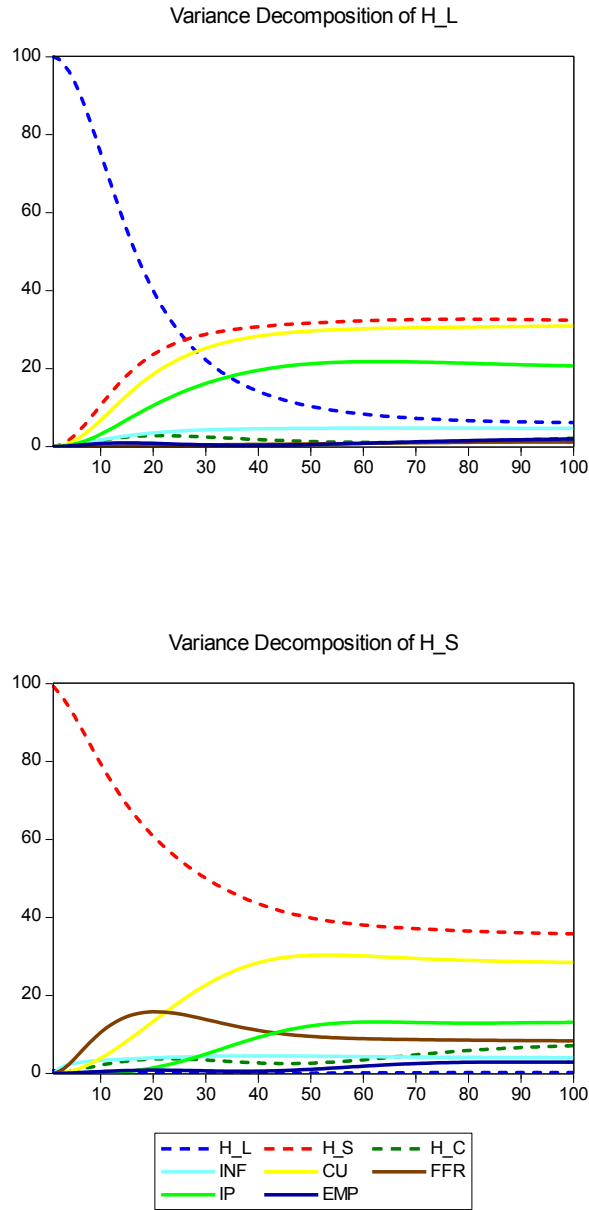


Figure 9: Prediction error decomposition of the level and slope volatility. Based on a VAR(1) model of volatility factors and macro factors using a Cholesky decomposition of the covariance.

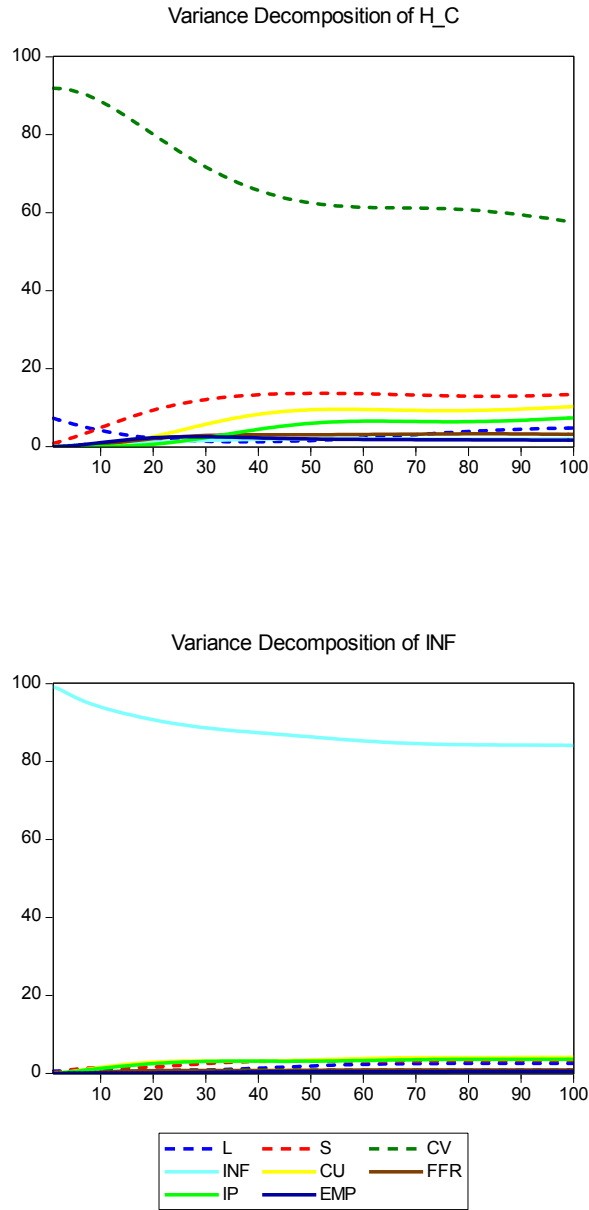


Figure 10: Prediction error decomposition of the curvature volatility and DCPI. Based on a VAR(1) model of volatility factors and macro factors using a Cholesky decomposition of the covariance.

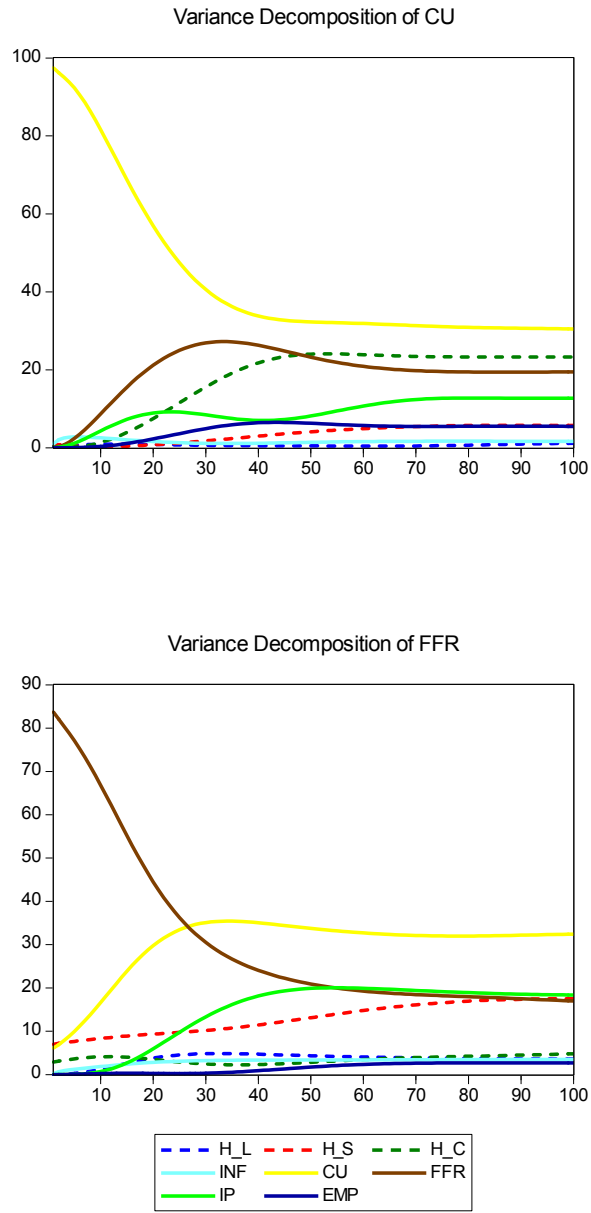


Figure 11: Prediction error decomposition of capacity utilization and of the federal funds rate. Based on a VAR(1) model of volatility factors and macro factors using a Cholesky decomposition of the covariance.

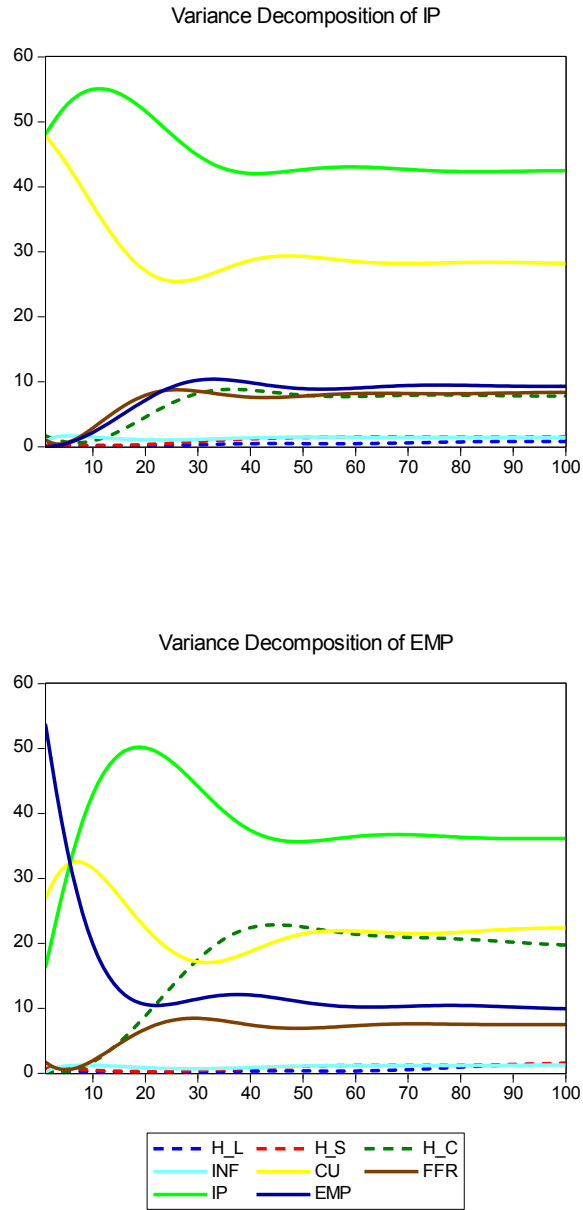


Figure 12: Prediction error decomposition of industrial production and of the employment growth rate. Based on a VAR(1) model of volatility factors and macro factors using a Cholesky decomposition of the covariance.

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