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TERM PREMIA IMPLICATIONS OF MACROECONOMIC REGIME CHANGES

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In 2014 all ECB publications feature a motif taken from the €20 banknote.

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Abstract

Term premia are shown to provide crucial information for discriminating among alternative sources of change in the economy, and 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 80s and 90s. While hardly distinguishable through the lens of standard DSGE models, lower non-policy 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 U.S. improved macroeconomic stability in the 80s and 90s: term premia would have shifted downwards, a fact inconsistent with the evidence of higher premia from early 80s onwards, where term premia are derived following Cochrane and Piazzesi (2005). Conversely, favourable shifts in non-policy innovation variance imply movements in term premia which are at least qualitatively consistent with historical patterns.

JEL classification: E43; E52.

Keywords: Term premia; Regime switching; DSGE models.

Non-Technical Summary

The contribution of this paper is to show how term premia provide crucial information for discriminating among alternative sources of change in the economy, and namely shifts in the variance of structural shocks and changes in the conduct of monetary policy. Notably, a vast literature has identified them as competing explanations about time-varying features of the economic environment experienced by major industrial economies during the 80s and 90s. From one side, Stock and Watson (2002), Sims and Zha (2006) and Justiniano and Primiceri (2008) among many others find that specifications which allow for 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.[1] By contrast, Clarida, Gali, Gertler (2000), Lubik and Schorfheide (2004), Boivin and Giannoni (2006), among others, stress the role played by monetary policy in achieving better macroeconomic performance. In an attempt to reconcile such alternative findings, Benati and Surico (2009) for instance illustrate how VAR evidence, typically interpreted as supportive of time variation in shock variances, are in fact compatible with changes in monetary policy within a New Keynesian framework. 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 equally affect inflation persistence. More broadly, that such distinct sources of change appear to manifest themselves into 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, while 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 instead 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

[1] For a documentation of time variation in overall macroeconomic stability in the U.S. during the 80's and 90's, see, for instance, Kim and Nelson (1999) and McConnell and Perez-Quiroz (2000).

means that financial assets carry low payoffs in bad times, thus fail to provide insurance when most needed, and hence command positive premia. First, central to the finding that smaller variances of non-policy shocks are associated to 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, a tighter monetary policy regime brings about a reduction in term premia ultimately because it better insulates 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. All in all, the model's prediction is that a more stable macroeconomic environment is characterised by: (i) higher term premia if such improved stability results from a reduction in the variance of non-policy shocks; (ii) lower term premia if macroeconomi stability is instead induced by tighter monetary policies. When considering the U.S. improved macroeconomic stability of the 80s and 90s through the interpretative filter of a standard DSGE model, the implication is that a move to tighter monetary policy regimes alone cannot explain such better outturn. 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 downwards. But such prediction is inconsistent with the empirical evidence for U.S. of higher expected excess returns experienced from early 80s onwards. On the other hand, favourable shifts in the variance of non-policy innovations imply movements in expected returns which are at least qualitatively consistent with historical patterns.

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-80s by employing an affine asset pricing model. Such change is interpreted as stemming from 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. By contrast, our set-up 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 vector autoregressive (VAR) representation of DSGE model's solution makes a reduced-form approach unsuited, and by 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 instead 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.

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, and namely shifts in the variance of structural shocks and changes in the conduct of monetary policy. Notably, a vast literature has identified them as competing explanations about time-varying features of the economic environment experienced by major industrial economies during the 80s and 90s. From one side, Stock and Watson (2002), Sims and Zha (2006) and Justiniano and Primiceri (2008) among many others find that specifications which allow for 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 allowing for heteroskedasticity in structural shocks, there is no clear evidence of other changes in the economy, and notably in the monetary policy conduct.[2] By contrast, Clarida, Gali, Gertler (2000), Lubik and Schorfheide (2004), Boivin and Giannoni (2006), among others, stress the role played by monetary policy in achieving better macroeconomic performance; according to this interpretation, the change in U.S. monetary policy around the early 80s 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 such alternative findings, Benati and Surico (2009) for instance illustrate how VAR evidence, typically interpreted as supportive of time variation in shock variances, are in fact compatible with changes in monetary policy within a New Keynesian framework. In essence, their emphasis rests on the fact that variations in structural shock variances are hardly distinguishable from changes in policy rule coefficients within standard 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 equally affect inflation persistence. 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 quantitatively larger impact. More broadly, that such distinct sources of change appear to manifest them-

[2] For a documentation of time variation in overall macroeconomic stability in the U.S. during the 80’s and 90’s, see, for instance, Kim and Nelson (1999) and McConnell and Perez-Quiroz (2000).

selves into 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, while 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 instead 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, thus fail to provide insurance when most needed, and hence command positive premia. First, central to the finding that smaller variances of non-policy shocks are associated to 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 characterised by: (i) higher term premia if such improved stability results from a reduction in the variance of non-policy shocks; (ii) lower term premia if such better stability is instead induced by tighter monetary policies. When considering the U.S. improved macroeconomic stability of the 80s and 90s through the interpretative filter of a standard DSGE model, the implication is that a move to tighter monetary policy regimes alone cannot explain such better outturn. 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 downwards. But such prediction is inconsistent with the empirical evidence for U.S. of higher expected excess returns experienced from early 80s onwards. On the other hand, favourable shifts in the variance of non-policy innovations imply movements in expected returns which 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.[3]

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-80s by employing an affine asset pricing model. Such 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 on 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. By contrast, our set-up 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 vector autoregressive (VAR) representation of DSGE model's solution makes a reduced-form approach unsuited, and by 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

[3] 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.

intuition instead 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 for macroeconomic dynamics in an estimated medium-scale DSGE model which allow for regime changes in structural parameters and stochastic volatilities.

More broadly, the present analysis provides contribution to a growing literature that attempts to model jointly the dynamics of macroeconomic variables and bond yields within structural DSGE models.[4] 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 inference 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), Bekaert et al. (2010), Doh (2007), Amisano and Tristani (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.[5] 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

[4]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 vector autoregressive (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.

[5]Precisely, Bekaert et al. (2010) consider a log-linear log-normal 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 results from heteroskedasticity in the model structural shocks (time-variation in the "amount of risk").

variables and bond prices. Second, it contributes to sharpen parameter estimates, typically poorly identified in DSGE models' estimations.[6]

The rest of the paper is organised as follows. Section 2 briefly describes the modelling framework and motivates the paper. Section 3 incorporates the term structure into the analysis and derive 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 mid-80s 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).[7] Specifically, the model considered here takes the form of a Markov-switching rational expectations model of the type popularised by Farmer, Waggoner and Zha (2011). Regime switching is assumed to govern the process 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 80s and 90s. 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).

[6] For an investigation on identification issues in DSGE models, see Canova and Sala (2009) for instance.

[7] See also Goodfriend and King (1997) and Clarida, Galí and Gertler (1999).

2.1 The baseline model

The model consists of an intertemporal IS equation (2.1) and an expectations-augmented aggregate supply equation (2.2), which are derived by log-linear approximating optimal behaviour of households and firms. For present purposes, we consider the following specification:

$$\widehat{y}_t = \gamma E_t \widehat{y}_{t+1} + (1-\gamma)\widehat{y}_{t-1} - \sigma^{-1}(\widehat{i}_t - E_t \widehat{\pi}_{t+1}) + \varepsilon_{y,t} \tag{2.1}$$

$$\widehat{\pi}_t = \frac{\beta}{1+\alpha\beta} E_t \widehat{\pi}_{t+1} + \frac{\alpha}{1+\alpha\beta}\widehat{\pi}_{t-1} + k\widehat{y}_t + \varepsilon_{\pi,t} \tag{2.2}$$

where $\widehat{y}_t$ is the output gap, $\widehat{\pi}_t$ the inflation rate, $\widehat{i}_t$ 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:

$$\widehat{i}_t = \rho\widehat{i}_{t-1} + (1-\rho)(\phi_\pi(\xi_t^m)\widehat{\pi}_t + \phi_y(\xi_t^m)\widehat{y}_t) + \varepsilon_{i,t} \tag{2.3}$$

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 $p_{ij}^m \equiv Prob(\xi_t^m = i|\xi_{t-1}^m = j)$. Macroeconomic disturbances follow exogenous first-order autoregressive processes subject to switches in their conditional variances:

$$\varepsilon_{z,t} = \rho_z \varepsilon_{z,t-1} + \sigma_z(\xi_t^s)\epsilon_{zt} \quad for \quad z = i, \pi, y \tag{2.4}$$

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 recasted into the generalised form analysed by Farmer, Waggoner and Zha (2011)

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

where $\widehat{s}_t = [\widehat{i}_t \;\; \widehat{\pi}_t \;\; \widehat{y}_t \;\; E_t\widehat{\pi}_{t+1} \;\; E_t\widehat{y}_{t+1} \;\; \varepsilon_{i,t} \;\; \varepsilon_{\pi,t} \;\; \varepsilon_{y,t}]$, $\epsilon_t = [\epsilon_{i,t} \;\; \epsilon_{\pi,t} \;\; \epsilon_{y,t}]$, and $\xi_t = [\xi_t^m, \xi_t^s]$. Following closely Farmer, Waggoner and Zha (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 popularised by Sims and Zha (2006) and Sims at al. (2008), building on the seminal work by Hamilton (1989) in a univariate context. Specifically, the markov-switching vector autoregressive (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:

$$\begin{aligned} \widehat{x}_t &= \Phi(\xi_t)\widehat{x}_{t-1} + \Gamma(\xi_t)\epsilon_t \\ \epsilon_t &\sim N(0, \Sigma(\xi_t)\Sigma(\xi_t)') \end{aligned} \tag{2.5}$$

For ease of exposition, equation (2.5) can also be rearranged as:

$$\widehat{x}_t = \Phi(\xi_t)\widehat{x}_{t-1} + \Gamma(\xi_t)\epsilon_t = \Phi(\xi_t)\widehat{x}_{t-1} + \Gamma(\xi_t)\Sigma(\xi_t)u_t = \Phi\widehat{x}_{t-1} + v_t \tag{2.6}$$

where the reduced-form innovation covariance matrix is equal to

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

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, and namely: i) shifts in the variance of non-policy 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 above 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 non-policy shock regimes and changes in the monetary policy coefficients are in fact hardly distinguishable by merely looking at the model's dynamics implied by (2.5). 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 upon 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 to monetary policy under determinacy and indeterminacy turn out to imply only modest differences in the way macroeconomic variables respond to shocks over time. Even more difficult is to discriminate among the sets of impulse responses under the two monetary policy regimes once accounting for uncertainty around the median responses. 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 arguments, and defining $\widehat{\pi}_t = e'\widehat{x}_t$, where e is a 6×1 vector with a 1 in the entry associated to $\widehat{\pi}_t$, then 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 function of the persistence parameters associated to 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 trough the lens of standard DSGE models warns against simplistic interpretations of reduced-form evidence in structural terms, and ultimately cast some doubts about 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 above 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 in shedding some light on the nature of change in the model economy. Following closely Wu (2006), Bekaert et al. (2010), Nimark (2008) among others, the term structure is derived by log-normalising the model's Euler equation. When doing so, 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 previous section continues to characterise the dynamics of standard macroeconomic variables. Such a log-linear log-normal modelling 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 modelling framework, additional channels emerge through which the sources of change described above might manifest themselves differently. The yield curve appears a natural candidate for that purpose. At the same time, simply (log-)linearising the bond pricing equation would mean ignoring key aspects of the term structure, and notably the compensation for risk. Via log-normalisation 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 that 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:

$$P_t^{(n)}(\xi_t) = E_t(M_{t+1} P_{t+1}^{(n-1)}(\xi_{t+1})) \tag{3.7}$$

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 log-normal approximation of equation (3.7), the pricing equation can be expressed as:

$$p_t^{(n)}(h) = \sum_j p_{hj} \left(E_t[m_{t+1} + p_{t+1}^{(n-1)}(j)|\xi_{t+1} = j] + \frac{1}{2} Var_t[m_{t+1} + p_{t+1}^{(n-1)}(j)|\xi_{t+1} = j] \right) \quad (3.8)$$

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

$$m_{t+1} = -(\delta'\widehat{s}_t + \bar{i}) - 0.5\lambda(\xi_{t+1})'\Sigma(\xi_{t+1})\Sigma(\xi_{t+1})'\lambda(\xi_{t+1}) - \lambda(\xi_{t+1})'\epsilon_{t+1} \quad (3.9)$$

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 (3.8), (3.9) and the law of motion of endogenous variables (2.5), bond prices, and thus yields, are affine functions of macroeconomic variables:

$$\begin{aligned} p_t^{(n)}(h) &= \widetilde{a}_n(h) + \widetilde{b}_n(h)\widehat{s}_t \\ i_t^{(n)}(h) &= -\frac{1}{n}p_t^{(n)}(h) = -\frac{\widetilde{a}_n(h)}{n} - \frac{\widetilde{b}_n(h)}{n}\widehat{s}_t = a_n(h) + b_n(h)\widehat{s}_t \end{aligned} \quad (3.10)$$

$$\begin{aligned} \widetilde{a}_n(h) &= \sum_j p_{hj}[\widetilde{a}_{n-1}(j) - \bar{i} + 0.5\widetilde{b}'_{n-1}(j)\Gamma(j)\Sigma(j)\Sigma(j)'\Gamma(j)'\widetilde{b}_{n-1}(j) - \widetilde{b}'_{n-1}(j)\Gamma(j)\Sigma(j)\Sigma(j)'\lambda(j) \\ \widetilde{b}_n(h) &= \sum_j p_{hj}(-\delta' + \widetilde{b}'_{n-1}(j)\Phi(j))' \end{aligned}$$

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) return on buying a n-period bond at time t and sell 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:

$$\begin{aligned} E_t\vartheta_{t+1}^{(n)}(h) &= \sum_j p_{hj} \left(-Cov_t[m_{t+1}, p_{t+1}^{(n-1)}(j)|\xi_{t+1} = j] - \frac{1}{2}Var_t[p_{t+1}^{(n-1)}(j)|\xi_{t+1} = j] \right) \quad (3.11) \\ &= \sum_j p_{hj} \left(-Cov_t[m_{t+1}, E_{t+1}\sum_{i=1}^{n-1} m_{t+1+i}|\xi_{t+1} = j] - \frac{1}{2}Var_t[p_{t+1}^{(n-1)}(j)|\xi_{t+1} = j] \right) \end{aligned}$$

By using the model notation, expected returns take the following form:

$$E_t \vartheta_{t+1}^{(n)}(h) = \sum_j p_{hj} \left(\lambda'(j)\Sigma(j)\Sigma(j)'\Gamma(j)'\widetilde{b}_{n-1}(j) - 0.5\widetilde{b}'_{n-1}(j)\Gamma(j)\Sigma(j)\Sigma(j)'\Gamma(j)'\widetilde{b}_{n-1}(j) \right) \tag{3.12}$$

Specifically, the compensation for carrying certain units of risk is captured by the conditional covariance between bond prices and the pricing kernel.[8] 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 above-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 volatility. Specifically, the sample period ranges from 1980Q1 to July 2007, and the macro data used in the estimation comprise: inflation (annualised 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); nominal interest rate (annualised 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 maximise the log posterior, and second a Random Walk Metropolis algorithm

[8]Expected excess returns can be equivalently expressed in terms of the conditional covariance between current and expected future nominal pricing kernel.

to draw from the posterior distribution. The parameters specifically associated to the regime of high macroeconomic volatility are calibrated as follows. High values of variance of non-policy shock, $\sigma_{\varepsilon\pi}$ and $\sigma_{\varepsilon y}$, are calibrated by scaling up the posterior median estimates by the factor 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.[9] Finally, the probabilities to persist in a given regime, p_{ii}^k, 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 non-policy shock variances Fig 1 illustrates how a switch to a regime of lower non-policy shock variance is associated to 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 non-policy shock, while the remaining parameters are kept fixed at their posterior estimates.[10]

[FIGURE 1 ABOUT HERE]

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 risk-free short rate because their expected payoffs ($E_{t+1}p_{t+1}$) 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

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

[10] For convenience of further exposition, we assume that this situation is one in which the equation (3.11) takes the simplified form where only $\Sigma(j)$ is function of the state j of the Markov-switching process, while λ and Γ and $\widetilde{b}_{n-1}$ are not function of the state j.

a negative premia 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 (3.11). Intuitively, by having higher returns when most needed, long bonds provide an 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 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[11]

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

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

$$E_t \vartheta_{t+1}^{(2)} = -Cov_t[m_{t+1}, E_{t+1} p_{t+1}] = -Cov_t[m_{t+1}, E_{t+1} m_{t+2}] \approx -Cov_t[\Delta y_{t+1}, \Delta y_{t+2}] - Cov_t[\pi_{t+1}, \pi_{t+2}]$$

From the last equation, it is then evident how larger variances of non-policy 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 considering also the cross-covariance terms, those including $\varepsilon_{y,t+1}$ similarly imply that larger variance of demand shocks weigh negatively on term premia, as they induce a larger positive covariance between demand shocks and inflation and output (growth). Admittedly, these simplified lines of explanations are somewhat lost when considering more general cases in which σ and γ are different from 1, as additional terms enter in the definition of term

[11] The convexity term due to Jensen's inequality is disregarded here.

premia. However, for a wide range of alternative parameter calibrations, the rise in the autocovariance of inflation and output gap associated to larger variance of non-policy shocks remain quantitatively the prevailing forces behind our finding.

To gain further formal insight, simple manipulations of (3.12) lead to the following alternative expression for excess returns under the time-t regime $s_t = h$

$$E_t \vartheta_{t+1}^{(n)}(h) = \sum_j p_{hj} \left(\lambda_i \sigma_{\varepsilon i}^2 (\Gamma'_{.1} \widetilde{b}_{n-1}) + \lambda_\pi \sigma_{\varepsilon\pi}^2(j)(\Gamma'_{.2} \widetilde{b}_{n-1}) + \lambda_y \sigma_{\varepsilon y}^2(j)(\Gamma'_{.3} \widetilde{b}_{n-1}) \right) \tag{3.13}$$

where λ_k is the k^{th} entry of the vector of prices of risk λ, and $\Gamma'_{.k}$ is the (transpose of the) k^{th} column vector of the matrix Γ, vector that captures the impact of the k^{th} macroeconomic shock. In other words, (3.13) expresses excess returns at different maturities in terms of the contribution of monetary, cost-push, and demand shocks, respectively. The difference between excess returns under the low (l) and the high (h) non-policy shock variance regimes can then be expressed as:

$$E_t \vartheta_{t+1}^{(n)}(l) - E_t \vartheta_{t+1}^{(n)}(h) = \lambda_\pi (\Gamma'_{.2} \widetilde{b}_{n-1})(\widetilde{\sigma}_{\varepsilon\pi}^2(l) - \widetilde{\sigma}_{\varepsilon\pi}^2(h)) + \lambda_y (\Gamma'_{.3} \widetilde{b}_{n-1})(\widetilde{\sigma}_{\varepsilon y}^2(l) - \widetilde{\sigma}_{\varepsilon y}^2(h)) \tag{3.14}$$

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 (3.14) 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 to cost-push and demand shocks, respectively. As evident from equation (3.9), 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 namely λ_π is positive.[12]. Similarly, λ_y is also positive. Intuitively, a demand shock brings about an increase in both inflation and output which unambiguously leads to a fall in the nominal pricing kernel,

[12] The robustness of various findings will be investigated in the next section.

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.[13] With λ_π and λ_y positive, $(\Gamma'_{.2}\widetilde{b}_{n-1,.})$ and $(\Gamma'_{.3}\widetilde{b}_{n-1,.})$ negative, and $(\widetilde{\sigma}^2_{\varepsilon_k}(l) - \widetilde{\sigma}^2_{\varepsilon_k}(h))$ negative for $k = \pi, y$, both multiplicative terms in (3.14) are positive. As a result, excess returns are unambiguously larger under smaller variances of non-policy shocks.[14] All in all, the first model's prediction can be summarised as follows: moving to a more stable macroeconomic environment would bring about an upward shift in term premia when such improved stability results from lower variance of non-policy structural disturbances.

Regime switching in the monetary policy conduct By 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 Fig 2 by displaying that excess returns under the tight policy regime are lower than under the loose regime at all maturities (horizontal axis).

[FIGURE 2 ABOUT HERE]

To isolate the impact of changing monetary policy on excess returns, regime switching is assumed to characterise the dynamic of policy coefficients only, while 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 correlated with long bond prices. As a result, long bonds command lower premia under tighter regimes. In essence, these regimes make inflation less negatively

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

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

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:

$$\begin{aligned} E_t\vartheta_{t+1}^{(n)}(h) - E_t\vartheta_{t+1}^{(n)}(l) \;=\;& \sigma^2_{\varepsilon i}[\widetilde{\lambda}_i(h)(\Gamma'_{.1}(h)\widetilde{b}_{n-1,.}(h)) - \widetilde{\lambda}_i(l)(\Gamma'_{.1}(l)\widetilde{b}_{n-1,.}(l)))] + \\ & \sigma^2_{\varepsilon\pi}[\widetilde{\lambda}_\pi(h)(\Gamma'_{.2}(h)\widetilde{b}_{n-1,.}(h)) - \widetilde{\lambda}_\pi(l)(\Gamma'_{.2}(l)\widetilde{b}_{n-1,.}(l))] + \\ & \sigma^2_{\varepsilon y}[\widetilde{\lambda}_y(h)(\Gamma'_{.3}(h)\widetilde{b}_{n-1,.}(h)) - \widetilde{\lambda}_y(l)(\Gamma'_{.3}(l)\widetilde{b}_{n-1,.}(l))] \end{aligned} \quad (3.15)$$

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 k^{th} shock of the vector ϵ_t under the policy regime $j = h, l$.

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

$$\begin{aligned} E_t\vartheta_{t+1}^{(n)}(h) - E_t\vartheta_{t+1}^{(n)}(l) \;=\;& \widetilde{\sigma}^2_{\varepsilon i}[\lambda_i(h)(\Gamma'_{.1}(h)\widetilde{b}_{n-1,.}(h)) - \lambda_i(l)(\Gamma'_{.1}(l)\widetilde{b}_{n-1,.}(l))] + \\ & \widetilde{\sigma}^2_{\varepsilon\pi}[\lambda_\pi(h)(\Gamma'_{.2}(h)\widetilde{b}_{n-1,.}(h)) - \lambda_\pi(l)(\Gamma'_{.2}(l)\widetilde{b}_{n-1,.}(l))] + \\ & \widetilde{\sigma}^2_{\varepsilon y}[\lambda_y(h)(\Gamma'_{.3}(h)\widetilde{b}_{n-1,.}(h)) - \lambda_y(l)(\Gamma'_{.3}(l)\widetilde{b}_{n-1,.}(l))] \end{aligned} \quad (3.16)$$

The first term in (3.16) 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.[15] 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

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

to covary less in absolute value. The second term in (3.16) 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, while 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, and namely that λ_π is lower under the tight regime. From the other side, it implies that the tight regime is characterised 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. While 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 (3.16) 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.

4 Sensitivity analysis (via Bayesian predictive checks)

This sections confirms that the findings described in the previous 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 parameters 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 a case captures looser monetary policy regime, characterised by bringing down the drawn values of the policy coefficients at the boundary with the indeterminacy region.

Fig 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 5000 parameters draws. Points below (above) the 45-degree line mean that term premia are lower (higher) under the high shock variance regime.

[FIGURE 3 ABOUT HERE]

Our finding of higher excess returns associated to lower shock variances appears very robust, as illustrated graphically by the concentration of all points below the 45-degree line. In light of the analysis in the previous section, the robustness of our finding is only partly surprising considering that scaling up non-policy 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, Fig 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 looser monetary policy rule (vertical axis). The finding that tighter regimes are associated to lower excess returns is confirmed, as illustrated by the fact that almost all points are located above the 45-degree line. This finding appears only slightly less robust than in the previous case, since in just around 1.8 percent of the cases the second model's prediction is reversed.[16] Yet, the overall robustness of the finding is remarkable taking into account the large region of the parameter space spanned by the draws, illustrated indirectly by the wide range of values taken by term premia.

[FIGURE 4 ABOUT HERE]

[16] 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.

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 mid-80s 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 which fit relatively well model-free ex-post excess returns. While, 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. Rudebusch and Swanson (2008) for instance documents such inability of DSGE models. Admittedly, more recently, Amisano and Tristani (2010) find that once accounting for stochastic regime shifts in structural shocks, standard DSGE models, solved to a second order approximation and estimated on US data, generate non-negligeble time-varying risk premia. While this result is in itself important, in particular in light of the unsuccessful previous attempts in the DSGE literature, yet their framework falls short to generate term premia which are as sizeable and variable as those extracted, for instance, from reduced-form approaches.

Finally, the focus on excess returns as a measure of term premia allows to net out the level of inflation and of interest rates, level which could be well influenced by time-variation in the underlying inflation target, or in agents' beliefs about the inflation target, both channels not considered in this framework.[17]

[17] Wright (2011), for instance, considers term premia as difference 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 90's for a group of industrialised 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

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

$$rx_{t+1}^{(n)} = b_n(\gamma_0 + \gamma_1 i_t^{(1)} + \gamma_2 f_t^{(2)} + ...\gamma_5 f_t^{(5)}) + \nu_{t+1}^{(n)} \tag{5.17}$$

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 separately identified, the average value of b_n is normalised to 1. Equation (5.17) 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 following regressions

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

where $(\gamma' f_t)$ is the single linear combination of forward rates between $t + n - 1$ and $t + n$ which predicts excess returns at all maturities. On the basis of such approach by Cochrane and Piazzesi (2005), one-year holding period return for U.S. government bonds with maturities from 2 to 5 years are thus derived for the period 1965Q1-2003Q4. Fig 5 depicts the time series for average (across maturities) expected returns, along with the corresponding ex-post excess returns.

[FIGURE 5 ABOUT HERE]

Overall, ex-post excess returns tend to display substantially fluctuation over time, as captured by the standard deviation been as high as 4 percent over the relevant period. Of main interest for present purposes is the extent to which expected returns have experienced a regime change around mid-80s, iin ways that can be informative about the sources of

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 predict excess returns at all maturities with an R^2 value as high as 0.44

economic transformation experienced by the U.S. economy in that period. Specifically, two distinct sub-periods characterizing the post-WW II US macroeconomic history are considered, and namely before and after the (end of the) Volcker disinflation, respectively.

First, over the sub-period from 1965Q1 to 1979Q3, the mean of (average-across maturities) ex-post and expected excess returns is around -0.8 and -0.2 percent respectively. Focusing more specifically on the 70's, a decade characterised by particularly high macroeconomic instability, the mean of ex-post and expected excess returns is -0.5 and 0.3 percent respectively, while the standard deviation is almost unchanged in comparison to the sample period 1965Q1:1979Q3 for both ex-post and expected excess returns. Second, over the sub-period from 1984Q1 to 2003Q4, the mean of (average-across maturities) ex-post and expected excess returns is instead as high as 2.4 and 1.5 percent respectively, while the standard deviation of both ex-post and expected excess returns remains broadly unchanged in comparison to the first sub-period. 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 a direct targeting of monetary aggregates, and of the two recessionary episodes occurred in such 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 above, 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, being the null hypothesis of no structural break rejected at 5% confidence interval for the number 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 tests sequentially k versus $k+1$

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

number of breaks, indicate the presence of two breaks. Such 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

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

where $j = 1,\ \ m + 1$, with m being the total number of breaks, identified to be 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, 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 sub-samples which roughly correspond to the above investigated sub-periods, and namely the pre-Volcker period, the Volcker disinflation period, and the post-84 period. While not statistically different from zero in the first sub-period, the mean of excess returns is estimated to be as high as 4 percent during the second sub-period, and 1 percent in the third sub-period. 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 80s.

Notably, when using this finding to interpret the sources of the U.S. improved macroeconomic stability in the 80s and 90s through the lens of a small-scale DSGE model, the implication is that changes in monetary policy alone cannot explain such better outturn. 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 downwards, a fact inconsistent with the empirical evidence of rising expected returns experienced from the early 80s onwards. On the other hand, favourable shifts in the variances of non-policy innovations imply movements in expected returns which are at least qualitatively consistent with historical patterns. Finally, this finding is consistent with a number of studies which identify the change in the variance of structural shocks as the major source behind the U.S. economic transformation in the late 70's and early 80s, studies which 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 dynamic 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 rationalising consumption and investment decisions ultimately rests on their ability to also capture important features of those markets which are relevant for these economic decisions, and notably the bond markets. Moreover, adding bond price data into the econometric analysis might contribute to mitigate 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 non-policy 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 80s and 90s. While 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 interpreting the sources of U.S. improved macroeconomic stability of the 80s and 90s in light of these findings, the implication is that a move to tighter monetary regimes alone cannot explain such better outturn, as this would have implied lower expected excess returns, in contrast with the empirical evidence of rising (average) expected excess returns experienced from early 80s onwards. Favourable shifts in the variance of non-policy innovations instead imply movements in expected returns which are at least qualitatively consistent with empirical evidence. Admittedly, there are two major caveats to our findings. First, changes in monetary policy are here modelled in terms of switches in the response coefficients of the policy rule. While this represents the most common characterisation 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 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 which 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 immediately evident. Second, canonical DSGE models fall short in characterising an empirically plausible term structure. In particular, these models fail to generate term premia which 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 to shifts in shock variance and monetary policy regimes, respectively. Being the term structure 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 simultaneously macroeconomic and term structure evidence: 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, while still preserving a good fit of main macroeconomic variables. However, as estimating models solved to third order approximation is currently unfeasible, such 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 US data, generate non-negligible time-varying risk premia once accounting for stochastic regime shifts; yet their framework falls somewhat short to generate term premia which are as sizeable and variable as those extracted, for instance, from reduced-form approaches such the one by Cochrane and Piazzesi (2005) employed here.

References

[1] Amisano, G. and O. Tristani (2010) A nonlinear DSGE model of the term structure with regime shifts. 2010 Meeting Papers 234, Society for Economic Dynamics.

[2] Ang, A. and M. Piazzesi (2003) A No-Arbitrage Vector Autoregression of Term Structure Dynamics with Macroeconomic and Latent Variables. *Journal of Monetary Economics* 50, 745-787.

[3] Bai, J., and P. Perron (2003) Computation and analysis of multiple structural change models. *Journal of Applied Econometrics* 18(1), 1-22.

[4] Bai J, and P. Perron (1998) Estimating and testing linear models with multiple structural changes. *Econometrica* 66 47–78.

[5] Benati L. and P. Surico (2009) VAR Analysis and the Great Moderation. *American Economic Review* 99(4), 1636-52.

[6] Bekaert, G., Cho, S., and A. Moreno (2010) New-Keynesian macroeconomics and the term structure. *Journal of Money, Credit and Banking* 42(1), 33-62.

[7] Bianchi, F. (2012) Regime switches, Agents' Beliefs, and Post-World WAR II U.S. Macroeconomic Dynamics. Working Papers No. 120 (ERID), Duke University.

[8] Bikbov, R. and M. Chernov, (2008) Monetary Policy Regimes and the Term Structure of Interest Rates. CEPR Discussion Papers 7096.

[9] Calvo, G. (1983) Staggered Prices in a Utility-Maximizing Framework. *Journal of Monetary Economics* 47, 383-398.

[10] Canova F., L. Sala (2009) Back to Square One: Identification Issues in DSGE Models. *Journal of Monetary Economics* 56 (4), 431-449.

[11] Christiano, L., M. Eichenbaum and C. Evans (2005) Nominal rigidities and the dynamic effects of a shock to monetary policy. *Journal of Political Economy* 113, 1-45.

[12] Clarida, R., J. Gali, and M. Gertler (2000) Monetary Policy Rules and Macroeconomic Stability: Evidence and Some Theory. *Quarterly Journal of Economics* CXV(1), 147-180.

[13] Clarida, R., Galí, J., and M. Gertler (1999) The Science of Monetary Policy: A New Keynesian Perspective. *Journal of Economic Literature* 37, 1661-1707.

[14] Cochrane, J. H., and M. Piazzesi (2005) Bond Risk Premia. *American Economic Review* 95(1), 138-160.

[15] Cogley, T., and T.J. Sargent (2005) Drifts and Volatilities: Monetary Policies and Outcomes in the Post WWII U.S.. *Review of Economic Dynamics* 8, 262-302.

[16] Davig, T., and T. Doh (2009) Monetary Policy Regime Shifts and Inflation Persistence", Federal Reserve Bank of Kansas City working paper, 2008.

[17] Del Negro, M., and S. Eusepi (2011) Fitting Observed Inflation Expectations. *Journal of Economic Dynamics and Control* 35(12), 2105-2131.

[18] Farmer R. E. A., D. F. Waggoner, and T. Zha (2011) Minimal State Variable Solutions To Markov-Switching Rational Expectations Models. *Journal of Economic and Dynamic Control* 35(12), 2150-2166.

[19] Hamilton, J. D. (1989) A New Approach to the Economic Analysis of Nonstationary Time Series and the Business Cycle. *Econometrica* 57(2), pages 357-384.

[20] Hördahl, P., O. Tristani and D. Vestin (2006) A joint econometric model of macroeconomic and term structure dynamics. *Journal of Econometrics* 131(1-2), 405-44.

[21] Hördahl, P., O. Tristani and D. Vestin (2008) The Yield Curve and Macroeconomic Dynamics. *Economic Journal*, 118(533) 1937-1970.

[22] Kim, C., and C. Nelson (1999) Has The U.S. Economy Become More Stable? A Bayesian Approach Based On A Markov-Switching Model Of The Business Cycle. *The Review of Economics and Statistics* 81(4), 608-616.

[23] Levin, A., and J. B. Taylor (2010) Falling Behind the Curve: A Positive Analysis of Stop-Start Monetary Policies and the Great Inflation. NBER Working Papers 15630.

[24] Liu, Z., D.F. Wagonner, and T. Zha (2009) Asymmetric Expectation Effects of Regime Switches in Monetary Policy. *Review of Economic Dynamics* 12(2), 284-303.

[25] Liu, Z., D.F. Wagonner, and T. Zha (2011) Sources of Macroeconomic Fluctuations: A regime-switching DSGE approach. *QuantitativeEconomics* 2(2), 251-301.

[26] Lubik, T., and F. Schorfheide (2003) Computing Sunspot Equilibria in Linear Rational Expectations Models. *Journal of Economic Dynamics and Control* 28(2), 273-285.

[27] Lubik, T., and F. Schorfheide (2004) Testing for Indeterminacy: An Application to U.S. Monetary Policy. *American Economic Review* 94(1), 190-217.

[28] Lucas, R. E. (1976) Econometric Policy Evaluation: A Critique, Carnegie-Rochester Conference Series on Public Policy, 1, 19-46.

[29] Nimark, P. K. (2008) Monetary Policy with Signal Extraction from the Bond Market. *Journal of Monetary Economics* 55, 1389-1400.

[30] McConnell, M.M., and G. Perez-Quiros (2000) Output Fluctuations in the United States: What Has Changed since the Early 1980s?. *American Economic Review* 90, 1464-1476.

[31] Primiceri, G. E. (2005) Time Varying Structural Vector Autoregressions and Monetary Policy. *Review of Economic Studies* 72(3), 821–52.

[32] Ravenna, F. and J. Seppälä (2006) Monetary Policy and Rejections of the Expectations Hypothesis. mimeo.

[33] Rudebusch, G. D. and E. T. Swanson (2008) Examining the bond premium puzzle with a DSGE model. *Journal of Monetary Economics* 55, 111-126.

[34] Rudebusch, G. D. and E. T. Swanson (2009) The bond premium in a DSGE model with long-run real and nominal risks. Federal Reserve Bank of San Francisco, Working Paper Series 2008-31.

[35] Rudebusch G. D. and T. Wu (2007) Accounting for a Shift in Term Structure Behavior with No-Arbitrage and Macro-Finance Models, Journal of Money, Credit and Banking, 39(2-3), pp. 395-422.

[36] Schorfheide, F. (2005) Learning and monetary policy shifts. *Review of Economic Dynamics* 8, 392-419.

[37] Sims, C. (1996) Solving linear rational expectations models. Manuscript, Yale University.

[38] Sims, C. and T. Zha, (2006) Were There Regime Switches in U.S. Monetary Policy?. *American Economic Review* 96(1), 54-81.

[39] Sims, C., D.F. Waggoner,and T. Zha, (2008) Methods for Inference in Large Multiple-Equation Markov-Switching Models. *Journal of Econometrics* 146, 113-144.

[40] Stock, J. H., and M. W. Watson (2003) Has the Business Cycle Changed and Why ?. In *NBER Macroeconomics Annual 2002*, ed. Mark Gertler and Kenneth Rogoff, pages 159–218. Cambridge, MA: MIT Press.

[41] Woodford, M. (2003) *Interest and Prices: Foundations of a Theory of Monetary Policy*, Princeton University Press.

[42] Wright, J. H. (2011) Term premia and Inflation Uncertainty: Empirical Evidence from an International Panel Dataset. *American Economic Review* 101(4), 1514-1534.

[43] Wu T. (2006) Macro Factors and the Affine Term Structure of Interest Rates. *Journal of Money, Credit and Banking* 38(7), 1847-1875.

Table 1: Bayesian estimation of the model parameters

Parameter	Domain	Density	Prior distribution: Mode	Prior distribution: Standard deviation	Posterior distribution: Median and 90 percent coverage percentile	
k	$\mathbb{R}^+$	Gamma	0.05	0.01	0.04	[0.03; 0.06]
σ	$\mathbb{R}^+$	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	$\mathbb{R}^+$	Gamma	1.00	0.50	2.11	[1.50; 2.78]
ϕ_y	$\mathbb{R}^+$	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]
$\sigma^2_{\varepsilon i}$	$\mathbb{R}^+$	InvGamm	0.5	1.00	1.08	[0.86; 1.40]
$\sigma^2_{\varepsilon\pi}$	$\mathbb{R}^+$	InvGamm	0.9	0.80	0.33	[0.25; 0.46]
$\sigma^2_{\varepsilon y}$	$\mathbb{R}^+$	InvGamm	0.5	0.50	0.10	[0.06; 0.16]

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)

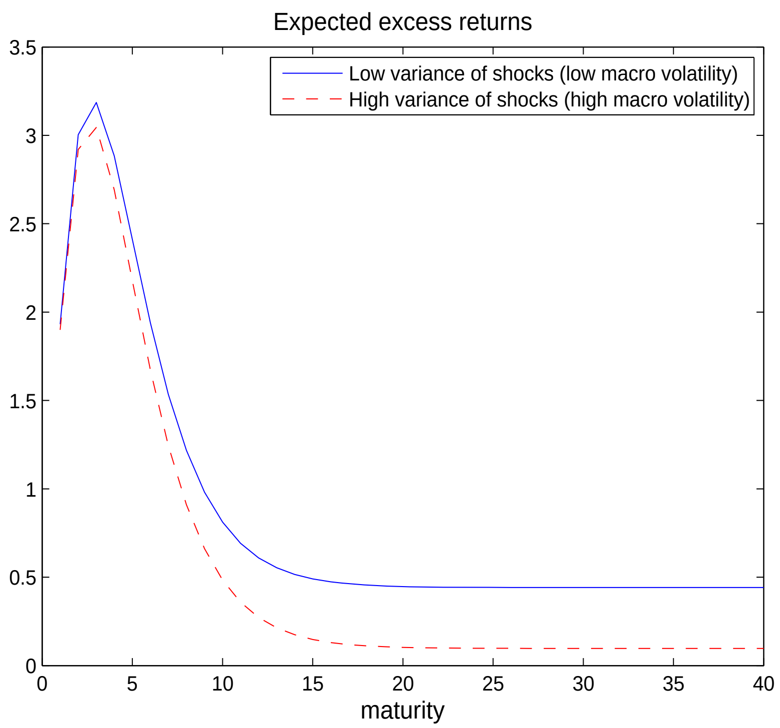

Figure 1: Expected excess returns under low and high non-policy shock variance regimes.

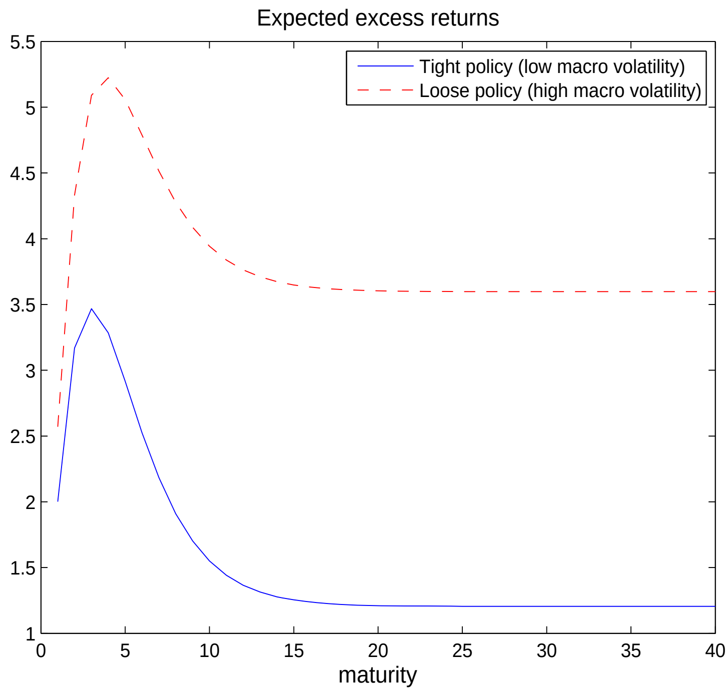

Figure 2: Expected excess returns under tight and loose monetary policy regimes.

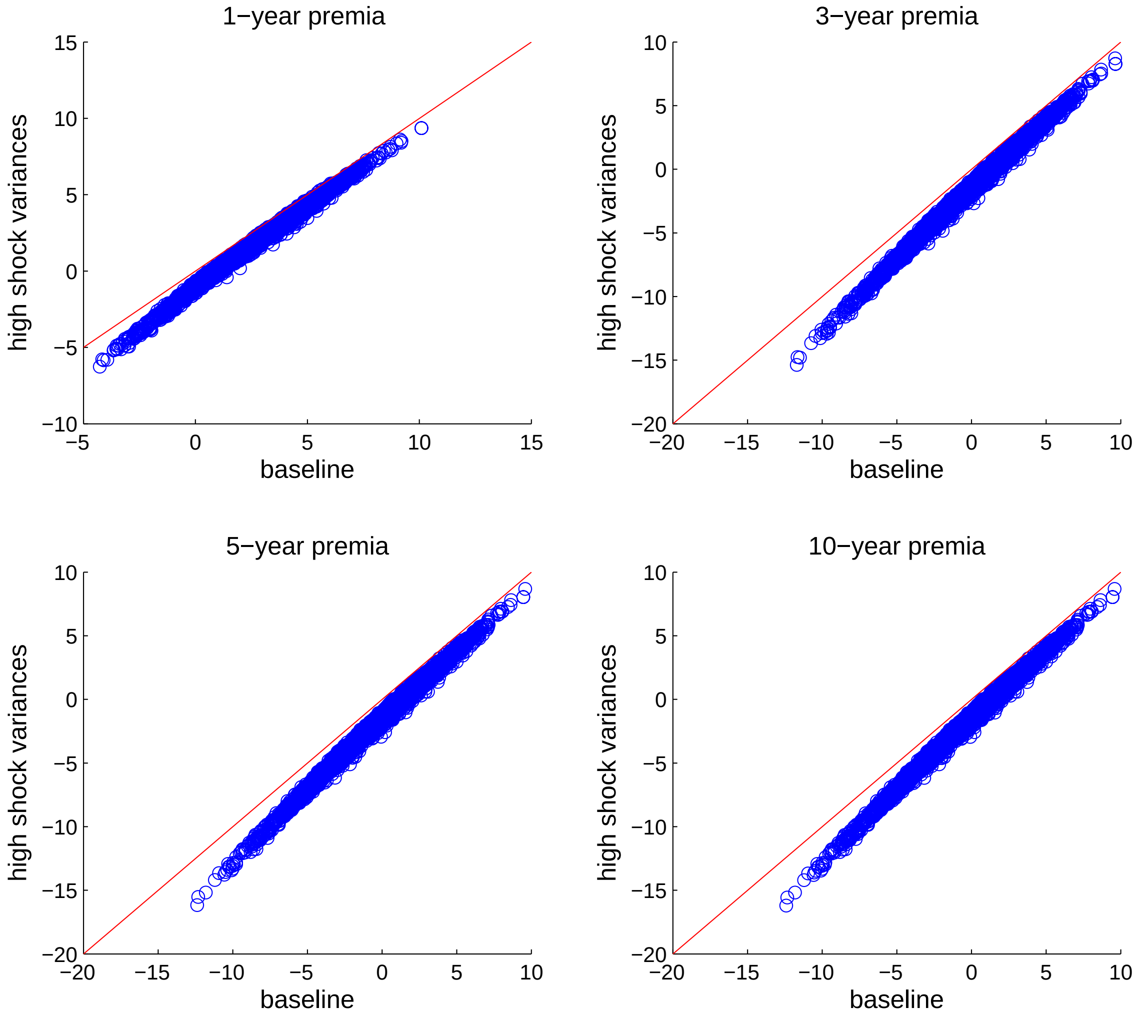

Figure 3: Scatterplot between excess returns under the high shock variances regime and under the low macroeconomic volatility baseline.

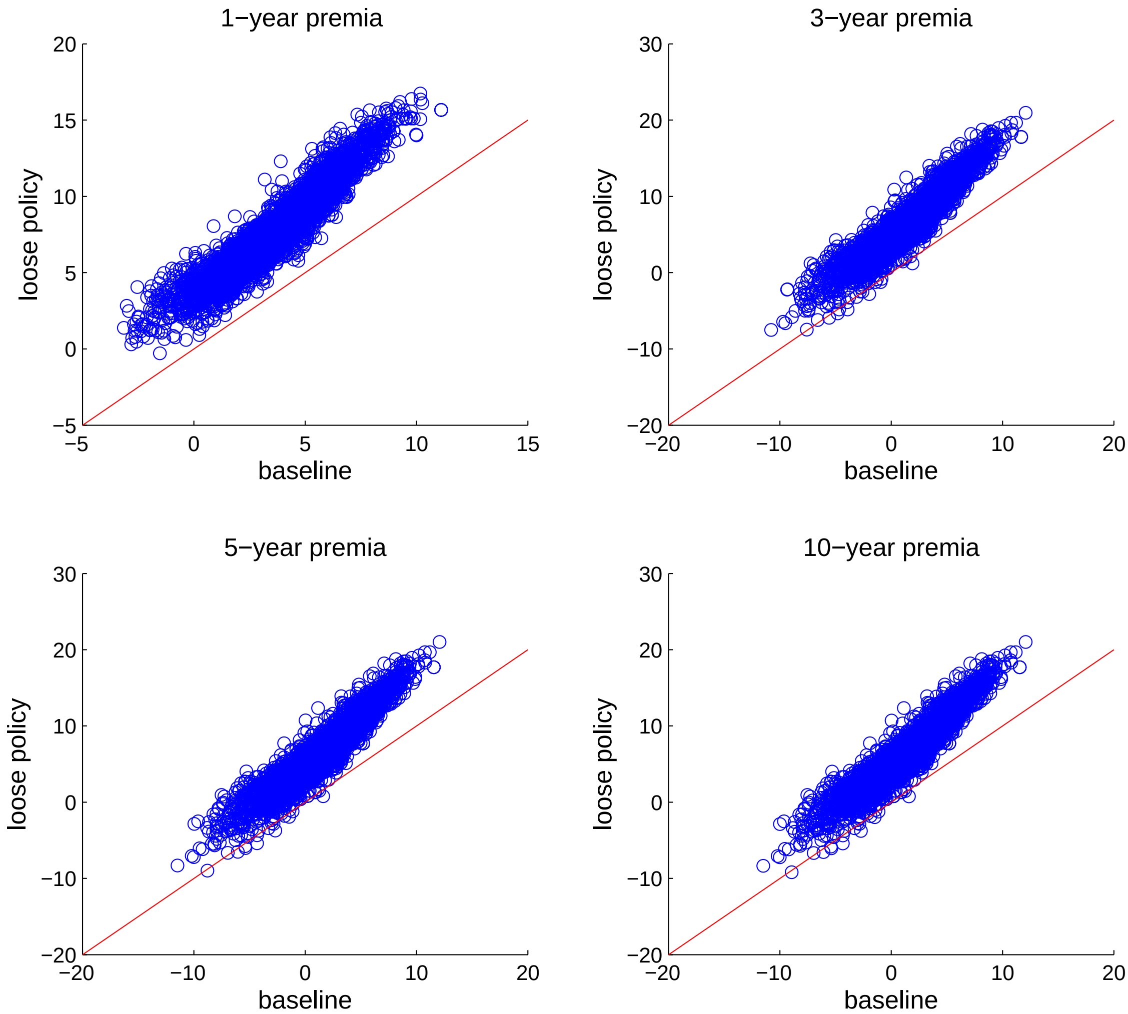

Figure 4: Scatterplot between excess returns under the loose monetary policy regime and under the low macroeconomic volatility baseline.

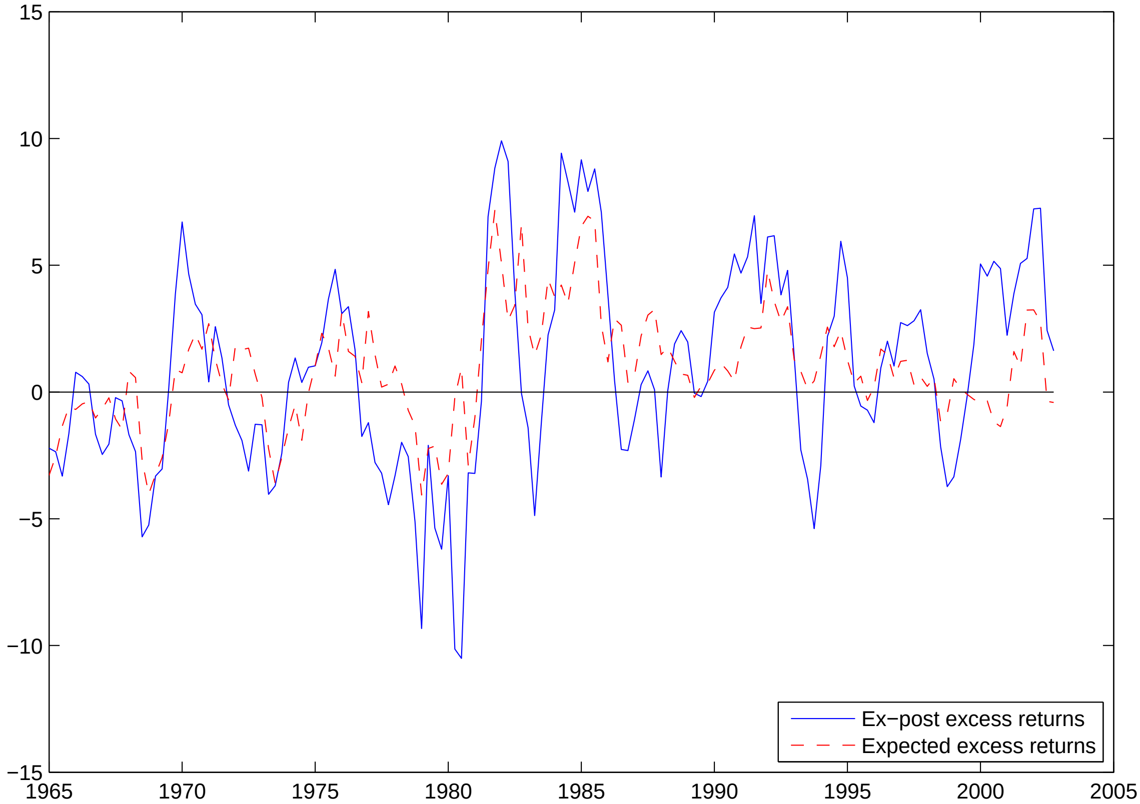

Figure 5: Historical paths of U.S. ex-post and expected excess returns (averages across maturities). The ex-post returns series is shifted to the left so as to line up with the expectation.