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TERM PREMIA IMPLICATIONS OF MACROECONOMIC REGIME CHANGES

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Term premia are shown to provide crucial information for discriminating among alternative sources of change in the economy, namely shifts in the variance of structural shocks and in monetary policy. These sources have been identified as competing explanations for time-varying features of major industrial economies during the 1980s and 1990s. Although hardly distinguishable through the lens of standard dynamic stochastic general equilibrium (DSGE) models, lower nonpolicy shock variances and "tighter" monetary policy regimes imply higher and lower term premia, respectively. As a result, moving to tighter monetary policy alone cannot explain the improved U.S. macroeconomic stability in the 1980s and 1990s: term premia would have shifted downward, a fact inconsistent with the evidence of higher premia from early 80s onward. Conversely, favorable shifts in nonpolicy innovation variance imply movements in term premia that are at least qualitatively consistent with historical patterns.

Keywords: Term Premia, Regime Switching, DSGE Models

1. INTRODUCTION

The contribution of this paper is to show how term premia provide crucial information for discriminating among alternative sources of change in the economy, namely shifts in the variance of structural shocks and changes in the conduct of monetary policy. Notably, a vast literature has identified these as competing explanations for time-varying features of the economic environment experienced by major industrial economies during the 1980s and 1990s. On one hand, Stock and Watson (2002), Sims and Zha (2006), and Justiniano and Primiceri (2008), among many others, find that specifications that allow time variation in the variances of structural disturbances fit best the changes in the reduced-form properties of the U.S. economy over the last decades; once heteroskedasticity in structural shocks is allowed for, there is no clear evidence of other changes in the economy, and notably in the conduct of monetary policy.[1] In contrast, Clarida et al. (2000), Lubik and Schorfheide (2004), and Boivin and Giannoni (2006), among others, stress the role

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played by monetary policy in achieving better macroeconomic performance; according to this interpretation, the change in U.S. monetary policy around the early 1980s from a "passive" to an "active" regime led eventually to greater macroeconomic stability by achieving equilibrium determinacy and thereby suppressing economic fluctuations induced by self-fulfilling expectations. In an attempt to reconcile these alternative findings, Benati and Surico (2009), for instance, illustrate how vector autoregressive (VAR) evidence, typically interpreted as supportive of time variation in shock variances, is in fact compatible with changes in monetary policy within a New Keynesian framework. In essence, their emphasis is on the fact that variations in structural shock variances are hardly distinguishable from changes in policy rule coefficients within standard dynamic stochastic general equilibrium (DSGE) models. In a similar vein, Davig and Doh (2009) make the point that, within a standard DSGE model, regime changes in monetary policy and in shock volatility affect inflation persistence equally. Both sources are found to have contributed to the reduction of historical inflation persistence for the U.S. economy; however, shifts in monetary regime have a quantitatively larger impact. More broadly, that such distinct sources of change appear to manifest themselves in similar economic dynamics within standard DSGE models casts doubts about the usefulness of such models for interpreting major economic events and thus informing policy decisions.

Our main finding is that, although similarly implying a reduction in the volatility of both inflation and output within standard DSGE models, regimes of lower shock variances and of tighter monetary policy imply higher and lower term premia, respectively. Defined in terms of expected excess returns, term premia are captured by the covariance between long bond prices and the pricing kernel. Our finding draws precisely upon the impact of the two alternative sources of change on the covariances between the stochastic discount factor and the relevant macro variables. Intuitively, as implied by standard consumption-based asset pricing models, a negative covariance between long bond prices and the pricing kernel means that financial assets carry low payoffs in bad times, and thus fail to provide insurance when it is most needed, and hence command positive premia. First, central to the finding that smaller variances of nonpolicy shocks are associated with higher premia is the induced fall in the positive autocovariances of both real output and inflation; this in turn translates into lower positive covariance between long bond prices and the pricing kernel. With low payoffs in bad times, long bonds end up commanding higher premia. Second, tighter monetary policy regimes bring about a reduction in term premia ultimately because they better insulate inflation and output from various shocks; this implies that the pricing kernel also tends to be less responsive to macroeconomic disturbances, and then less negatively correlated with long bond prices. In essence, tighter regimes induce lower premia by making inflation less negatively correlated with output growth, and thus long bonds less risky. All in all, the model's prediction is that a more stable macroeconomic environment is characterized by (i) higher term premia if such improved stability results from a reduction in the variance of nonpolicy shocks; (ii) lower term premia if such better stability is instead induced by tighter monetary policies. When the improved U.S. macroeconomic stability of the 1980s and 1990s is considered through the interpretative filter of a standard DSGE model, the implication is that a move to tighter monetary policy regimes alone cannot explain such a better outcome. Indeed, had the transition from a high- into a low-volatility environment been merely the result of tighter policy regimes, excess returns would have shifted downward. But such prediction is inconsistent with the empirical evidence for the United States of higher expected excess returns experienced from the early 1980s onward. On the other hand, favorable shifts in the variance of nonpolicy innovations imply movements in expected returns that are at least qualitatively consistent with historical patterns. Expected returns are derived here by employing the regression analysis of Cochrane and Piazzesi (2005), with the general idea of using estimates of risk premia extracted *outside the model* to discriminate between alternative model specifications.[2]

This paper builds in particular on the idea in Rudebusch and Wu (2007) and in Bikbov and Chernov (2008) that variations in term structure dynamics may shed some light on the nature of changes in the macroeconomic landscape. Rudebusch and Wu (2007) find significant changes in the U.S. term structure around the mid-1980s by employing an affine asset pricing model. This change originates from time variation in the pricing of risk associated with a so-called "level" factor, namely a factor that affects yields broadly uniformly at all maturities. In turn, this time-varying price of risk is interpreted in terms of agents' beliefs about central bank's inflation objectives. One important limitation of their contribution is that the interpretation of their reduced-form analysis in structural terms is clearly problematic. In contrast, our setup provides a laboratory for an internally consistent exercise. On the other hand, Bikbov and Chernov (2008) consider a rational expectations model with regime switching both in the variances of exogenous shocks and in the monetary policy regimes. Term structure information is employed to improve the identification of those regimes. Our assessment differs from theirs in two main dimensions. First, and contrary to Bikbov and Chernov (2008), we consider a fully microfounded model, which by construction is explicit about the deep sources of change in the economy, as well as about their impact on macroeconomic variables and term premia. The fact that the sources of changes considered here are hardly distinguishable from the VAR representation of the DSGE model's solution makes a reduced-form approach unsuitable, and in contrast vindicates our structural approach. Second, Bikbov and Chernov (2008) find that the yield curve is informative for identifying regime switching in the monetary policy but not in the variance of shocks. Our intuition is that both sources of changes manifest themselves in the term premia by affecting the (conditional) covariances between current and future values of the stochastic discount factor. We discriminate between these two sources of changes in the model economy by comparing the predicted response of term premia with information extracted *outside the model,* in the form of expected excess holding period returns of U.S. government bonds.

Finally, in its emphasis on the sources of changes in the macroeconomic environment, the present analysis is also related to the recent contributions by Davig and Doh (2009) and Bianchi (2012). However, the latter both abstract from the informational content of the term structure. Bianchi (2012) focuses specifically on the role of agents' beliefs in macroeconomic dynamics in an estimated medium-scale DSGE model that allows regime changes in structural parameters and stochastic volatilities.

More broadly, the present analysis provides a contribution to a growing literature that attempts to model the dynamics of macroeconomic variables and bond yields jointly within structural DSGE models.[3] For instance, Nimark (2008) considers the situation in which term structure information is used by agents *inside the model*, notably the central bank, to make inferences about the state of the economy. Alternatively, a number of studies include term structure data in the information set of the econometrician *outside the model*. For example, Hördahl et al. (2006), Doh (2007), Amisano and Tristani (2010), and Bekaert et al. (2010) augment the set of standard macroeconomic observables to include bond price data when estimating versions of the New Keynesian model complemented with affine term structure specifications.[4] There are at least two main purposes in doing so. First, it is a way to evaluate the ability of standard DSGE models to capture the joint dynamics of macroeconomic variables and bond prices. Second, it contributes to sharpening parameter estimates, typically poorly identified in DSGE models' estimations.[5]

The rest of the paper is organized as follows. Section 2 briefly describes the modelling framework and motivates the paper. Section 3 incorporates the term structure into the analysis and derives the model implications. After the robustness analysis performed in Section 4, Section 5 makes use of the empirical evidence on U.S. term premia around the mid-1980s to draw inference on the sources of economic transformations experienced by the U.S. economy in that period. Section 6 concludes.

2. INTERPRETING CHANGES IN THE ECONOMY WITHOUT TERM STRUCTURE INFORMATION

This section outlines the baseline model, which builds on the standard New Keynesian framework originally introduced by Calvo (1983) and extensively reviewed by Woodford (2003).[6] Specifically, the model considered here takes the form of a Markov-switching rational expectations model of the type popularized by Farmer et al. (2011). Regime switching is assumed to govern the processes for (i) the variance of shocks and (ii) the systematic component of monetary policy. Notably, these two sources of change have represented alternative explanations for time-varying features of the economic environment experienced by major industrial economies in the 1980s and 1990s. Model specifications similar to the one employed here have been extensively used for assessing the relative contributions of changes in the variance of shocks and in the systematic component of monetary policy. However, these sources of change are in fact hardly distinguishable by merely looking at the baseline model dynamics, as briefly illustrated in the final part of this section by drawing on the contributions of Benati and Surico (2009) and Davig and Doh (2009).

2.1. The Baseline Model

The model consists of an intertemporal IS equation (1) and an expectations-augmented aggregate supply equation (2), which are derived by log-linear approximating optimal behavior of households and firms. For present purposes, we consider the following specification:

$$\widehat{y}_t = \gamma E_t \widehat{y}_{t+1} + (1-\gamma)\widehat{y}_{t-1} - \sigma^{-1}(\widehat{i}_t - E_t \widehat{\pi}_{t+1}) + \varepsilon_{y,t}, \tag{1}$$

$$\widehat{\pi}_t = \frac{\beta}{1+\alpha\beta} E_t \widehat{\pi}_{t+1} + \frac{\alpha}{1+\alpha\beta}\widehat{\pi}_{t-1} + k\widehat{y}_t + \varepsilon_{\pi,t}, \tag{2}$$

where $\widehat{y}_t$ is the output gap, $\widehat{\pi}_t$ the inflation rate, and $\widehat{i}_t$ the short-term nominal interest rate, all expressed in terms of deviations from their respective steady-state levels, and $\varepsilon_{y,t}$ and $\varepsilon_{\pi,t}$ are demand and cost-push shocks, respectively. The central bank is assumed to set the short-term nominal interest rate according to a Taylor-type rule,

$$\widehat{i}_t = \rho \widehat{i}_{t-1} + (1-\rho)[\phi_\pi(\xi_t^m)\widehat{\pi}_t + \phi_y(\xi_t^m)\widehat{y}_t] + \varepsilon_{i,t}, \tag{3}$$

where ξ_t^m is a Markov chain variable switching between two states intended to capture alternative monetary policy regimes. The transition matrix P_m collects the probabilities $p_{ij}^m \equiv \text{Prob}(\xi_t^m = i | \xi_{t-1}^m = j)$. Macroeconomic disturbances follow exogenous first-order autoregressive processes subject to switches in their conditional variances,

$$\varepsilon_{z,t} = \rho_z \varepsilon_{z,t-1} + \sigma_z(\xi_t^s)\epsilon_{zt} \quad for \quad z = i, \pi, y, \tag{4}$$

where ξ_t^S is a Markov chain variable that governs regime switching in the volatility of exogenous shocks, and ϵ_{zt} are i.i.d. innovations normally distributed with mean zero and unit variance. In essence, this specification can be recast into the generalized form analyzed by Farmer et al. (2011),

$$A(\xi_t)\widehat{s}_t = B(\xi_t)\widehat{s}_{t-1} + \Psi(\xi_t)\epsilon_t + \Pi\eta_t,$$

where $\widehat{s}_t = [\widehat{i}_t \ \widehat{\pi}_t \ \widehat{y}_t \ E_t\widehat{\pi}_{t+1} \ E_t\widehat{y}_{t+1} \ \varepsilon_{i,t} \ \varepsilon_{\pi,t} \ \varepsilon_{y,t}]$, $\epsilon_t = \left[\epsilon_{i,t} \ \epsilon_{\pi,t} \ \epsilon_{y,t}\right]$, and $\xi_t = [\xi_t^m, \xi_t^s]$. Following closely Farmer et al. (2011), the associated MSV solution, provided it exists, has the form

$$\widehat{s}_t = \Theta_1(\xi_t)\widehat{s}_{t-1} + \Theta_0(\xi_t)\epsilon_t.$$

The solution to the structural model can also be expressed as a vector autoregression with regime switching, of the kind popularized by Sims and Zha (2006)

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and Sims at al. (2008), building on the seminal work of Hamilton (1989) in a univariate context. Specifically, the Markov-switching VAR representation for the standard macroeconomic variables $\widehat{i}_t$, $\widehat{\pi}_t$, and $\widehat{y}_t$, collected in the vector $\widehat{x}_t$ together with their lagged values, can be written in companion form as

$$\begin{aligned} \widehat{x}_t &= \Phi(\xi_t)\widehat{x}_{t-1} + \Gamma(\xi_t)\epsilon_t, \\ \epsilon_t &\sim N[0, \Sigma(\xi_t)\Sigma(\xi_t)']. \end{aligned} \tag{5}$$

For ease of exposition, equation (5) can also be rearranged as

$$\widehat{x}_t = \Phi(\xi_t)\widehat{x}_{t-1} + \Gamma(\xi_t)\epsilon_t = \Phi(\xi_t)\widehat{x}_{t-1} + \Gamma(\xi_t)\Sigma(\xi_t)u_t = \Phi\widehat{x}_{t-1} + v_t, \tag{6}$$

where the reduced-form innovation covariance matrix is equal to

$$\mathrm{Var}(v_t|I_t) = \widetilde{\Sigma}_{\xi_t}\widetilde{\Sigma}'_{\xi_t},$$

where $\widetilde{\Sigma} \equiv \Gamma(\xi_t)\Sigma(\xi_t)$.

2.2. Alternative Sources of Change in the Model Economy

The present analysis focuses on two alternative sources of change in the macroeconomic environment, namely (i) shifts in the variance of nonpolicy shock regimes; (ii) changes in the systematic component of monetary policy, and more precisely in the response coefficients. A vast literature has used models similar to the one briefly described earlier to discriminate among these sources of change. However, as illustrated for instance in Benati and Surico (2009) and Davig and Doh (2009), shifts in the variance of nonpolicy shock regimes and changes in the monetary policy coefficients are in fact hardly distinguishable by merely looking at the model's dynamics implied by (5). First, following Benati and Surico (2009), for instance, as both policy coefficients and shock volatilities affect $\widetilde{\Sigma}$ (via Γ and Σ, respectively), observed variations in $\widetilde{\Sigma}$ are in principle compatible with both sources of change in the economy. Admittedly, the two alternatives are not strictly speaking observationally equivalent, primarily because changes in the response coefficients do also impact Φ. In practice, however, they are hardly distinguishable, as can be inferred from the estimation exercise conducted by Benati and Surico (2009). Indeed, changes in Φ associated with monetary policy under determinacy and indeterminacy turn out to imply only modest differences in the way macroeconomic variables respond to shocks over time. It is even more difficult to discriminate among the sets of impulse responses under the two monetary policy regimes once uncertainty around the median responses is accounted for. Second, following Davig and Doh (2009) in their investigation of historical patterns of inflation persistence for the U.S. economy, both monetary and volatility regimes are shown to affect the model-consistent autocorrelation of inflation. Indeed, following their line of argument, and defining $\widehat{\pi}_t = e'\widehat{x}_t$, where e is a 6×1 vector

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with a 1 in the entry associated with $\widehat{\pi}_t$, the first-order autocorrelation of inflation can be expressed as

$$\rho_\pi = e'\Phi C_x(0)e[e'C_x(0)e]^{-1}$$

where $C_x(0)$ is the stationary variance matrix of $\widehat{x}_t$ given by

$$C_x(0) \equiv E_t[\widehat{x}_t\widehat{x}_t] = \Phi C_x(0)\Phi' + \Gamma\Sigma\Sigma'\Gamma'.$$

The serial correlation of inflation ρ_π is a function of the persistence parameters associated with endogenous and exogenous variables, collected into Φ, as well as of the variance matrix $C_x(0)$. Therefore, ρ_π is equally affected by changes in policy coefficients and in shock volatilities, with the former entering in Φ and Γ, and the latter via Σ.

More broadly, that such major macroeconomic transformations are hardly distinguishable through the lens of standard DSGE models warns against simplistic interpretations of reduced-form evidence in structural terms, and ultimately casts some doubt on the usefulness of these models for interpreting the economic landscape and informing policy decisions.

3. INTERPRETING CHANGES IN THE ECONOMY WITH TERM STRUCTURE INFORMATION

A natural way to address the concerns described in the preceding is to consider additional channels along which the two alternative sources of change might manifest themselves differently. With this in mind, the term structure is incorporated into the analysis; as both shock variances and policy coefficients affect bond prices, the yield curve might in principle provide useful information shedding some light on the nature of change in the model economy. Following closely Wu (2006), Bekaert et al. (2010), and Nimark (2008), among others, the term structure is derived by log-normalizing the model's Euler equation. When this is done, the main implication is that the compensation for risk enters into the bond pricing equation. At the same time, the system of linear equations described in the preceding section continues to characterize the dynamics of standard macroeconomic variables. Such a log-linear lognormal modeling approach represents a compromise between the idea of performing a "within-the-model" exercise and the intention of capturing basic aspects of the term structure. Specifically, the idea here is to investigate whether, within a relatively standard modeling framework, additional channels emerge through which the sources of change described previously might manifest themselves differently. The yield curve appears a natural candidate for that purpose. At the same time, simply (log-)linearizing the bond pricing equation would mean ignoring key aspects of the term structure, and notably the compensation for risk. Via log normalization instead, risk consideration has a role to play in the determination of bond yields. Overall, by providing an additional channel through which variances of shocks and policy coefficients enter into the model, the term structure can in principle serve the role of an identification device. The extent to which this is the case in practice will be investigated in the following sections.

3.1. The Model-Consistent Term Structure

Nominal bonds in the economy are priced via the standard equation,

$$P_t^{(n)}(\xi_t) = E_t[M_{t+1} P_{t+1}^{(n-1)}(\xi_{t+1})], \tag{7}$$

where $P_t^{(n)}(s_t)$ is the price of a nominal zero-coupon bond at time t with n periods to maturity and conditional on the regime in place at time t, s_t; M_{t+1} is the nominal stochastic discount factor that satisfies the standard condition

$$M_{t+1}(\xi_{t+1}) \equiv \beta \frac{U_{ct+1}(\xi_{t+1}) P_t(\xi_t)}{U_{ct}(\xi_t) P_{t+1}(\xi_{t+1})}.$$

By taking a lognormal approximation of equation (7), the pricing equation can be expressed as

$$p_t^{(n)}(h) = \sum_j p_{hj} \times \left(E_t[m_{t+1} + p_{t+1}^{(n-1)}(j)|\xi_{t+1} = j] + \frac{1}{2}\mathrm{Var}_t[m_{t+1} + p_{t+1}^{(n-1)}(j)|\xi_{t+1} = j] \right), \tag{8}$$

where $m_{t+1} \equiv \log(M_{t+1})$ and $p_t^{(n)}(h) \equiv \log(P_{t+1}(h))$. Via simple manipulations, the stochastic discount factor consistent with the model's Euler equation can be expressed as

$$m_{t+1} = -(\delta' \widehat{s_t} + \bar{i}) - 0.5\lambda(\xi_{t+1})' \Sigma(\xi_{t+1}) \Sigma(\xi_{t+1})' \lambda(\xi_{t+1}) - \lambda(\xi_{t+1})' \epsilon_{t+1}, \tag{9}$$

where $\lambda(\xi_{t+1})$ is the regime-dependent vector of prices of risk restricted from the structural parameters of the households' FOC. The nominal interest rate i_t is derived from the policy rule equation $i_t \equiv \widehat{i_t} + \bar{i} = \delta' \widehat{s_t} + \bar{i}$. Using equations (8) and (9) and the law of motion of endogenous variables (5), bond prices, and thus yields, are affine functions of macroeconomic variables:

$$p_t^{(n)}(h) = \widetilde{a}_n(h) + \widetilde{b}_n(h) \widehat{s_t}, \tag{10}$$

$$i_t^{(n)}(h) = -\frac{1}{n} p_t^{(n)}(h) = -\frac{\widetilde{a}_n(h)}{n} - \frac{\widetilde{b}_n(h)}{n} \widehat{s_t} = a_n(h) + b_n(h) \widehat{s},_t ,$$

where

$$\widetilde{a}_n(h) = \sum_j p_{hj} [\widetilde{a}_{n-1}(j) - \bar{i} + 0.5 \widetilde{b}'_{n-1}(j) \Gamma(j) \Sigma(j) \Sigma(j)' \Gamma(j)' \widetilde{b}_{n-1}(j) - \widetilde{b}'_{n-1}(j) \Gamma(j) \Sigma(j) \Sigma(j)' \lambda(j)],$$

$$\widetilde{b}_n(h) = \sum_j p_{hj} [-\delta' + \widetilde{b}'_{n-1}(j) \Phi(j)]',$$

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where h is the state of the regime in place at time t. Throughout the paper, we refer to term premia as the expected excess holding-period returns, defined as the (expected) returns on buying an n-period bond at time t and selling it in period $t+1$ in excess of the risk-free short rate. Via simple manipulations, expected excess returns, defined as $E_t \vartheta_{t+1}^{(n)}(h) \equiv E_t[p_{t+1}^{(n-1)} - p_t^{(n)}(h) - i_t]$, can then be expressed as

$$\begin{aligned} E_t \vartheta_{t+1}^{(n)}(h) &= \sum_j p_{hj} \\ &\quad \times \left\{ -\mathrm{Cov}_t[m_{t+1}, p_{t+1}^{(n-1)}(j) | \xi_{t+1} = j] - \frac{1}{2}\mathrm{Var}_t[p_{t+1}^{(n-1)}(j) | \xi_{t+1} = j] \right\} \qquad (11) \\ &= \sum_j p_{hj} \\ &\quad \times \left\{ -\mathrm{Cov}_t(m_{t+1}, E_{t+1} \sum_{i=1}^{n-1} m_{t+1+i} | \xi_{t+1} = j) - \frac{1}{2}\mathrm{Var}_t[p_{t+1}^{(n-1)}(j) | \xi_{t+1} = j] \right\}. \end{aligned}$$

When the model notation is used, expected returns take the following form:

$$\begin{aligned} E_t \vartheta_{t+1}^{(n)}(h) &= \sum_j p_{hj} \left[\lambda'(j)\Sigma(j)\Sigma(j)'\Gamma(j)'\widetilde{b}_{n-1}(j) \right. \\ &\quad \left. - 0.5\widetilde{b}_{n-1}'(j)\Gamma(j)\Sigma(j)\Sigma(j)'\Gamma(j)'\widetilde{b}_{n-1}(j) \right]. \qquad (12) \end{aligned}$$

Specifically, the compensation for carrying certain units of risk is captured by the conditional covariance between bond prices and the pricing kernel.[7] The conditional variance of bond prices is simply due to Jensen's inequality and it is negligible. As in any consumption-based asset pricing model, a negative covariance between long bond prices and the pricing kernel translates into positive risk premia. Intuitively, when carrying low payoffs (low value of $p_{t+1}^{(n-1)}$) when these are valued more (m_{t+1} is high), long bonds fail to provide insurance when needed, and hence command positive risk premia. Finally, note how both the variance of shocks and the policy coefficients affect term premia: the former via $\Sigma(\xi_{t+1})\Sigma(\xi_{t+1})'$, and the latter via both $\Gamma(\xi_{t+1})'\widetilde{b}_{n-1}(\xi_{t+1})$ and $\lambda(\xi_{t+1})$.

3.2. The Response of the Term Structure to Changes in the Model Economy

This section investigates the response of term premia to the previously mentioned sources of change in the model economy. The structural parameters are calibrated as follows. Most parameters are calibrated by estimating a constant-parameter specification of the model described in Section 2.1. In essence, by spanning the period of macroeconomic stability under the Volcker and Greenspan chairmanships, this estimation provides the calibration for the regime of low macroeconomic

TABLE 1. Bayesian estimation of the model parameters

Parameter	Domain	Density	Prior distribution: Mode	Prior distribution: Standard deviation	Posterior distribution: Median and 90% coverage percentile
k	$\mathbf{R}^+$	Gamma	0.05	0.01	0.04 [0.03; 0.06]
σ	$\mathbf{R}^+$	Gamma	2.00	1.00	7.78 [5.87; 10.25]
α	[0, 1]	Beta	0.20	0.20	0.24 [0.07; 0.48]
γ	[0, 1]	Beta	0.95	0.10	0.84 [0.73; 0.92]
ρ	[0, 1]	Beta	0.75	0.20	0.69 [0.62; 0.76]
ϕ_π	$\mathbf{R}^+$	Gamma	1.00	0.50	2.11 [1.50; 2.78]
ϕ_y	$\mathbf{R}^+$	Gamma	0.15	0.25	0.81 [0.44; 1.30]
ρ_i	[0, 1)	Beta	0.3	0.10	0.30 [0.23; 0.38]
ρ_π	[0, 1)	Beta	0.3	0.10	0.32 [0.24; 0.41]
ρ_y	[0, 1)	Beta	0.3	0.10	0.64 [0.57; 0.70]
$\sigma^2_{\varepsilon i}$	$\mathbf{R}^+$	InvGamm	0.5	1.00	1.08 [0.86; 1.40]
$\sigma^2_{\varepsilon \pi}$	$\mathbf{R}^+$	InvGamm	0.9	0.80	0.33 [0.25; 0.46]
$\sigma^2_{\varepsilon y}$	$\mathbf{R}^+$	InvGamm	0.5	0.50	0.10 [0.06; 0.16]

volatility. Specifically, the sample period ranges from 1980Q1 to July 2007, and the macro data used in the estimation are inflation (annualized quarter-to-quarter percentage change of GDP deflator), output gap (percentage deviation of real GDP from its potential, where the latter is the CBO estimate), and nominal interest rate (annualized Federal Funds Rate). Table 1 presents the priors and the posterior estimates of the model parameters, where the latter are derived by using first a simulated annealing algorithm to maximize the log posterior, and second a random walk Metropolis algorithm to draw from the posterior distribution. The parameters specifically associated with the regime of high macroeconomic volatility are calibrated as follows. High values of variance of nonpolicy shock, $\sigma_{\varepsilon\pi}$ and $\sigma_{\varepsilon y}$, are calibrated by scaling up the posterior median estimates by a factor of 1.1. The switch of monetary policy from a "tight" (hawkish) to a "loose" (dovish) regime is captured by assuming that the policy coefficients ϕ_π and ϕ_y are brought down from the posterior median estimates, where the latter represent the "tight" (hawkish) regime. Specifically, under the loose regime, the policy coefficients ϕ_π and ϕ_y are calibrated at 1.0 and 0.4, respectively, and namely at values for which the model solution is close to the boundary with the indeterminacy region.[8] Finally, the probabilities of persisting in a given regime, p^k_{ii}, for $k = m, s$, are calibrated symmetrically at 0.9, implying an average duration for a given state of 10 quarters.

Regime switching in the nonpolicy shock variances. Figure 1 illustrates how a switch to a regime of lower nonpolicy shock variance is associated with higher term premia at all maturities. In essence, under this scenario, the assumption is to consider a regime switch only in the process governing the variance of

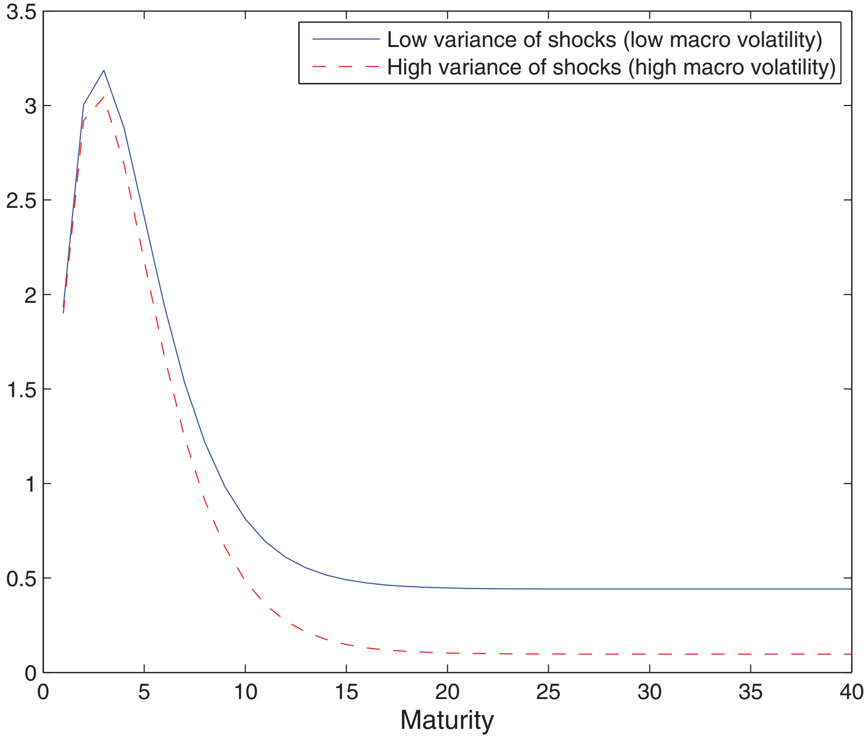

FIGURE 1. Expected excess returns under low and high nonpolicy shock variance regimes.

nonpolicy shock, whereas the remaining parameters are kept fixed at their posterior estimates.[9]

The mechanism underpinning our finding closely relates to the implications of any standard consumption-based asset pricing model. Term premia are positive if marginal utility is negatively correlated with expected changes in future marginal utilities. In this case, long bonds commands a premium over the risk-free short rate because their expected payof ($E_{t+1} p_{t+1}$) is low when most needed (m_{t+1} is high). Alternatively, a positive covariance of expected future bond prices with m_{t+1} means that long bonds are attractive assets, as their payoffs tend to be high when most valued (m_{t+1} is high). In this case, long bonds command a negative premium over short rates. Central to our finding is that higher variances of demand and cost-push shocks lead to larger positive autocovariances of both real output and inflation. This induces m_{t+1} to covary more positively with its expected future values, thus leading to lower expected returns, as evident from (11). Intuitively, by having higher returns when most needed, long bonds provide insurance against bad times and thus command lower premia.

To see this better, consider the following arguments. First, assume the simplified case $\sigma = \gamma = 1$, and for convenience a constant-regime environment; in this case

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m_{t+1} takes the simple form

$$m_{t+1} = \overline{m} - y_{t+1} + y_t - \pi_{t+1} - \varepsilon_{y,t}.$$

Second, focusing on two-period maturity, excess returns are given by[10]

$$\begin{aligned} E_t \vartheta_{t+1}^{(2)} &= -\operatorname{Cov}_t(m_{t+1}, E_{t+1} p_{t+1}) = -\operatorname{Cov}_t(m_{t+1}, E_{t+1} m_{t+2}) \\ &= -\operatorname{Cov}_t(\Delta y_{t+1}, \Delta y_{t+2}) - \operatorname{Cov}_t(\Delta y_{t+1}, \varepsilon_{y,t+1}) - \operatorname{Cov}_t(\Delta y_{t+1}, \pi_{t+2}) \\ &\quad - \operatorname{Cov}_t(\pi_{t+1}, \Delta y_{t+2}) - \operatorname{Cov}_t(\pi_{t+1}, \pi_{t+2}) - \operatorname{Cov}_t(\pi_{t+1}, \varepsilon_{y,t+1}). \end{aligned}$$

Third, ignoring for a moment the cross-covariance terms, excess returns simplify further to

$$\begin{aligned} E_t \vartheta_{t+1}^{(2)} &= -\operatorname{Cov}_t[m_{t+1}, E_{t+1} p_{t+1}] = -\operatorname{Cov}_t[m_{t+1}, E_{t+1} m_{t+2}] \\ &\approx -\operatorname{Cov}_t[\Delta y_{t+1}, \Delta y_{t+2}] - \operatorname{Cov}_t[\pi_{t+1}, \pi_{t+2}]. \end{aligned}$$

From the last equation, it is then evident how larger variances of nonpolicy shocks weigh down on term premia, by making real output growth and inflation covary more positively with their respective future values. Quantitatively these are the crucial forces behind our finding. When the cross-covariance terms are also considered, those including $\varepsilon_{y,t+1}$ similarly imply that higher variance of demand shocks weighs negatively on term premia, as it induce a greater positive covariance between demand shocks and inflation and output (growth). Admittedly, these simplified lines of explanation are somewhat lost when more general cases are considered in which σ and γ are different from 1, as additional terms enter into the definition of term premia. However, for a wide range of alternative parameter calibrations, the rises in the autocovariance of inflation and output gap associated with higher variance of nonpolicy shocks remain quantitatively the prevailing forces behind our finding.

To gain further formal insight, simple manipulations of (12) lead to the following alternative expression for excess returns under the time-t regime $s_t = h$:

$$\begin{aligned} E_t \vartheta_{t+1}^{(n)}(h) = \sum_j p_{hj} \big[&\lambda_i \sigma_{\varepsilon i}^2 (\Gamma'_{.1} \widetilde{b}_{n-1}) + \lambda_\pi \sigma_{\varepsilon\pi}^2(j)(\Gamma'_{.2} \widetilde{b}_{n-1}) \\ &+ \lambda_y \sigma_{\varepsilon y}^2(j)(\Gamma'_{.3} \widetilde{b}_{n-1}) \big], \end{aligned} \quad \mathbf{(13)}$$

where λ_k is the kth entry of the vector of prices of risk λ, and $\Gamma'_{.k}$ is the (transpose of the) kth column vector of the matrix Γ, the vector that captures the impact of the kth macroeconomic shock. In other words, (13) expresses excess returns at different maturities in terms of the contributions of monetary, cost-push, and demand shocks, respectively. The difference between excess returns under the low (l) and the high (h) nonpolicy shock variance regimes can then be expressed as

$$E_t\vartheta_{t+1}^{(n)}(l) - E_t\vartheta_{t+1}^{(n)}(h) = \lambda_\pi(\Gamma'_{.2}\widetilde{b}_{n-1})[\widetilde{\sigma}_{\varepsilon\pi}^2(l) - \widetilde{\sigma}_{\varepsilon\pi}^2(h)]$$
$$+ \lambda_y(\Gamma'_{.3}\widetilde{b}_{n-1})[\widetilde{\sigma}_{\varepsilon y}^2(l) - \widetilde{\sigma}_{\varepsilon y}^2(h)], \tag{14}$$

where $\widetilde{\sigma}_\varepsilon^2(l) \equiv (p_{ll} - p_{hl})\sigma_\varepsilon^2(l)$, $\widetilde{\sigma}_\varepsilon^2(h) \equiv (p_{hh} - p_{lh})\sigma_\varepsilon^2(h)$, and $\sigma(l)_{\varepsilon k}$ and $\sigma(h)_{\varepsilon k}$ denote low and high standard deviations of shock k, respectively. Both additive terms in equation (14) are shown to be positive, similarly implying that lower shock volatility regimes induce higher term premia. To shed some light on this, it is convenient to consider each of the three factors of the two additive terms in turn. First, both λ_π and λ_y are positive, reflecting positive risk premia on financial assets carrying one unit of risk associated with cost-push and demand shocks, respectively. As evident from equation (9), a positive λ_π means that m_{t+1} falls in response to ε_π. Intuitively, a positive cost-push shock causes an increase in inflation and a contraction in output as the central bank raises the real interest rate in response to such inflationary pressures. With inflation increasing and output decreasing, the final effect on the nominal pricing kernel is in principle ambiguous. In practice, under a wide range of alternative calibrations, the nominal pricing kernel tends to fall, and λ_π is positive.[11]Similarly, λ_y is also positive. Intuitively, a demand shock brings about an increase in both inflation and output that unambiguously leads to a fall in the nominal pricing kernel, thus implying a positive market price of risk λ_y. Second, both $(\Gamma'_{.2}\widetilde{b}_{n-1,.})$and $(\Gamma'_{.3}\widetilde{b}_{n-1,.})$ are negative, reflecting the fall in bond prices in response to cost-push and demand shocks. Intuitively, both shocks call for a rise in the short-term policy rate, which, by transmitting itself across the term structure, leads to a fall in long bond prices.[12]

With λ_π and λ_y positive, $(\Gamma'_{.2}\widetilde{b}_{n-1,.})$and $(\Gamma'_{.3}\widetilde{b}_{n-1,.})$ negative, and $(\widetilde{\sigma}_{\varepsilon_k}^2(l) - \widetilde{\sigma}_{\varepsilon_k}^2(h))$ negative for $k = \pi, y$, both multiplicative terms in (14) are positive. As a result, excess returns are unambiguously larger under smaller variances of nonpolicy shocks.[13] All in all, the first model's prediction can be summarized as follows: moving to a more stable macroeconomic environment would bring about an upward shift in term premia when such improved stability resulted from lower variance of nonpolicy structural disturbances.

Regime switching in the monetary policy conduct. In contrast, when improved macroeconomic stability is achieved via a tighter monetary policy, and here is the second model's prediction, excess returns would actually fall. This is illustrated in Figure 2 by displaying that excess returns under the tight policy regime are lower than under the loose regime at all maturities (horizontal axis).

To isolate the impact of changing monetary policy on excess returns, regime switching is assumed to characterize the dynamic of policy coefficients only, whereas all the other parameters, including the standard deviations of various shocks, are kept fixed at their posterior estimates. Intuitively, by better insulating inflation and output from various shocks, tighter regimes make the pricing kernel less responsive to macroeconomic disturbances and ultimately less negatively

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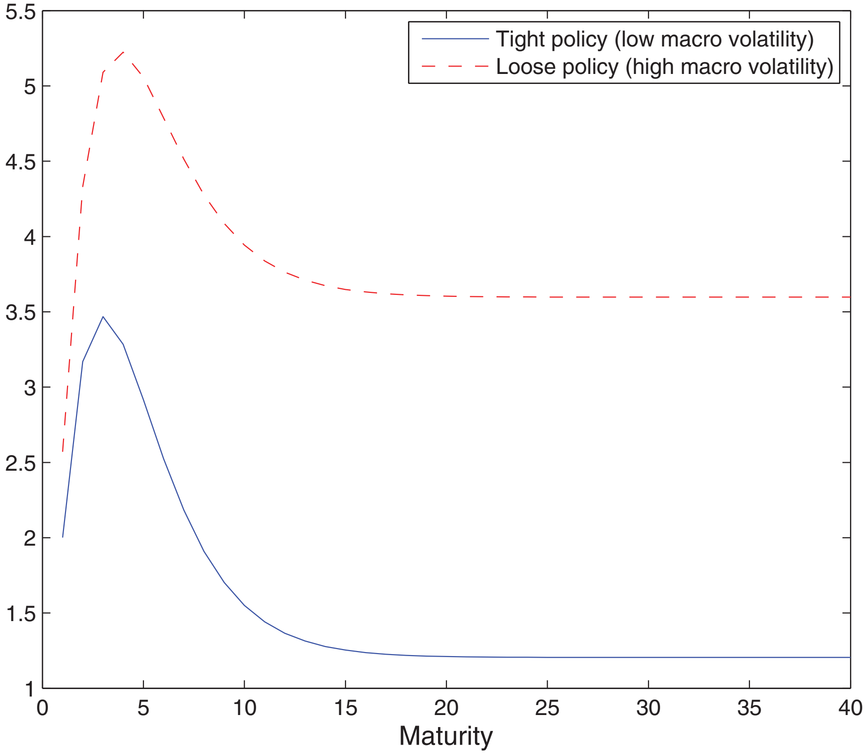

FIGURE 2. Expected excess returns under tight and loose monetary policy regimes.

correlated with long bond prices. As a result, long bonds command lower premia under tighter regimes. In essence, these regimes make inflation less negatively correlated with output growth, and thus long bonds less risky. With only policy coefficients subject to regime switching, excess returns can be expressed as

$$E_t\vartheta_{t+1}^{(n)}(h) = \sum_j p_{hj}\left[\lambda'(j)\Sigma\Sigma'\Gamma'(j)\widetilde{b}_{n-1}(j) - 0.5\widetilde{b}'_{n-1}(j)\Gamma(j)\Sigma\Sigma'\Gamma'(j)\widetilde{b}_{n-1}(j)\right],$$

where the regime h refers to the policy coefficient regime at time t. Again, to gain further insight, it is useful to consider the difference between excess returns under the tight (h) and the loose (l) regime:

$$\begin{aligned} E_t\vartheta_{t+1}^{(n)}(h) - E_t\vartheta_{t+1}^{(n)}(l) &= \sigma_{\varepsilon i}^2\{\widetilde{\lambda}_i(h)[\Gamma'_{.1}(h)\widetilde{b}_{n-1,.}(h)] - \widetilde{\lambda}_i(l)[\Gamma'_{.1}(l)\widetilde{b}_{n-1,.}(l)]\} \\ &\quad + \sigma_{\varepsilon\pi}^2\{\widetilde{\lambda}_\pi(h)[\Gamma'_{.2}(h)\widetilde{b}_{n-1,.}(h)] - \widetilde{\lambda}_\pi(l)[\Gamma'_{.2}(l)\widetilde{b}_{n-1,.}(l)]\} \\ &\quad + \sigma_{\varepsilon y}^2\{\widetilde{\lambda}_y(h)[\Gamma'_{.3}(h)\widetilde{b}_{n-1,.}(h)] - \widetilde{\lambda}_y(l)[\Gamma'_{.3}(l)\widetilde{b}_{n-1,.}(l)]\}, \end{aligned} \quad (\mathbf{15})$$

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where $\widetilde{\lambda}(h) \equiv (p_{hh} - p_{lh})\lambda(h)$, $\widetilde{\lambda}(l) \equiv (p_{ll} - p_{hl})\lambda(l)$ and the term $\lambda_k(j)(\Gamma'_{.k}(j)\widetilde{b}_{n-1,.}(j)$ captures the negative covariance between the pricing kernel and bond prices induced by the kth shock of the vector ϵ_t under the policy regime $j = h, l$.

Under the assumption of a symmetric transition probability matrix, the previous expression simplifies further to

$$
\begin{aligned}
&E_t\vartheta^{(n)}_{t+1}(h) - E_t\vartheta^{(n)}_{t+1}(l) \\
&\quad = \widetilde{\sigma}^2_{\varepsilon i}\{\lambda_i(h)[\Gamma'_{.1}(h)\widetilde{b}_{n-1,.}(h)] - \lambda_i(l)[\Gamma'_{.1}(l)\widetilde{b}_{n-1,.}(l)]\} \\
&\qquad + \widetilde{\sigma}^2_{\varepsilon\pi}\{\lambda_\pi(h)[\Gamma'_{.2}(h)\widetilde{b}_{n-1,.}(h)] - \lambda_\pi(l)[\Gamma'_{.2}(l)\widetilde{b}_{n-1,.}(l)]\} \\
&\qquad + \widetilde{\sigma}^2_{\varepsilon y}\{\lambda_y(h)[\Gamma'_{.3}(h)\widetilde{b}_{n-1,.}(h)] - \lambda_y(l)[\Gamma'_{.3}(l)\widetilde{b}_{n-1,.}(l)]\}.
\end{aligned} \tag{16}
$$

The first term in (16) is negative, capturing the fact that the covariance between m_{t+1} and bond prices induced by a monetary policy shock is, in absolute value, lower under the tight regime.[14] Intuitively, a monetary policy shock brings about a contraction in inflation and output under both regimes. However, under the tight regime, the impact of the policy shock on macroeconomic variables is largely mitigated by the more aggressive opposite response of the systematic component of policy. This implies that, under the tight regime, both the pricing kernel and bond prices respond less to a monetary policy shock, and hence tend to covary less in absolute value. The second term in (16) is also negative, reflecting the fact that the covariance between the pricing kernel and bond prices induced by a cost-push shock, and given by $-[\lambda_\pi(\Gamma'_{.2}\widetilde{b}_{n-1,.})]$, is greater under the tight regime. Specifically, although bringing about a similar rise in inflation under both regimes, cost-push shocks prompt an interest rate reaction that leads to a more pronounced fall in output under the tight regime. From one side, this means that the pricing kernel declines less, namely that λ_π is lower under the tight regime. On the other hand, it implies that the tight regime is characterized by a stronger reaction of long rates in the face of inflationary shocks as a result of a more effective transmission of short-term policy rate across the term structure. Although the final effect on the covariance term is in principle ambiguous, in practice the pricing kernel tends to covary more positively with bond prices under the tight regime, thus implying lower premia. The third term in (16) is instead positive and small, under the baseline calibration, meaning that the covariance between the pricing kernel and bond prices induced by a demand shock is lower under the tight regime. Intuitively, under the latter regime, a demand shock leads to a less pronounced fall in the price kernel and to a larger decline in bond prices. Similarly to the case of the cost-push shock, the relative strength of the covariance between the pricing kernel and bond prices in the two regimes is in principle ambiguous. It turns out that the pricing kernel and bond prices covary less positively under the tight regimes, and thus long bonds end up commanding slightly higher premia in the face of demand shocks. All in all, with the first two negative terms dominating the third small positive one, expected excess returns turn out to be lower under tighter policy regimes.

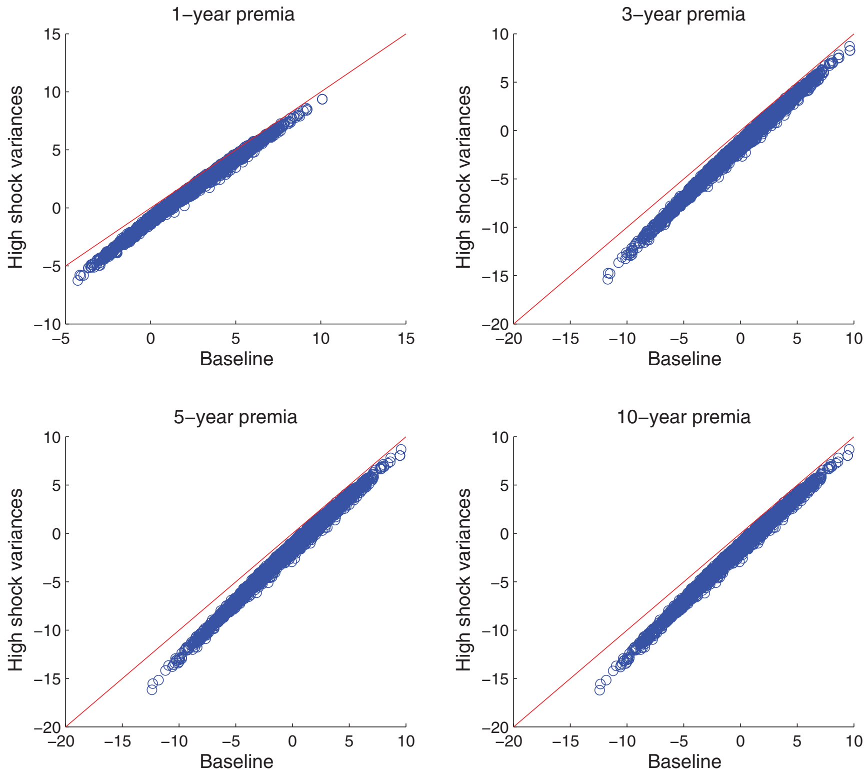

FIGURE 3. Scatterplot between excess returns under the high-shock-variance regime and under the low-macroeconomic-volatility baseline.

4. SENSITIVITY ANALYSIS (VIA BAYESIAN PREDICTIVE CHECKS)

This section confirms that the findings described in the preceding section hold also for a wide range of alternative calibrations of structural parameters. To illustrate this in a systematic manner, parameter values are first drawn from their posterior distribution. Then, for each parameter draw, excess returns are computed under two alternative cases for the high-volatility regime. The first case reflects higher shock variances, where $\sigma^h_{\varepsilon\pi}$ and $\sigma^h_{\varepsilon y}$ are obtained by scaling up the draws for $\sigma_{\varepsilon\pi}$ and $\sigma_{\varepsilon y}$. A second case captures a looser monetary policy regime, characterized by bringing down the drawn values of the policy coefficients at the boundary with the indeterminacy region.

Figure 3 shows the scatterplot between excess returns for selected maturities under the low-volatility regime (horizontal axis) and under the high-volatility regime when the latter stems from high shock variances (vertical axis). The scatterplot depicts 5,000 parameter draws. Points below (above) the 45° line mean that term premia are lower (higher) under the high-shock-variance regime.

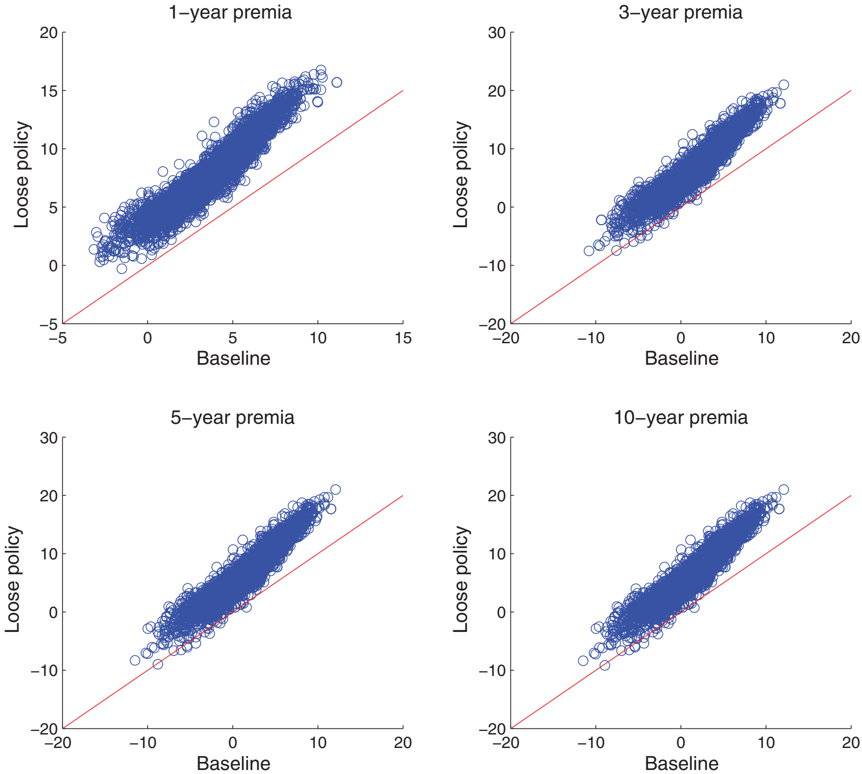

FIGURE 4. Scatterplot between excess returns under the loose-monetary-policy regime and under the low-macroeconomic-volatility baseline.

Our finding of higher excess returns associated with lower shock variances appears very robust, as illustrated graphically by the concentration of all points below the 45° line. In light of the analysis in the preceding section, the robustness of our finding is only partly surprising, considering that scaling up nonpolicy shock variances translates into lower excess returns as long as the market price of risks remain positive.

Turning to the scenario of regime changes in the monetary policy rule, Figure 4 shows the scatterplot between excess returns for selected maturities under the low-volatility regime (horizontal axis) and under the high-volatility regime when the latter stems from a looser monetary policy rule (vertical axis). The finding that tighter regimes are associated with lower excess returns is confirmed, as illustrated by the fact that almost all points are located above the 45° line. This finding appears only slightly less robust than in the previous case, because in just about 1.8% of the cases the second model's prediction is reversed.[15] Yet the overall robustness of the finding is remarkable, taking into account the large

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region of the parameter space spanned by the draws, illustrated indirectly by the wide range of values taken by term premia.

5. THE EXPERIMENT IN A HISTORICAL PERSPECTIVE

Does the comparison between the model's predicted responses and the pattern of U.S. term premia around the mid-1980s shed some light on the sources of the economic transformations experienced by the U.S. economy in that period ? To investigate this issue, after having derived the model's predictions in the last sections, we extract here the time series for expected excess holding period returns of U.S. government bonds by employing the regression analysis by Cochrane and Piazzesi (2005). There are two main reasons underpinning this approach. First, it is a way to use valuable information extracted *outside the model* to discriminate between alternative model specifications. Second, the regression analysis by Cochrane and Piazzesi (2005) appears successful in extracting time series for expected returns that fit model-free ex post excess returns relatively well. Although, in principle, estimates for expected returns could be recovered by directly estimating the structural DSGE model, in practice even much richer model specifications than the one considered here typically fail to generate sizeable time-varying term premia.[16] Moreover, the focus on excess returns as a measure of term premia makes it possible to net out the levels of inflation and of interest rates, which could well be influenced by time variation in the underlying inflation target, or in agents' beliefs about the inflation target, two channels not considered in this framework.[17]

As a result of these considerations, time series for expected returns are here derived by employing the approach of Cochrane and Piazzesi (2005).[18] Specifically, their regression equation is the following:

$$rx_{t+1}^{(n)} = b_n(\gamma_0 + \gamma_1 i_t^{(1)} + \gamma_2 f_t^{(2)} + \ldots \gamma_5 f_t^{(5)}) + \nu_{t+1}^{(n)}, \qquad (\mathbf{17})$$

where $rx_{t+1}^{(n)}$ is the one-year excess return at maturity n, and $f_t^{(n)}$ is the time-t forward rate. As b_n and γ cannot be identified separately, the average value of b_n is normalized to 1. Equation (17) is estimated following a two-step approach. First, the parameters γ are estimated by regressing the average (across maturities) excess return on forward rates:

$$\overline{rx}_{t+1} = \gamma' f_t + \overline{\nu}_{t+1},$$

where $\overline{rx}_{t+1}$ is the average excess return. Second, b_n are estimated by running the regressions

$$rx_{t+1}^{(n)} = b_n(\gamma' f_t) + \nu_{t+1}^{(n)} \quad \text{for} \quad n = 2, 3, 4, 5,$$

where $(\gamma' f_t)$is the single linear combination of forward rates between $t + n - 1$ and $t + n$ that predicts excess returns at all maturities. On the basis of this approach by Cochrane and Piazzesi (2005), one-year holding period returns for

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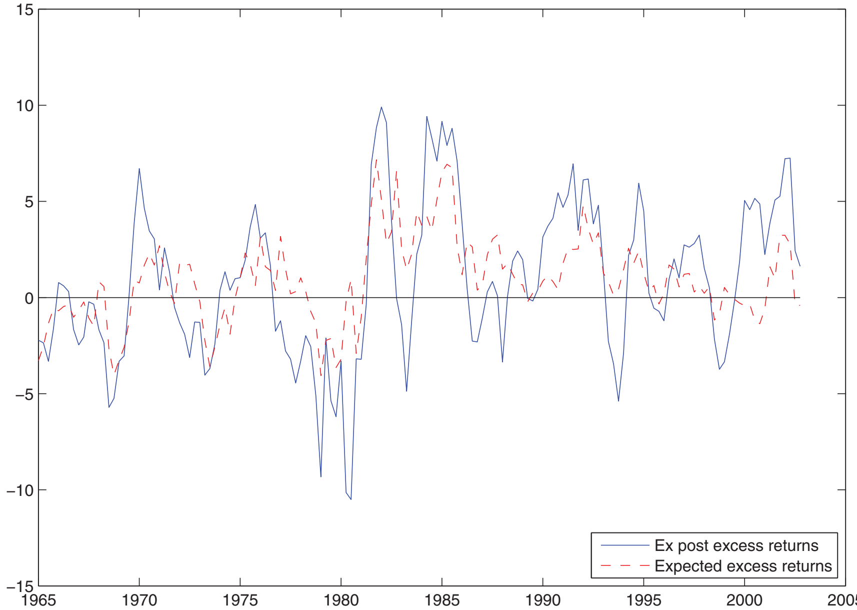

FIGURE 5. Historical paths of U.S. ex post and expected excess returns (averages across maturities). The ex post returns series is shifted to the left to line up with the expectations.

U.S. government bonds with maturities from 2 to 5 years are thus derived for the period 1965Q1–2003Q4. Figure 5 depicts the time series for average (across maturities) expected returns, along with the corresponding ex post excess returns.

Overall, ex post excess returns tend to display substantial fluctuation over time, as captured by the standard deviation having been as high as 4% over the relevant period. Of main interest for present purposes is the extent to which expected returns experienced a regime change around the mid-1980s, in ways that can be informative about the sources of economic transformation experienced by the U.S. economy in that period. Specifically, two distinct subperiods characterizing the post–World War II U.S. macroeconomic history are considered, before and after the (end of the) Volcker disinflation, respectively.

First, over the subperiod from 1965Q1 to 1979Q3, the mean of (average across maturities) ex post and expected excess returns is around −0.8 and −0.2%, respectively. Focusing more specifically on the 1970s, a decade characterized by particularly high macroeconomic instability, the means of ex post and expected excess returns are −0.5 and 0.3%, respectively, whereas the standard deviations are almost unchanged in comparison to the sample period 1965Q1:1979Q3 for both ex post and expected excess returns. Second, over the subperiod from 1984Q1 to 2003Q4, the means of (average across maturities) ex post and expected excess returns are as high as 2.4 and 1.5%, respectively, whereas the standard deviations of both ex post and expected excess returns remain broadly unchanged from

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TABLE 2. Testing for structural breaks in U.S. expected excess returns

Tests						
$\sup F_T(1)$	$\sup F_T(2)$	$\sup F_T(3)$	$\sup F_T(4)$	$\sup F_T(5)$	UDmax	Wdmax
6.73	13.10	9.71	7.68	4.75	13.10	14.69
$F_T(2\|1)$	$F_T(3\|2)$	$F_T(4\|2)$				
13.91	2.96	0.87				
Numbers of breaks identified						
Sequential	2					
BIC	2					
LWZ	2					
Estimates with two breaks						
μ_1	μ_2	μ_3		T_1	T_2	
−0.36	4.04	1.06		81Q1	86Q3	
(0.54)	(0.66)	(0.42)		(79Q4–83Q2)	(83Q3–88Q3)	

Notes: The significance of the tests is at the 5% level. In parentheses, below the estimates, are the standard errors (robust to serial correlation).

the first subperiod. Finally, during the first four years of the Federal Reserve chairmanship of Paul Volcker, which comprises the bulk of the U.S. disinflation, excess returns were extremely volatile, presumably on account of the Federal Reserve experimenting with direct targeting of monetary aggregates, and of the two recessionary episodes that occurred in such a short period of time.

To address more formally the issue of structural changes in expected excess returns series, we employ the approach of Bai and Perron (2003), who largely draw on the theoretical results in Bai and Perron (1998). In essence, this approach allows estimating the number of breaks and the break dates, along with the associated confidence intervals under various hypotheses on the structure of the data and errors. We replicate the empirical analysis in Bai and Perron (2003) for the expected excess returns series derived previously, allowing for up to five structural breaks and accounting for serial correlation in the data. Table 2 reports the results. The $\sup F_T(k)$ tests signal the presence of more than one break, the null hypothesis of no structural break being rejected at the 5% confidence interval for numbers of breaks, k, from 2 up to 5.[19] This finding is confirmed by the results of UDmax and Wdmax, which test the null hypothesis of no structural break against the alternative hypothesis of an unknown number of breaks. Furthermore, the $\sup F_T(k+1|k)$ statistics, which test sequentially k versus $k+1$ of breaks, indicate the presence of two breaks. This finding is also confirmed by the BIC and the modified Schwarz criterion.

As a result of this analysis, expected returns series are estimated on the basis of the following specification. assuming the presence of two structural breaks:

$$\overline{rx}_{t+1} = z_t'\mu_j + u_{t+1} \quad t = T_{j-1}+1, \ldots, T_j,$$

where $j = 1, m + 1$, with m being the total number of breaks, identified as equal to two. As z_t includes the constant as the only regressor, changes in μ_j represent breaks in the mean of expected excess returns, which are thus consistent with the model's prediction about the mean levels of excess returns. Specifically, the procedure jointly estimates the unknown regression coefficients μ_j together with the break points T_j, and allows for serial correlation in the errors u_{t+1}. The results are reported in Table 2. Not surprisingly, the identified break dates define three subsamples, which roughly correspond to the previously investigated subperiods, namely the pre-Volcker period, the Volcker disinflation period, and the post-84 period. Although not statistically different from zero in the first subperiod, the mean of excess returns is estimated to be as high as 4% during the second subperiod, and 1% in the third subperiod. All in all, simple empirical evidence and more formal analysis point to an upward shift in the level of U.S. expected excess returns identified in the early 1980s.

Notably, when this finding is used to interpret the sources of the U.S. improved macroeconomic stability in the 1980s and 1990s through the lens of a small-scale DSGE model, the implication is that changes in monetary policy alone cannot explain this better outcome. Had the transition from a high- into a low-volatility environment been merely the result of tighter policy regimes, expected excess returns would have shifted downward, a fact inconsistent with the empirical evidence of rising expected returns experienced from the early 1980s onward. On the other hand, favorable shifts in the variances of nonpolicy innovations imply movements in expected returns that are at least qualitatively consistent with historical patterns. Finally, this finding is consistent with a number of studies that identify the change in the variance of structural shocks as the major source behind the U.S. economic transformation in the late 1970s and early 1980s, studies that include, for instance, Stock and Watson (2003), Primiceri (2005), and Sims and Zha (2006).

6. CONCLUSIONS

A large literature has increasingly attempted to capture jointly the dynamics of macroeconomic variables and of the term structure using structural general equilibrium models. Indeed, the extent to which these models can be considered suitable candidates for rationalizing consumption and investment decisions ultimately rests on their ability to also capture important features of those markets that are relevant to these economic decisions, and notably the bond markets. Moreover, adding bond price data into the econometric analysis might contribute to mitigating identification issues, particularly severe and widespread in standard DSGE models. In the present context, the term structure serves the role of discriminating among alternative sources of change in the model economy, sources otherwise hardly distinguishable by looking at the dynamics of standard macro variables within small scale DSGE models. These two sources of change are shifts in nonpolicy shock variances and changes in the systematic component of monetary policy, which represent competing ways to account for time-varying features of the economic environment experienced by major industrial economies during the 1980s and 1990s. Although similarly implying a reduction in the macroeconomic volatility, these two alternatives are found to manifest themselves differently in the model-consistent term structure, implying higher and lower term premia, respectively. Therefore, when the sources of improved U.S. macroeconomic stability of the 1980s and 1990s are interpreted in light of these findings, the implication is that a move to tighter monetary regimes alone cannot explain this better outcome, as this would have implied lower expected excess returns, in contrast with the empirical evidence of rising (average) expected excess returns experienced from the early 1980s onward. Favorable shifts in the variance of nonpolicy innovations instead imply movements in expected returns that are at least qualitatively consistent with empirical evidence.

Admittedly, there are two major caveats to our findings. First, changes in monetary policy are here modeled in terms of switches in the response coefficients of the policy rule. Although this represents the most common characterization in the literature, it is not the only one. Schorfheide (2005), for instance, documents time variation in the conduct of U.S. monetary policy in terms of a regime-switching inflation target. More recently, Levin and Taylor (2010) corroborate this view on the basis of the evolution of long-run inflation expectations. At the same time, when considering large DSGE models that include habit formation and various markup shocks, Liu et al. (2011) do not find compelling evidence of changes in the inflation target. Moreover, the impact of changes in the central bank's inflation objective on term premia would not be immediately evident. Second, canonical DSGE models fall short in characterizing an empirically plausible term structure. In particular, these models fail to generate term premia that are as sizeable and variable as those observed in the data, a failure well documented, for instance, by Rudebusch and Swanson (2008). In our context, term premia vary over time as a result of changes in the "amount" and "price" of risk, associated with shifts in shock variance and monetary policy regimes, respectively. When the term structure is solved to second-order approximation, conditional on a given regime, term premia are constant. More recently, Rudebusch and Swanson (2009) have drawn a more positive conclusion regarding the ability of structural models to match macroeconomic and term structure evidence simultaneously: a third-order approximate solution to an otherwise standard DSGE model, augmented with Epstein–Zin preferences as well as with long-run economic risks, implies sizeable and variable term premia, although still preserving a good fit of main macroeconomic variables. However, as estimating models solved to third-order approximation is currently unfeasible, this framework cannot be used to address empirical issues. More recently, Amisano and Tristani (2010) notably find that even small DSGE models solved to a second-order approximation, and estimated on U.S. data, generate non-negligible time-varying risk premia once accounting for stochastic regime shifts; yet their framework falls somewhat short of generating term premia that are as sizeable and variable as those extracted, for instance, from reduced-form approaches such as that of Cochrane and Piazzesi (2005) employed here.

NOTES

1. For a documentation of time variation in overall macroeconomic stability in the United States during the 1980s and 1990s, see, for instance, Kim and Nelson (1999) and McConnell and Perez-Quiroz (2000).

2. In a similar vein, for instance, Del Negro and Eusepi (2011) use inflation expectations obtained from the Survey of Professional Forecasters to discriminate between three variants of a prototypical DSGE model.

3. A complementary approach investigates the interaction between macroeconomic variables and bond yields within reduced-form models. When doing so, Ang and Piazzesi (2003), for instance, find that arbitrage-free VAR models with macro factors forecast better than models with only unobservable factors; moreover, macro factors are able to explain much of the variation in bond yields.

4. Precisely, Bekaert et al. (2010) consider a log-linear lognormal approach using the model-consistent pricing kernel, which then implies constant term premia. Hördahl et al. (2006) assume instead a flexible non-model-consistent specification for the pricing kernel, which induces time variation in term premia (via time variation in the "price of risk"). Doh (2007) and Amisano and Tristani (2010) solve the full model to a second-order approximation, and time-varying risk premia result from heteroskedasticity in the model structural shocks (time variation in the "amount of risk").

5. For an investigation of identification issues in DSGE models, see Canova and Sala (2009), for instance.

6. See also Goodfriend and King (1997) and Clarida et al. (1999).

7. Expected excess returns can be equivalently expressed in terms of the conditional covariance between current and expected future nominal pricing kernels.

8. As will be extensively investigated in the next section, the findings are robust to a wide range of alternative calibrations.

9. For convenience of further exposition, we assume that this situation is one in which the equation (11) takes the simplified form where only $\Sigma(j)$ is a function of the state j of the Markov-switching process, whereas λ and $\Gamma$and $\widetilde{b}_{n-1}$ are not functions of the state j.

10. The convexity term due to Jensen's inequality is disregarded here.

11. The robustness of various findings will be investigated in the next section.

12. Note that, in this framework, the determinacy of the equilibrium calls for a policy reaction that not only increases the nominal short-term rate in the face of inflationary pressures, but also does so aggressively enough to raise real interest rates.

13. A sufficient condition for $\widetilde{\sigma}_{\varepsilon}^2(l) < \widetilde{\sigma}_{\varepsilon}^2(h)$ is that the transition probability matrix is symmetric, and that $(p_{ll} - p_{hl}) = (p_{hh} - p_{lh})$.

14. Note again that the covariance between the pricing kernel and the bond price at maturity $n-1$, induced by the jth shock, is $-\lambda_j[\Gamma'_{.j}\widetilde{b}_{n-1,.})]$.

15. Notice that the extent to which the model's prediction fails to hold is not uniform across maturities. For instance, at 1-year maturity the model prediction does not hold in just 4% of the cases, in comparison with the 23% of the cases at 5-year maturity.

16. See, for instance, Rudebusch and Swanson (2008) for a documentation of this inability of DSGE models. Admittedly, more recently, Amisano and Tristani (2010) find that once stochastic regime shifts in structural shocks are accounted for, standard DSGE models, solved to a second-order approximation and estimated on U.S. data, generate non-negligible time-varying risk premia. Although this result is important in itself, in particular in light of the unsuccessful previous attempts in the DSGE literature, their framework falls short of generating term premia that are as sizeable and variable as those extracted, for instance, from reduced-form approaches.

17. Wright (2011), for instance, considers term premia as differences between long-bond nominal yields and expected future short-term rates. By using survey evidence for estimating expected future

interest rates, he relates the downward pattern of the survey-based term premia observed during the 1990s for a group of industrialized countries to the associated decline in long-term inflation uncertainty. Intuitively, falling inflation uncertainty might well relate to a learning process of economic agents toward the then newly adopted inflation targets.

18. Notably, when investigating time variation in expected excess returns in U.S. government bonds, they find that a single linear combination of forward rates predicts excess returns at all maturities with an R^2 value as high as 0.44.

19. The statistic for $\sup F_T(1)$ is just below the 5% critical value.

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