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Abstract

We use Bayesian time-varying parameters VARs with stochastic volatility to investigate changes in the marginal predictive content of the yield spread for output growth in the United States and the United Kingdom, since the Gold Standard era, and in the Eurozone, Canada, and Australia over the post-WWII period. Overall, our evidence does not provide much support for either of the two dominant explanations why the yield spread may contain predictive power for output growth, the monetary policy-based one, and Harvey's (1988) 'real yield curve' one. Instead, we offer a new conjecture.

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Keywords: Bayesian VARs; stochastic volatility; time-varying parameters; median-unbiased estimation; Monte Carlo integration.

JEL classification: E42, E43, E47.

Non Technical Summary

Although, since the end of the 1980s, a vast literature has documented the predictive content of the long-short nominal yield spread for future output growth, such finding is still to be regarded essentially as a stylised fact in search of a theory. Currently, there are two main explanations why the nominal yield spread may contain information on future output growth, one dealing with the workings of monetary policy, the other with the interaction between intertemporal consumption smoothing, on the one hand, and the stochastic properties of inflation, as determined by the underlying monetary regime, on the other.

According to the former explanation, a temporary monetary tightening can be expected to produce two results: first, a recession; and second, a fall in inflation, and therefore in inflation expectations. To the extent that the tightening—i.e., the increase in the short rate—is temporary, the fall in inflation expectations automatically guarantees that long rates increase less than short rates, thus causing a flattening of the yield curve. By the same token, a symmetrical argument explains why a monetary expansion causes both a steepening of the yield curve and an economic expansion.

According to the latter explanation, on the other hand, the predictive content of the spread pertains to the *real* term structure, rather than to the *nominal* one. The fact that the predictive content intrinsic to the real term structure translates, or does not translate, to the nominal term structure then crucially depends on the stochastic properties of inflation—in particular, inflation persistence—and therefore, ultimately, upon the nature of the underlying monetary regime. If inflation is, in the limiting case, a pure random walk, so that innovations are entirely permanent, a shock to inflation today shifts expected inflation at all horizons by an identical amount, thus leaving the nominal yield curve, for a given real yield curve, unaffected. In this case the predictive content of the spread intrinsic to the real yield curve translates one-to-one to the nominal yield curve. If, on the other hand, inflation has little persistence—as it was the case under metallic standards, and currently is the case for several inflation-targeting countries—a shock to inflation uniquely increases short-term inflation expectations, leaving instead long-term expectations unaffected, and therefore, for a given real yield curve, by increasing short rates and leaving long rates unchanged, it twists the nominal yield curve, thus 'blurring' the informational content of the real curve.

In order to empirically assess the two theories, and to tentatively discriminate between them, in this paper we investigate changes in the marginal predictive content of the yield spread for future output growth in the United States and the United Kingdom, since the Gold Standard era, and in the Eurozone, Canada, and Australia over the post-WWII period.

Within a univariate context, based on Stock and Watson's Classical time-varying parameters median-unbiased estimation methodology we detect strong evidence of random-walk time-variation in output growth regressions for all countries, with com-

paratively large median-unbiased estimates of the overall extent of parameter drift. Moving to a multivariate context, evidence based on Bayesian time-varying parameters VARs with stochastic volatility does not provide, overall, full support for *either* of the two explanations why the yield spread may contain predictive power for output growth. On the one hand, the 'real yield curve' explanation is contradicted by the fact that in both the United States, the United Kingdom, and Canada over the post-WWII era, the broad decrease in inflation persistence we identify for all the three countries over the second part of the sample was not accompanied by a corresponding decrease in the marginal predictive content of the spread compared to the information already encoded in past output growth. The monetary policy-based explanation, on the other hand, appears incompatible with the fact that, for example, results for the United States during both the interwar and the post-WWII periods clearly point towards several periods during which the spread exhibited predictive power for output growth over and above that already encoded in the short rate. In particular, during both the Volcker recession, and the 2000-2001 one, the spread clearly appears to have possessed additional information compared with that contained in the simplest measure of the monetary policy stance.

1 Introduction

Since the end of the 1980s, a large literature has investigated the predictive content of the long-short nominal yield spread for both inflation, and for the rates of growth of GDP and individual expenditure components.[1] While the spread's predictive content for inflation, once having controlled for lagged inflation, has almost uniformly been found to be low or non-existent, several papers have documented how, both in the United States, and in other OECD countries, the yield spread appears to have contained information on future output growth *independent* of that contained in other macroeconomic aggregates, thus allowing forecasting improvements upon models including standard predictors like indices of leading indicators, inflation measures, etc..[2] Especially intriguing is the finding, documented by Estrella and Hardouvelis (1991) and Plosser and Rouwenhorst (1994), that the informational content of the spread appears to have been independent of both nominal and real short-term interest rates, thus providing *prima facie* evidence that the spread's information may be (at least partly) independent of monetary policy actions.[3] Interestingly, as first documented by Dotsey (1998) and Estrella, Rodrigues, and Schich (2003), in the United States the marginal predictive content of the spread for output growth appears to have largely disappeared in recent years.[4]

Although the predictive content of the spread for output growth has now been systematically documented for almost two decades, such finding is still to be regarded essentially as a stylised fact in search of a theory. Currently, there are two main explanations why the nominal yield spread may contain information on future output growth, one dealing with the workings of monetary policy, the other with the interaction between intertemporal consumption smoothing, on the one hand, and the stochastic properties of inflation, as determined by the underlying monetary regime, on the other.

A simple, 'introductory macro' description of the first explanation runs as follows. A temporary monetary tightening can be expected to produce two results: first, a recession; and second, a fall in inflation, and therefore in inflation expectations. To the extent that the tightening—i.e., the increase in the short rate—is temporary, the

[1] For a literature survey, see Stock and Watson (2003), section 3.1.

[2] As stressed by Stock and Watson (2003), the predictive content of the nominal yield spread for output growth was discovered independently by Laurent (1988), Laurent (1989), Harvey (1988), Harvey (1989), Stock and Watson (1989), Chen (1991), and Estrella and Hardouvelis (1991).

[3] As discussed by Estrella and Hardouvelis (1991), however, this does *not* imply that such information might be systematically used by the monetary authority, as by the Lucas critique—and by Goodhart's law—the spread's informational content could not be reasonably thought to remain intact in the face of systematic attempts to exploit it on the part of the policymaker.

[4] Estrella and Hardouvelis (1991) perceptively conjectured such a phenomenon for the very latest years of their sample period, 1955-1988. In recent months, the apparent breakdown of the relationship between the yield spread and output growth has also received a lot of attention in the press—see e.g. Jennifer Hughes' article on the *Financial Times* of February 9, 2006, page 13 ('A World Turned Inside Out: Why Investors Are Re-Evaluating the Predictive Power of Bonds').

fall in inflation expectations automatically guarantees that long rates increase less than short rates, thus causing a flattening of the yield curve. By the same token, a symmetrical argument explains why a monetary expansion causes both a steepening of the yield curve and an economic expansion. An important point to stress is that, according to this explanation, the predictive content of the spread for future output growth is *entirely spurious*, in the sense that fluctuations in both the spread and future output growth are caused by a third variable, monetary policy actions.

According to the second explanation,[5] on the other hand, the informational content of the spread for output growth is not spurious, but rather *intrinsic*—to put it differently, it finds its origin in the workings of the deep structure of the economy, rather than in monetary policy actions. According to such a view, first clearly articulated by Harvey (1988), based on standard intertemporal consumption smoothing arguments, the predictive content of the spread pertains to the *real* term structure, rather than to the *nominal* one. The fact that the predictive content intrinsic to the real term structure translates, or does not translate, to the nominal term structure then crucially depends on the stochastic properties of inflation—in particular, inflation persistence—and therefore, ultimately, upon the nature of the underlying monetary regime. If inflation is, in the limiting case, a pure random walk, so that innovations are entirely permanent, a shock to inflation today shifts expected inflation at all horizons by an identical amount, thus leaving the nominal yield curve, for a given real yield curve, unaffected. In this case the predictive content of the spread intrinsic to the real yield curve translates one-to-one to the nominal yield curve. If, on the other hand, inflation has little persistence—as it was the case under metallic standards, and currently is the case for several inflation-targeting countries[6]—a shock to inflation uniquely increases short-term inflation expectations, leaving instead long-term expectations unaffected, and therefore, for a given *real* yield curve, by increasing short rates and leaving long rates unchanged, it twists the *nominal* yield curve, thus 'blurring' the informational content of the real curve.

Which, if either, of the two explanations is correct? Or might it be the case that they are both wrong? As stressed by Bordo and Haubrich (2004, page 3),

> [w]hether the yield curve's ability to predict [output growth] emerges as a general property of the American business cycle or depends sensitively on the structure of the economy, financial markets, and monetary policy seems an obvious question. Particularly since a subtext of the yield curve's

[5] See in particular Harvey (1988), Plosser and Rouwenhorst (1994), and Bordo and Haubrich (2004).

[6] See in particular Bordo and Schwartz (1999). The essential white-noise character of U.S. inflation under metallic standards has been extensively documented, e.g., by Shiller and Siegel (1977) and Barsky (1987). Benati (2006) documents how, in the United Kingdom, inflation persistence has been entirely absent under both metallic standards and the current inflation-targeting regime, while Benati (2007b) shows how persistence is currently low to non-existent in three other inflation-targeting countries, Canada, New Zealand, and Sweden.

> predictive ability has been the instability of its relationship with output growth, looking at a long time series seems warranted. *A broader historical perspective may also shed some light on the reasons behind the yield curve's ability to predict future output—for example, one simply cannot ascribe twists in the yield curve during the 1880s to an FOMC ratcheting up short-term rates.* (emphasis added)

Bordo and Haubrich hit upon a crucial point: if one wants to discriminate between the two previously discussed theories, (s)he has to examine sample periods during which the two theories are not observationally equivalent, and, as a simple matter of logic, the chance of finding such periods increases (1) with the length of the sample period considered; and (2) with the variety of monetary arrangements examined, for the simple reason that, in both explanations, monetary policy plays, either directly or indirectly, a crucial role. Under this respect, both the recent experience of inflation targeting countries, and the U.S. and U.K. experience under metallic standards, should be regarded as potentially valuable, as they should provide sufficient variation in the monetary policy rule to discriminate between the two rival explanations.

It is, therefore, quite surprising that—with the single exception of Kessel (1965)—Bordo and Haubrich (2004) is the *only* paper to have ever attempted a systematic investigation of (changes over time in) the predictive content of the spread based on long spans of data. Although conceptually pathbreaking, the work of Bordo and Haubrich (2004) suffers however, in our opinion, from a crucial drawback, in that it only investigates whether the yield spread contains information beyond that already encoded in lagged output growth, being therefore by definition silent on the crucial issue of whether the spread contains information which is not already encoded in other macroeconomic variables, first and foremost measures of the monetary policy stance such as short-term interest rates. Such a problem is unfortunately quite common in the literature, with several papers only having *one* regressor, the spread,[7] and another group of papers having, like Bordo and Haubrich (2004), only one additional regressor beyond the spread, the lagged dependent variable.[8]

1.1 Issues addressed in the present work

Based on data for the United States and the United Kingdom, since the Gold Standard era, and the Eurozone, Canada, and Australia over the post-WWII period (Figures 1-3 show the raw data used in this paper), in this paper we use Bayesian time-varying parameters VARs with stochastic volatility along the lines of Cogley and Sargent (2005), first, to re-examine the crucial issue in this literature:

[7] See e.g. Bernard and Gerlach (1996).

[8] See e.g. Bonser-Neal and Morley (1997) and Haubrich and Dombrovsky (1996).

- *'Once controlling for the information contained in other regressors—in particular, in measures of the monetary policy stance, such as the short rate—does the yield spread still contain information useful for predicting output growth?'*

Although, in principle, a proper attempt to provide an answer to this question would require an examination of every available macroeconomic indicator, in this paper we limit ourselves to inflation, output growth, and a short rate. There are two reasons for this. First, as a matter of practicality: the time-varying Bayesian methodology used herein is extremely computer intensive, to the point that expanding the benchmark dataset beyond four variables would become prohibitively cumbersome.[9] Second, it can reasonably be argued that these three variables provide a sufficiently exhaustive statistical summary of the properties of any advanced economy,[10] so that they should provide a reasonably robust benchmark against which to measure the informational content of the spread. We then tackle three additional issues:

- *'Has the marginal predictive content of the spread remained broadly unchanged over time, or has it exhibited significant time-variation?'*

- *'In case it has changed over time, do such changes bear any clear relationship with changes in the underlying monetary regime?'*

- *'Does our evidence clearly falsify/reject either of the two explanations we previously discussed in section 1.1?'*

1.2 Key results

Based on Stock and Watson's Classical time-varying parameters median-unbiased estimation (henceforth, TVPMUB) methodology, we detect strong evidence of random-walk time-variation, against the null of time-invariance, in output growth regressions for all countries, with comparatively large median-unbiased estimates of the overall extent of parameter drift. These results provide strong *prima facie* evidence—but, it is important to stress, only *prima facie* evidence—that both ouput growth's overall extent of predictability, and the marginal predictive power of individual regressors for output growth, may have changed over time. A proper assessment of both issues necessarily calls, however, for multivariate methods, as it requires a (time-varying) estimate of the entire spectral density matrix of the data.

Moving to a multivariate context, overall our evidence does not provide full support for *either* of the two previously discussed explanations why the yield spread may

[9] This is known in the literature as the 'curse of dimensionality', and is extensively discussed, e.g., by DelNegro (2003). According to our own experience, even the algorithm proposed by DelNegro (2003) only provides a partial solution to this problem.

[10] Cogley and Sargent (2002), for example, justify the inclusion, in their time-varying VAR, of inflation, a short rate, and (the logit of) the unemployment rate, precisely along these lines.

contain predictive power for output growth. On the one hand, the 'real yield curve' explanation is contradicted by the fact that in both the United States, the United Kingdom, and Canada over the post-WWII era, the broad decrease in inflation persistence we identify for all the three countries over the second part of the sample was not accompanied by a corresponding decrease in the marginal predictive content of the spread compared to the information already encoded in past output growth. The monetary policy-based explanation, on the other hand, appears incompatible with the fact that, for example, results for the United States during both the interwar and the post-WWII periods clearly point towards several periods during which the spread exhibited predictive power for output growth over and above that already encoded in the short rate. In particular, during both the Volcker recession, and the 2000-2001 one, the spread clearly appears to have possessed additional information compared with that contained in the simplest measure of the monetary policy stance.

The paper is organised as follows. The next section presents preliminary evidence on the presence of (random-walk) time-variation in univariate regressions for output growth, based on the Stock-Watson time-varying parameters median-unbiased estimation methodology. Section 3 describes the Bayesian methodology we use to estimate time-varying parameters VARs with stochastic volatility, while section 4 discusses the methodology we use to compute, at each point in time, measures of overall and marginal predictability at the various horizons. Section 5 presents the empirical evidence. Section 6 discusses the implications of our findings for the two explanations discussed in the introduction, and proposes yet another conjecture. Section 7 concludes.

2 Searching for Time-Variation in GDP Growth Regressions

Before delving into the time-varying multivariate analysis of the next four sections, a useful preliminary step is to provide evidence of instability in real GDP growth regressions. In this section we therefore apply the Classical TVP-MUB methodology due to Stock and Watson (1996) and Stock and Watson (1998) to test for the presence of random-walk time-variation in output growth regressions, and to estimate its extent.

Our preference for the Stock-Watson methodology over a currently popular alternative, structural breaks tests,[11] has to do with its greater robustness to uncertainty concerning the specific form of time-variation present in the data. While time-varying parameters models are well known for being capable of successfully tracking processes subject to structural breaks, both Cogley and Sargent (2005) and Benati (2007a) have shown break tests to possess a sometimes remarkably low power when the true data-generation process (henceforth, DGP) is characterised by random walk time

[11] See, e.g., Bai and Perron (1998) and Bai and Perron (2003). For an application of break tests to the issue at hand, see e.g. Estrella, Rodrigues, and Schich (2003).

variation.[12] To put it differently, break tests can be expected to perform well *if and only if* the DGP is subject to discrete structural breaks, while TVP models can be expected to perform well under *both* scenarios.

The regression model we consider in this section is

$$y_t = \mu + \alpha(L)y_{t-1} + \beta(L)\pi_{t-1} + \gamma(L)r_{t-1} + \delta(L)s_{t-1} + u_t \equiv \theta' Z_t + u_t \qquad (1)$$

where where y_t, π_t, r_t, and s_t are output growth, inflation, a short-term rate, and the nominal yield spread, respectively (Appendix A provides a detailed description of the dataset);[13] $\alpha(L)$, $\beta(L)$, $\gamma(L)$, and $\delta(L)$ are lag polynomials; θ=[μ, $\alpha_0(L)$, ..., $\delta_0(L)$]$'$ and Z_t=[1, y_{t-1}, ..., s_{t-p}]$'$. In what follows we set the lag order of all polynomials to p=4.

Letting θ_t=[μ_t, $\alpha_t'(L)$, ..., $\delta_t'(L)$]$'$, the time-varying parameters version of (1) is given by

$$y_t = \theta_t' Z_t + u_t \qquad (2)$$

$$\theta_t = \theta_{t-1} + \eta_t \qquad (3)$$

with η_t *iid* $N(0_{4p+1}, \lambda^2\sigma^2 Q)$, with 0_{4p+1} being a (4p+1)-dimensional vector of zeros; σ^2 being the variance of u_t; Q being a covariance matrix; and $E[\eta_t u_t]$=0. Following Nyblom (1989) and Stock and Watson (1996, 1998), we set Q=$[E(Z_t Z_t')]^{-1}$. Under such a normalisation, the coefficients on the transformed regressors, $[E(Z_t Z_t')]^{-1/2}Z_t$, evolve according to a (4p+1)-dimensional standard random walk, with λ^2 being the ratio between the variance of each 'transformed innovation' and the variance of u_t.[14] In what follows we estimate the matrix Q as in Stock and Watson (1996) as

$$\hat{Q} = \left[T^{-1}\sum_{t=1}^{T} z_t z_t'\right]^{-1}. \qquad (4)$$

Our approach closely follows Stock and Watson (1996, Section 2). The point of departure is the OLS estimate of (1), conditional on which we compute the residuals, $\hat{u}_t$, and the estimate of the innovation variance, $\hat{\sigma}^2$, and we perform an *exp*- and a

[12] Cogley and Sargent (2005) report the following values for the power of the test for the equations for the nominal rate, unemployment, and inflation in their Bayesian time-varying parameters VAR. Andrews (1993)'s *sup*-LM test: 0.136, 0.172, and 0.112. Nyblom (1989)-Hansen (1992) test: 0.076, 0.170, 0.086. Andrews (1993)'s *sup*-Wald test: 0.173, 0.269, 0.711. Conditional on taking the estimated Stock-Watson TVP-MUB models for labor productivity growth as DGPs, Benati (2007a) reports values of the power of the the tests for breaks in the mean between 0.319 and 0.374 for Andrews and Ploberger (1994)'s *exp*-Wald statistic, and between 0.310 and 0.390 for Bai and Perron (1998)'s *WDmax* test statistic.

[13] In principle, we could add more regressors to (1). The main reason why we are only considering these four variables is for consistency with the next sections, in which computational constraints (on this, see footnote 9) compel us to work with, at most, four series.

[14] To be precise, given that the Stock-Watson methodology is based on local-to-unity asymptotics, λ is actually equal to the ratio between τ, a small number which is fixed in each sample, and T, the sample length.

sup-Wald joint test for a single break at an unknown point in the sample in μ and in the sums of the α's, β's, γ's, and δ's, using the Andrews (1991) HAC covariance matrix estimator to control for possible autocorrelation and/or heteroskedasticity in the residuals. We then build up the empirical distribution of the test statistic as in Stock and Watson (1996, Section 2.4), by considering a 30-point grid of values for λ over the interval [0, 0.1], let's call it Λ. For each $\lambda_j \in \Lambda$ we compute the corresponding estimate of the covariance matrix of η_t as $\hat{Q}_j=\lambda_j^2\hat{\sigma}^2\hat{Q}$, and conditional on $\hat{Q}_j$ we simulate model (2)-(3) 1,000 times, drawing the pseudo innovations from pseudo random *iid* $N(0, \hat{\sigma}^2)$. For each simulation, we then compute an *exp*- and a *sup*-Wald test[15] thus building up their empirical distribution. We compute the median-unbiased estimate of λ as that particular value of λ_j for which the median of the simulated empirical distribution of the test is closest to the test statistic previously computed based on the actual data. In case the *exp*- or *sup*-Wald test statistics computed based on the actual data are greater than the corresponding medians of the empirical distributions conditional on λ_j=0.1, we add one more step to the grid, and we estimate λ as 0.10345. Finally, we compute the *p*-value based on the empirical distribution of the test conditional on λ_j=0.

Table 1 reports the results. Two main findings emerge from the table. First, we detect very strong evidence of random-walk time-variation for all countries and sample periods.[16] Second, in the vast majority of the cases, the MUB estimates of λ clearly suggest the DGPs to be characterised by a significant extent of random-walk drift. Overall, these results provide therefore strong *prima facie* evidence that the forecasting power of individual regressors for output growth might have changed over the sample periods. A proper assessment of changes over time both in output growth's overall predictability, and in the marginal predictive content of the spread, necessarily calls, however, for time-varying multivariate methods, as it requires a time-varying estimate of the entire spectral density matrix of the data. We therefore move to the next section, in which we describe the multivariate Bayesian methodology we use to characterise the evolution of the stochastic properties of the series under investigation.

[15] Quite obviously, without computing the Andrews (1991) HAC correction ...

[16] For the United Kingdom we consider only the post-WWII era, as, first, data for the Gold Standard period are annual, thus preventing the application of a 'data hungry' methodology like the Stock-Watson TVP-MUB; and second, although for the interwar era we do have quarterly data, the sample period is simply too short to allow us to obtain reliable results.

3 A Time-Varying Parameters VAR With Stochastic Volatility

Following Cogley and Sargent (2005), we work with the following time-varying parameters VAR(p) model:

$$Y_t = B_{0,t} + B_{1,t}Y_{t-1} + ... + B_{p,t}Y_{t-p} + \epsilon_t \equiv X_t'\theta_t + \epsilon_t \tag{5}$$

where the notation is obvious, with Y_t being a vector of endogenous variables to be discussed below. Consistent with the vast majority of the literature—and mostly for reasons of computational feasibility—the lag order is set to p=2. Following, e.g., Cogley and Sargent (2002), Cogley and Sargent (2005), Primiceri (2005), and Canova and Gambetti (2005), the VAR's time-varying parameters, collected in the vector θ_t, are postulated to evolve according to

$$p(\theta_t \mid \theta_{t-1}, Q) = I(\theta_t)\ f(\theta_t \mid \theta_{t-1}, Q) \tag{6}$$

with $I(\theta_t)$ being an indicator function rejecting unstable draws—thus enforcing a stationarity constraint on the VAR[17]—and with $f(\theta_t \mid \theta_{t-1}, Q)$ given by

$$\theta_t = \theta_{t-1} + \eta_t \tag{7}$$

with $\eta_t \sim N(0, Q)$. The VAR's reduced-form innovations in (5) are postulated to be zero-mean normally distributed, with $\mathrm{Var}(\epsilon_t) \equiv \Omega_t$ which, following Cogley and Sargent (2005), we factor as

$$\Omega_t = A^{-1}H_t(A^{-1})' \tag{8}$$

with H_t and A defined as

$$H_t \equiv \begin{bmatrix} h_{1,t} & 0 & ... & 0 \\ & h_{2,t} & ... & 0 \\ ... & ... & ... & ... \\ 0 & 0 & ... & h_{N,t} \end{bmatrix} \qquad A \equiv \begin{bmatrix} 1 & 0 & ... & 0 \\ \alpha_{2,1} & 1 & ... & 0 \\ ... & ... & ... & ... \\ \alpha_{N,1} & \alpha_{N,2} & ... & 1 \end{bmatrix} \tag{9}$$

The $h_{i,t}$'s are postulated to evolve according to geometric random walks,

$$\ln h_{i,t} = \ln h_{i,t-1} + \nu_{i,t} \tag{10}$$

with $\mathrm{Var}(\nu_{i,t}) \equiv \sigma_i^2$. For future reference, we define $h_t \equiv [h_{1,t}, h_{2,t}, ..., h_{N,t}]'$, $\sigma^2 \equiv [\sigma_1^2, \sigma_2^2, ..., \sigma_N^2]'$, and $\alpha \equiv [\alpha_{2,1}, \alpha_{3,1}, ..., \alpha_{N,N-1}]'$. Finally, and uniquely for the sake of simplicity, we follow Cogley and Sargent (2005) in assuming independence between the innovations to the random-walk coefficients and the VARs' reduced-form shocks:

$$\begin{bmatrix} \epsilon_t \\ \eta_t \end{bmatrix} \sim N(0, V_t), \text{ with } V_t = \begin{bmatrix} \Omega_t & 0 \\ 0 & Q \end{bmatrix} \tag{11}$$

[17] The reason for imposing such a constraint is in order to be able to Fourier-transform the time-varying VAR at each point in time—see section 4.

We estimate (5)-(11) *via* Bayesian methods. Appendix B discusses our choices for the priors, and the Markov-Chain Monte Carlo procedure we use to simulate the posterior distribution of the hyperparameters and the states conditional on the data.

4 Measuring the Marginal Predictive Content of the Yield Spread for Output Growth

Since we are interested in assessing the marginal predictive content of the spread for output growth compared to *both* the predictive power uniquely encoded in past output growth, *and* the information contained in past values of output growth, inflation, and the short-term rate, we consider two 'benchmark' models, with $Y_t \equiv y_t$ and $Y_t \equiv [y_t, \pi_t, r_t]'$, respectively. We then have the corresponding 'augmented' models with $Y_t \equiv [y_t, s_t]'$ and $Y_t \equiv [y_t, \pi_t, r_t, s_t]'$, respectively. We define the (time-varying) *marginal* predictive content of the spread for output growth as the difference between the two multivariate R^2s' of output growth implied by the augmented and by the benchmark models, respectively.

We compute the time-varying multivariate R^2 statistics for output growth at horizon k implied by each of the estimated TVP VARs along the lines of Cogley (2005). In order to establish notation, let

$$Y_t^* = F_t Y_{t-1}^* + \epsilon_t^* \tag{12}$$

be the companion form of (5), with $Y_t^* \equiv [Y_t', Y_{t-1}']$, $\epsilon_t^* \equiv [\epsilon_t', 0_{1\times N}]'$, where N is the dimension of Y_t, and $\text{Var}(\epsilon_t^*) \equiv \Omega_t^*$. If the VAR were time-invariant—so that θ_t=θ, Ω_t=Ω, and therefore F_t=F, and Ω_t^* =Ω^*—a multivariate R^2 statistic at horizon k could be trivially computed based on (12). Given that both the VAR's coefficients and its covariance matrix are time-varying such an approach is however unfeasible. In what follows we, therefore, compute the time-varying R^2 statistics by taking account of the uncertainty originating from future time-variation in both θ_t and Ω_t *via* the following Monte Carlo integration procedure.

Let $F_{t|T}$ and $\Omega_{t|T}^*$ be the two-sided estimates of F_t and Ω_t^* produced by the Gibbs sampler. For each t = 1, 2, ..., T, and for each of the 2,000 iterations of the Gibbs sampler which constitute the ergodic distribution, we start by simulating $F_{t|T}$ and $\Omega_{t|T}^*$ into the future[18] from t+1 to t+k, thus getting simulated paths $F_{t+1|T}$, ..., $F_{t+k|T}$

[18] It is important to stress that, for the entire exercise to be *exactly* correct, the VARs should be estimated for recursive samples. This would allow us to simulate into the future the proper objects—$F_{t|t}$ and $\Omega_{t|t}^*$—instead of $F_{t|T}$ and $\Omega_{t|T}^*$. Given the staggering computational burden associated with re-estimating the model for every single quarter (the reader should keep in mind that here we have five countries, and that for the U.S. we have three different sample periods ...), we have decided to perform the exercise based on the smoothed (i.e, two-sided) output of the Gibbs sampler conditional on the full sample. This implies that our results should be regarded as approximations to the authentic out-of-sample objects that would result from a proper recursive estimation. It

and $\Omega^*_{t+1|T}$ ..., $\Omega^*_{t+k|T}$. With $Y^*_{t+k|T}$ equal to

$$Y^*_{t+k|T} = \epsilon^*_{t+k|T} + F_{t+k|T}\epsilon^*_{t+k-1|T} + F_{t+k|T}F_{t+k-1|T}\epsilon^*_{t+k-2|T} + \tag{13}$$

$$+ ... + F_{t+k|T}F_{t+k-1|T}...F_{t+3|T}F_{t+2|T}\epsilon^*_{t+1|T} + F_{t+k|T}F_{t+k-1|T}...F_{t+3|T}F_{t+2|T}F_{t+1|T}Y^*_{t|T}$$

the forecast error for $Y^*_{t+k|T}$ conditional on $Y^*_{t|T}$ and on the simulated path $F_{t+1|T}$, ..., $F_{t+k|T}$ is given by[19]

$$\epsilon^*_{t+k|T} + F_{t+k|T}\epsilon^*_{t+k-1|T} + F_{t+k|T}F_{t+k-1|T}\epsilon^*_{t+k-2|T} + \tag{14}$$

$$+ ... + F_{t+k|T}F_{t+k-1|T}...F_{t+3|T}F_{t+2|T}\epsilon^*_{t+1|T}$$

Based on $F_{t+k|T}$ and $\Omega^*_{t+k|T}$ we can trivially compute the time-varying spectral density matrix of $Y^*_{t+k|T}$, $S_{Y^*_{t+k|T}}(\omega)$, as

$$S_{Y^*_{t+k|T}} = \frac{1}{2\pi}\left[I - F_{t+k|T}(e^{-i\omega})\right]^{-1}\Omega^*_{t+k|T}\left\{\left[I - F_{t+k|T}(e^{i\omega})\right]^{-1}\right\}' \tag{15}$$

By the same token, based on (14) we can compute the time-varying spectral density matrix of the forecast error at horizon k as

$$S_{\epsilon^*_{t+k|T}} = \frac{1}{2\pi}\left\{\Omega^*_{t+k|T} + \sum_{j=1}^{k-1}\left[\prod_{h=1}^{j} F_{t+k+1-h|T}\left(e^{-i\omega}\right)\right]\Omega^*_{t+k-j|T}\left[\prod_{h=1}^{j} F_{t+k+1-h|T}\left(e^{i\omega}\right)\right]'\right\} \tag{16}$$

The uncertainty originating from future time-variation in θ_t and Ω_t can then be integrated out by repeating this procedure N times, and by averaging across the $S^j_{Y^*_{t+k|T}}(\omega)$ and $S^j_{\epsilon^*_{t+k|T}}(\omega)$, j = 1, 2, ..., N. For future reference, let $\tilde{S}_{Y^*_{t+k|T}}(\omega)$ and $\tilde{S}_{\epsilon^*_{t+k|T}}(\omega)$ be the averages of $S^j_{Y^*_{t+k|T}}(\omega)$ and $S^j_{\epsilon^*_{t+k|T}}(\omega)$ across the N simulations. For each t = 1, 2, ..., T, and for each of the 2,000 iterations of the Gibbs sampler which constitute the ergodic distribution, we perform Monte Carlo integration based on N=100.

Based on $\tilde{S}_{Y^*_{t+k|T}}(\omega)$ and $\tilde{S}_{\epsilon^*_{t+k|T}}(\omega)$ we then compute the time-varying multivariate R^2 for output growth at horizon k implied by each of the TVP-VARs along the lines

is worth stressing that precisely because of the enormous computational burden associated with recursive estimation of Bayesian time-varying parameters VARs, computing predictability measures at each point in time, based on the two-sided output of the Gibbs sampler, is quite common in the literature—see e.g. Cogley and Sargent (2002, section 3.3).

[19] An intuitive way of understanding why (14) is indeed the forecast error for $Y^*_{t+k|T}$ is the following:- Suppose you are at time t—so that you have $Y^*_{t|T}$ and F_t—and you are asked to produce a forecast for $Y^*_{t+k|T}$. How would you proceed? Quite obviously, you would first simulate F_t into the future, thus getting the simulated path F_{t+1}, ..., F_{t+k}, and then, based on that, you would compute the forecast as $F_{t+k}F_{t+k-1}...F_{t+3}F_{t+2}F_{t+1}Y^*_{t|T}$. It is important to stress that this way of proceeding is exactly correct because of the crucial assumption of orthogonality between ϵ_t and η_t—see (11).

of Cogley (2005), as[20]

$$R^2_{y,t+k|T} = 1 - \frac{\tilde{\Omega}_{y,t+k|T}}{\sigma^2_{y,t+k|T}} \tag{17}$$

where

$$\sigma^2_{y,t+k|T} = \int_{-\pi}^{\pi} f_{y,t+k|T}(\omega) d\omega \tag{18}$$

is $y_{t+k|T}$'s estimated overall time-varying variance; $f_{y,t+k|T}(\omega)$ is the time-varying estimate of $y_{t+k|T}$'s spectral density, based on $\tilde{S}_{Y^*_{t+k|T}}(\omega)$; and

$$\tilde{\Omega}_{y,t+k|T} = 2\pi \exp\left\{\frac{1}{2\pi}\int_{-\pi}^{\pi} \ln\left[f_{\epsilon^*_{t+k|T}}(\omega) d\omega\right]\right\} \tag{19}$$

is the time-varying variance of the forecast error at horizon k, computed based on Kolmogorov's formula, with $f_{\epsilon^*_{t+k|T}}(\omega)$ being the time-varying estimate of its spectral density.

A final, important point to be stressed is the following. Although we are here investigating the marginal predictive content of the spread at horizon k by simulating the VARs into the future, *our results ought to be regarded, in terms of informational content, as in-sample ones*, as a (pseudo) out-of-sample analysis would have required recursive estimation of the TVP-VARs for each single quarter, conceptually in line with (e.g.) Stock and Watson (1999). As we have already mentioned in footnote 18, the strongest possible justification for not doing this is the staggering computational burden associated with recursive estimation of the time-varying parameters VARs for each country and for each single quarter.

5 Empirical Evidence

5.1 Evidence for the United States

Figure 4 shows, for the three sample periods, the medians of the distributions, and the 16th and 84th percentiles, for the time-varying overall one-quarter-ahead R^2s' implied by the bivariate VARs for y_t and s_t (in the top row), together with the corresponding figures for the marginal predictive content of the spread (in the bottom row), computed as the difference between the two R^2's implied by the bivariate and by the univariate models, respectively. Focussing on the median estimates,[21] a key finding emerging from Figure 4 is the comparatively large marginal predictive content

[20] The only difference between (5) and the analogous expression found in Cogley (2005) is that he uniquely focusses on one-step-ahead predictability (i.e., he only considers k=1).

[21] Our focus on median estimates, and our disregard for the associated extent of uncertainty, is motivated by the fact that, as it is well known, time-varying parameters models are intrinsically characterised by a comparatively large extent of econometric uncertainty.

of the spread (henceforth, MPCS)—over and above the information already encoded in past output growth—under the Classical Gold Standard, starting just shy of 10% around mid-1880s, peaking around 18% around the mid-1890s, and then decreasing rapidly towards the end of the century, fluctuating, until the outbreak of WWI, between 2% and 5%. While our results are fully compatible with those of Bordo and Haubrich (2004),[22] our use of time-varying parameters techniques allows us to identify exactly the precise extent of the MPCS at each point in time, thus showing how their finding results from 'lumping together' two different sub-periods: a former one stretching until the end of the 19th century, characterised by historically large values of the MPCS, and a latter one starting around the turn of the century and extending until the outbreak of WWI, during which the spread exhibited a markedly lower predictive content. By contrast, during the interwar era the MPCS clearly decreases, to the point that, based on median estimates, it is even estimated to be negative, although statistically indistinguishable from zero even based on the one standard deviation percentiles. Given the implausibility, on logical grounds, of a negative MPCS, the correct interpretation of such result is however that, during that period, the MPCS was essentially zero. This will have to be kept in mind throughout the rest of the paper, as, as we will see, in a few cases the MPCS will be estimated to have been negative. The key point to stress here is that the object we are dealing with—the MPCS—is, within the present context, stochastic, so that, even in those cases in which we uniquely report, for the sake of simplicity, a median estimate, it would never be possible to reject at conventional significance levels the null that the negative MPCS is equal to zero. Finally, results for the post-WWII period show how the MPCS for output growth previously documented by many authors clearly appears to have been associated with the chairmanship of Paul Volcker, while it was virtually *nil* before Volcker, and, in line with Dotsey (1998) and Estrella, Rodrigues, and Schich (2003), it appears to have disappeared under Alan Greenspan, with the possible exception of the period around the 2000-2001 recession, when, as Figure 11 clearly shows, it appears to have increased quite significantly.

Turning to the marginal predictive content exhibited by the spread at the different horizons, the top row of Figure 6 shows the medians of the distributions of the time-varying overall 1-, 4-, and 8-quarters ahead R^2s' implied by the bivariate VAR for y_t and s_t, while the top row of figure 7 shows the corresponding figures for the MPCS.[23] Several findings emerge from Figure 7. First, the peculiar time pattern we identified at the one-quarter horizon for the Gold Standard era all but disappears at longer horizons, with near-perfectly flat median estimates of the MPCS both one

[22] See Bordo and Haubrich (2004, Tables 2 and 3).

[23] While the one-step-ahead objects can be computed directly based on the output of the Gibbs sampler—i.e., without the need to perform Monte Carlo integration—the corresponding objects for $k>1$ cannot. Given the remarkable computational intensity of the Monte Carlo integration procedure described in section 4, all the results for $k>1$ reported in this paper (see Figures 6, 7, 10, and 11) are computed once out of every four quarters.

and two years ahead. Further, while the spread still appears to possess some limited forecasting power at the one-year horizon, at the two-year horizon the MPCS appears to have entirely vanished. Second, for the interwar period, the slightly negative MPCS we identified one-quarter-ahead does not carry over to longer horizons, with some modest marginal predictive content identified at both the one- and the two-year horizons. Third, for the post-WWII era, the MPCS increases from the 1- to the 4-quarter-ahead horizon by a broadly similar extent over the entire sample period, and then markedly decreases at the two-year horizon.

Let's now turn to the more interesting issue of the additional predictive content of the spread over and above the information already contained, not only in past output growth, but also in past values of inflation and of a short-term rate. Figure 8, and the top rows of Figures 10 and 11, show the same objects shown in Figure 4 and in the top rows of Figures 6 and 7, respectively, but this time with the 'benchmark' and 'augmented' VARs estimated for $Y_t \equiv [y_t, \pi_t, r_t]'$ and $Y_t \equiv [y_t, \pi_t, r_t, s_t]'$, respectively. Focusing, again, on median estimates, three findings emerge from Figure 8. First, under the Gold Standard, the inclusion of inflation and a short rate as additional regressors essentially 'kills off' the MPCS for output growth, with the only exception of the very last years of the sample, for which some modest additional predictive power, around 5%, still remains. Second, during the interwar period the MPCS fluctuated roughly between -10% and 10%, being consistently positive during the years leading up to the Great Crash, turning negative around the time of the Great Depression, and returning again mostly in positive territory following the abandonment of the gold parity in April 1933.[24] Finally, in the post-WWII era we confirm, for the 1-quarter-ahead horizon, the previous finding that the MPCS appears to have essentially been associated with the chairmanship of Paul Volcker. Interestingly, however, we detect clear evidence of the reappearance of some predictive power around the time of the 2000-2001 recession, with the median marginal R^2 increasing to a peak of about 8%. Turning to longer horizons, the most interesting results relate, once again, to the post-WWII period. In particular, the MPCS at the one-year horizon has consistently been greater than at the one-quarter horizon over the entire sample period, and it clearly appears to have been present not only during the Volcker chairmanship, but also during previous years. During the chairmanship of Alan Greenspan, the MPCS at the one-year horizon has markedly decreased, but, consistently with our findings for the one-quarter horizon, it has significantly increased around the time of the 2000-2001 recession. Finally, for the two-year horizon the MPCS appears to have been virtually *nil*, and often even slightly negative, over the entire sample period, with the only exception of the 2000-2001 recession, during which, again, it increased to a peak of 5%.

[24] See Eichengreen (1992).

5.2 Evidence for other countries

Based on time-invariant VARs Tables 2 and 3[25] show, for the United Kingdom during the Gold Standard and the interwar era,[26] respectively, both output growth's overall predictability, and the marginal predictive content of the yield spread computed based on the two previously discussed benchmark models, the univariate one, and the one with $Y_t \equiv [y_t,\ \pi_t, r_t]'$. For both overall predictability and the MPCS we report two sets of results, a first one based on the same frequency-domain methodology described in section 4,[27] and a second one, time-domain-based, in which the R^2s for both the benchmark and the 'extended' models are computed based on OLS regressions. Figures 5 and 9, and the bottom rows of Figures 6, 7, 10, and 11, on the other hand, show, for the Eurozone, the United Kingdom, Canada, and Australia, the same objects shown for the United States in Figures 4 and 8, and in the top rows of Figures 6, 7, 10, and 11.

Starting with the United Kingdom, results for the Gold Standard consistently point towards *no* MPCS at either horizon, based on either of the two benchmark models, and based on either time- or frequency-domain methods. Only in one case—at the one-year horizon, based on the univariate benchmark, and time-domain methods—there is limited evidence of some MPCS, but the figure, 3.8%, is still quite low. The picture for the interwar era is more complex, with, first, and quite surprisingly, consistent evidence of MPCS at the two-year horizon based on either benchmark and on either time- or frequency-domain methods; and second, much weaker evidence of MPCS at shorter horizons.

Turning to the post-WWII era, and focusing, again, on median estimates, the results based on the univariate benchmark point towards a consistently negative (i.e, zero) MPCS at the one-quarter horizon over the entire sample period, a marked increase to about 10% at the one-year horizon over the entire sample, and the virtual disappearance of any marginal predictive power at the two-year horizon. Results based on the alternative benchmark model, on the other hand, paint a radically

[25]Results for the Gold Standard (based on annual data) are based on VARs with one lag. As for the interwar period, in order to exploit the maximal amount of information possible, we estimate the VARs at the monthly frequency, setting the lag order to six. Alternative sets of results for the two periods based on VARs with two and twelve lags, respectively, are qualitatively the same, and are available upon request.

[26]For both regimes/periods we eschew the time-varying parameters VARs, which we use for all other countries, and for the post-WWII United Kingdom (see below), because of the high 'data intensity' typical of time-varying parameters models. As for the interwar period, we only have, indeed, 68 quarterly observation (instead of the 91 for the U.S.), and given the necessity of using part of the sample as 'pre-sample', to get the priors, we have judged that the reliability of results from a time-varying parameters VAR would, in the end, probably turn out to be too low. As for the Gold Standard era the two sample periods are, in principle, sufficiently long, but unfortunately the data are only available at the annual frequency, so that the number of observations is, again, comparatively small.

[27]Given that in the present case we are estimating time-invariant VARs, however, results reported in Tables 2 and 3 are, quite obviously, not based on Monte Carlo integration methods.

different picture, with essentially no MPCS over the first part of the sample, a clear watershed corresponding to the United Kingdom's entry into the Exchange Rate Mechanism of the European Monetary System (October 1990), and a positive MPCS since then at both the 1- and the 4-quarter horizon.

Turning to other countries,

- for the Eurozone, results based on the univariate benchmark point towards a negligible MPCS for output growth at any horizon, while those based on the 'extended' benchmark suggest that, at least at the one-quarter horizon at the very beginning of the sample, and at the one-year horizon both at the beginning of the sample and around the second half of the 1990s, a significant MPCS has been present.

- As for Australia, results based on the univariate benchmark point towards a MPCS around 9% at the very beginning of the sample, decreasing monotonically over the following years, and then stabilising, after the introduction of inflation targeting, around 3%. In contrast with, e.g., the United Kingdom, the MPCS based on the univariate benchmark appears to have been monotonically declining with the forecasting horizon during the entire sample period. Results based on the alternative benchmark, on the other hand, point towards an essentially negligible MPCS at all horizons, with the exception of the one-year horizon towards the middle of the sample.

- Finally, results for Canada point towards a striking difference between the results based on the two benchmarks, with those based on the univariate one suggesting a remarkably high MPCS at both the 1- and the 4-quarter horizon over the entire sample period, and those based on the extended benchmark pointing towards marked fluctuations at all horizons.

Having documented time-variation in the MPCS in the five countries/economic areas since the Gold Standard era, let's now turn to the more interesting question of what implications our findings have for the two previously discussed theories for why the spread might contain information useful for predicting output growth.

6 Interpreting the Evidence

6.1 Implications for the 'real yield curve' explanation

As mentioned in the introduction, according to Harvey's (1988), theory whether that the MPCS encoded in the real yield curve does, or does not, translate to the nominal yield curve crucially depends on inflation persistence. Following Bordo and Haubrich (2004), a necessary preliminary step in order to be able to assess the implications of our findings for the real yield curve explanation is, therefore, measuring (time-variation in) the extent of inflation persistence. Figure 12 shows, for the United States during the Classical Gold Standard, the interwar period, and the post-WWII era, and for the other four countries over the post-WWII period, the median estimate and the 16th and 84th percentiles of the distribution of the normalised spectrum of inflation at ω=0, based on the TVP VARs for $Y_t \equiv [y_t, \pi_t, r_t, s_t]'$.[28] Results for the United States, first, confirm the well-known white noise character of inflation during the Gold Standard, documented e.g. by Shiller and Siegel (1977) and Barsky (1987); second, the data point towards some mild increase in persistence during the interwar years, especially, during the Great Depression episode, when inflation remained consistently negative for several years; and finally, broadly confirm, for the post-WWII era, the finding of Cogley and Sargent (2002, 2005) of high persistence around the time of the Great Inflation, and of a significant decline after the Volcker stabilisation. In line with Benati (2004), Benati (2006) and Cogley, Morozov, and Sargent (2003), results for the United Kingdom confirm the broad picture of comparatively high persistence during the period between the floating of the pound and the U.K.'s entry into the Exchange Rate Mechanism of the European Monetary System, and of a lower persistence since then. As for other countries, Canada exhibits a peak around the time of the Great Inflation, and a significant fall since then, and the Eurozone a systematic increase over the sample period, with an overall high estimated persistence.[29] Finally, as for the United Kingdom during the Gold Standard and the interwar period, Table B of Benati (2006) reports, for the two series used herein, Hansen (1999) 'grid bootstrap' median-unbiased estimates of the sum of the AR coefficients, and 90%-coverage confidence intervals, equal to 0.05 [-0.13; 0.22] and 0.37 [-0.05; 0.80] respectively. In line with the United States, these results point towards no and, respectively, some very mild persistence during the two periods.

Some of these results are at odds with the previously documented pattern of variation in the MPCS over and above the information already encoded in past output growth. For example,

- in both the United States, the United Kingdom, and Canada over the post-WWII era, the broad decrease in inflation persistence in the second part of the sample, compared to the first half, would automatically imply, according to the real yield curve explanation, a corresponding and analogous decrease in the MPCS, as the lower persistence of inflation shocks, by moving the nominal yield curve mainly at the short end, should end up blurring the predictive content intrinsic to the real yield curve. As the results in Figures 4, 5, and 7 show, this is clearly not the case. Focussing on the 1- and 4-quarter ahead horizons,

[28] For each country, and for each sample period, we compute the time-varying spectral density matrix of the VAR based on a 100-points grid for ω over the interval $[0; \pi]$. This implies that the estimates of the normalised spectrum at zero presented in the seven panels of Figure 12 are *exactly* comparable with one another.

[29] O'Reilly and Whelan (2005) estimate Euro area inflation to be essentially a unit root process.

in all three countries the MPCS computed taking the univariate model as the benchmark—the one which is relevant in the present case—exhibits a broadly hump-shaped pattern over the post-WWII era, with peaks around the time of the Volcker recession, of the beginning of the inflation targeting regime, and of the Great Inflation, respectively. The case of Canada is especially striking, with inflation persistence markedly decreasing during the first half of the 1980s, and the MPCS remaining at comparatively high levels from mid-1970s through mid-1990s.

- As for the Eurozone, the increase in inflation persistence documented in Figure 12 stands in marked contrast with the broad invariance of the MPCS at the various horizons shown in Figures 5 and 7.

- By the same token, the weakly hump-shaped pattern for inflation persistence in Australia would imply a corresponding hump-shaped pattern in the MPCS, while Figures 5 and 7 clearly point towards a gently decreasing MPCS at both the 1-and the 4-quarter horizons.

- Finally, as for the Gold Standard era, the unchanging white noiseness of inflation in the United States over the sample period is at odds with the time-varying pattern of the MPCS at the one-quarter horizon documented in Figures 5 and 7. Results for the United Kingdom, on the other hand—with inflation nearly white noise, and virtually no MPCS over and above that already encoded in past output growth—are broadly compatible with the real yield curve explanation.

6.2 Implications for the monetary policy-based explanation

Evaluating the monetary policy-based explanation of the MPCS for output growth is, on the other hand, less straightforward. Taken at face value, this explanation implies that the MPCS *uniquely* originates from monetary policy actions, so that, once appropriately controlling for the monetary policy stance, the spread should not exhibit any additional predictive content for output growth. The problem with assessing such an explanation based on the evidence produced herein is that it is not entirely clear that the short rate encodes *all* the information which is necessary for capturing the monetary policy stance.[30] Ideally, we would have liked to include, as an additional variable in our second benchmark model, the rate of growth of at least a (broad) monetary aggregate. Unfortunately, this would have implied that the augmented model including the spread would have featured *five* variables, thus running into the previously discussed 'curse of dimensionality'.[31] *Obtorto collo*, we

[30] In line with the monetarist tradition, and very much at odds with most of contemporary macroeconomics, Nelson (2003) argues, for example, for an additional, independent informational content of monetary aggregates, over and above that already encoded in the short rate.

[31] See footnote 9.

have therefore solely included, in our second benchmark, a short rate, and our results should therefore regarded as broadly tentative.

With this *caveat* in mind, the results in figures 8, 9, and 11 seem to clearly run against the notion that the MPCS may just be a reflex of the stance of monetary policy. For example,

- results for the United States during both the interwar and the post-WWII periods clearly point towards several periods during which the spread exhibited predictive power for output growth over and above that already encoded in the short rate. In particular, during both the Volcker recession, and that in 2000-2001, the spread appears to have possessed additional information compared to that contained in the simplest possible measure of the monetary policy stance.

- The case of the United Kingdom, with the MPCS appearing at both the 1- and the 4-quarter horizon after the end of the 1980s and rising again slightly in the early 2000.

That said, we are impressed by the coincidence between marked increases in MPCS and periods when the current (and future) monetary policy regime were most uncertain. If the monetary policy regime is known, and believed to be constant, agents should probably be able to deduce much of the path of future short rates from the trio of current, and past, short rates, output and inflation, included in our VAR (and so the MPCS should be zero). But when the monetary regime becomes unpredictable, at the least, some risk premia will have to be added to expectations of future short rates. Consider the number of cases to which this relationship can be applied:

- *United States*: 1890-1898, debate over Bimetallism; 1910-1913, debate over adoption of Federal Reserve System; 1979-1986, Volcker experiment with non-borrowed reserves (see Figure 4 and 10).

- *United Kingdom*: 1990-1992, ERM debate and introduction of inflation targets (see Figure 9).

- *Canada*: 1979-1981, connection with the U.S.; 1990-1992, introduction of inflation targets, and conflict between John Snow and Liberal Opposition (see Figure 9).

Note also that in all the post-WWII cases, the long rate during these years was much higher than consistent with perfect foresight, i.e. short rates dropped in subsequent years, because the inflation targets were maintained and achieved. So, there is a *prima facie* case that long rates were kept much *higher* (than consistent with the new monetary regime) because of uncertainty about its success. On this view, part of the cause of the subsequent output decline was not that the yield curve was negatively sloping, but that it was *insufficiently* so, i.e. long rates were above those consistent with the successful adoption of the new regime, thereby imparting additional downwards pressure on output growth.

While we do find this hypothesis attractive, there does, alas, appear to be a counter-example in our results. In the case of almost all the countries examined, but especially so in the U.K. and Canada, there appears to have been some rise in MPCS in the early 2000s, at a time when the monetary policy regime had been successfully established. Remember, however, that this was a period of low short-term interest rates, upwards sloping yield curves and continuing growth. If our above analysis is correct, the implication is that some external force(s) were holding down long yields (beyond those factors already in our VAR), and thereby imparting stronger growth. International influences, notably affecting the balance of world savings and investment, come to mind. But an attempt to establish this must remain for further research, so our own proposed explanations of the time-varying characteristics of the MPCS must be considered, for the time being, conjectural. Note, however, that this final conjecture, (about the reasons for the MPCS in the 2000s), relates to Harvey's (1988) view, whereas the explanation for the earlier periods based on monetary uncertainty is more in accord with the standard monetary policy approach; so both prior arguments may have *some*, but not full, validity.

7 Conclusions

In this paper we have used Bayesian time-varying parameters VARs with stochastic volatility to investigate changes in the marginal predictive content of the yield spread for output growth in the United States and the United Kingdom, since the Gold Standard era, and in the Eurozone, Canada, and Australia over the post-WWII period. Overall, our evidence has not provided full support for any of the two dominant explanations for why the yield spread may contain predictive power for output growth, the monetary policy-based one, and Harvey's (1988) 'real yield curve' one.

A The Data

Here follows a detailed description of the dataset.

A.1 Gold Standard

United States Quarterly series for the commercial paper rate, the yield on corporate bonds, the GNP deflator, and real GNP, all available from 1875:1 to 1983:4, are from Balke and Gordon (1986).

United Kingdom A series for real GNP for the period 1830-1913 is from National Accounts' Table 6 of Mitchell (1988), while a corresponding series for the GNP deflator has been computed as the ratio between the nominal GNP series from Mitchell's National Accounts' Table 5 and the just-mentioned real GNP series. A series for the short rate has been computed by linking Gurney's series for the interest rate on first-class three-months bills (available for the period 1824-1856) and the series for three-months banks bills (available for the period 1845-1938), both found in Financial Institutions' Table 15 of Mitchell (1988). Over the overlapping period, 1845-1856, the two series are near-identical, which justifies their linking. As for the long rate, we take the series for the rate on consols from Financial Institutions' Table 13 of Mitchell (1988), available for the period 1756-1980.

A.2 Interwar Period

United States Quarterly series for the commercial paper rate, the yield on comporate bonds, the GNP deflator, and real GNP, all available from 1875:1 to 1983:4, are from Balke and Gordon (1986).

United Kingdom Our output measure is the *Economist*'s seasonally adjusted index of business activity from Table 3.1 of Capie and Collins (1983), which is available for the period January 1920-December 1938. The seasonally unadjusted series for the retail price index available for the period July 1914-December 1982 is from Table III.(11) of Capie and Webber (1985). As for the short rate, we take the market rate of interest on best three-month bills (quoted at an annual rate), available for the period January 1919-December 1939, from Table 7.1 of Capie and Collins (1983). As for the long rate, we take the series for the yield on 2.5% consols (quoted at an annual rate) from Table 7.5 of Capie and Collins (1983), available for the period January 1919-December 1939. Given that the RPI series is seasonally unadjusted, prior to analysis we seasonally adjust it via the ARIMA X-12 procedure as implemented in *Eviews*.

A.3 Post-WWII

United States Quarterly seasonally adjusted series for real GDP (GDPC96) and the GDP deflator (GDPDEF) are from the *U.S. Department of Commerce, Bureau of*

Economic Analysis. A quarterly seasonally adjusted CPI series has been obtained by keeping the last observation from each quarter from the original monthly series from the *U.S. Department of Labor, Bureau of Labor Statistics* (CPIAUCSL: 'Consumer price index for all urban consumers, all items'). A quarterly seasonally adjusted import price index (11176.X.ZF...), a Treasury Bill Rate (11160C..ZF...), and a 10-year government bond yield (11161...ZF...) are from the *International Monetary Fund*'s International Financial Statistics.

Eurozone Quarterly seasonally adjusted series for real GDP (YER), the GDP deflator (YED), short-term and long-term interest rates (STN and LTN) are from the ECB's Area Wide Model's dataset. The overall sample period is 1970:1-2003:4.

United Kingdom Quarterly seasonally adjusted series for real GDP (ABMI) and the GDP deflator (YBGB) are from the *Office for National Statistics.* A quarterly seasonally adjusted import price index (11276.X.ZF...), a Treasury Bill Rate (11260C..ZF...), and a 10-year government bond yield (11261...ZF...) are from the *International Monetary Fund*'s International Financial Statistics. A quarterly seasonally adjusted RPI series has been obtained by keeping the last observation from each quarter from the original seasonally unadjusted monthly series from the *Office for National Statistics.*(CDKO), and then seasonally adjusting the resulting series via the ARIMA X-12 procedure as implemented in *Eviews.*

Canada Quarterly seasonally adjusted series for real GDP (15699BVRZF), the GDP deflator (15699BIRZF), short and long interest rates (15660C..ZF and 15661...ZF) are from the *International Monetary Fund*'s International Financial Statistics. The sample period is 1975:1-2005:2.

Australia Quarterly seasonally adjusted series for real GDP (19399BVRZF), the GDP deflator (19399BIRZF), and short and long interest rates (19360B..ZF and 19361...ZF) are from the *International Monetary Fund*'s International Financial Statistics. The overall sample period is 1957:1-2005:1.

B Details of the Markov-Chain Monte Carlo Procedure

We estimate (5)-(11) *via* Bayesian methods. The next two subsections describe our choices for the priors, and the Markov-Chain Monte Carlo algorithm we use to simulate the posterior distribution of the hyperparameters and the states conditional on the data, while the third section discusses how we check for convergence of the Markov chain to the ergodic distribution.

B.1 Priors

For the sake of simplicity, the prior distributions for the initial values of the states—θ_0 and h_0—which we postulate all to be normal, are assumed to be independent both from each other, and from the distribution of the hyperparameters. In order to calibrate the prior distributions for θ_0 and h_0 we estimate a time-invariant version of (5) based on the first 8 years of data, and we set

$$\theta_0 \sim N\left[\hat{\theta}_{OLS}, 4 \cdot \hat{V}(\hat{\theta}_{OLS})\right] \tag{B1}$$

where $\hat{V}(\hat{\theta}_{OLS})$ is the estimated asymptotic variance of $\hat{\theta}_{OLS}$. As for h_0, we proceed as follows. Let $\hat{\Sigma}_{OLS}$ be the estimated covariance matrix of ϵ_t from the time-invariant VAR, and let C be its lower-triangular Cholesky factor—i.e., $CC' = \hat{\Sigma}_{OLS}$. We set

$$\ln h_0 \sim N(\ln \mu_0, 10 \times I_N) \tag{B2}$$

where μ_0 is a vector collecting the logarithms of the squared elements on the diagonal of C. As stressed by Cogley and Sargent (2005), 'a variance of 10 is huge on a natural-log scale, making this weakly informative' for h_0.

Turning to the hyperparameters, we postulate independence between the parameters corresponding to the two matrices Q and A—an assumption we adopt uniquely for reasons of convenience—and we make the following, standard assumptions. The matrix Q is postulated to follow an inverted Wishart distribution,

$$Q \sim IW\left(\bar{Q}^{-1}, T_0\right) \tag{B3}$$

with prior degrees of freedom T_0 and scale matrix $T_0\bar{Q}$. In order to minimize the impact of the prior, thus maximizing the influence of sample information, we set T_0 equal to the minimum value allowed, the length of θ_t plus one. As for $\bar{Q}$, we calibrate it as $\bar{Q}= \gamma \times \hat{\Sigma}_{OLS}$, setting γ=3.5$\times$10^{-4}, the same value used in Cogley and Sargent (2005). As for α, we postulate it to be normally distributed woth a 'large' variance,

$$f(\alpha) = N(0,\ 10000 \cdot I_{N(N-1)/2}). \tag{B4}$$

Finally, as for the variances of the stochastic volatility innovations, we follow Cogley and Sargent (2002, 2005) and we postulate an inverse-Gamma distribution for $\sigma_i^2 \equiv \mathrm{Var}(\nu_{i,t})$:

$$\sigma_i^2 \sim IG\left(\frac{10^{-4}}{2}, \frac{1}{2}\right) \tag{B5}$$

B.2 Simulating the posterior distribution

We simulate the posterior distribution of the hyperparameters and the states conditional on the data *via* the following MCMC algorithm, as found in Cogley and Sargent (2005). In what follows, x^t denotes the entire history of the vector x up to time t—i.e. $x^t \equiv [x_1', x_2', \quad, x_t']'$—while T is the sample length.

(a) Drawing the elements of θ_t Conditional on Y^T, α, and H^T, the observation equation (5) is linear, with Gaussian innovations and a known covariance matrix. Following Carter and Kohn (2004), the density $p(\theta^T|Y^T,\alpha,H^T)$ can be factored as

$$p(\theta^T|Y^T,\alpha,H^T)=p(\theta_T|Y^T,\alpha,H^T)\prod_{t=1}^{T-1}p(\theta_t|\theta_{t+1},Y^T,\alpha,H^T) \quad \text{(B6)}$$

Conditional on α and H^T, the standard Kalman filter recursions nail down the first element on the right hand side of (A6), $p(\theta_T|Y^T,\alpha,H^T)=N(\theta_T,P_T)$, with P_T being the precision matrix of θ_T produced by the Kalman filter. The remaining elements in the factorization can then be computed via the backward recursion algorithm found, e.g., in Kim and Nelson (2000), or Cogley and Sargent (2005, appendix B.2.1). Given the conditional normality of θ_t, we have

$$\theta_{t|t+1}=\theta_{t|t}+P_{t|t}P^{-1}_{t+1|t}\left(\theta_{t+1}-\theta_t\right) \quad \text{(B7)}$$

$$P_{t|t+1}=P_{t|t}-P_{t|t}P^{-1}_{t+1|t}P_{t|t} \quad \text{(B8)}$$

which provides, for each t from T-1 to 1, the remaining elements in (5), $p(\theta_t|\theta_{t+1}$, Y^T, α, $H^T)$ = $N(\theta_{t|t+1}$, $P_{t|t+1})$. Specifically, the backward recursion starts with a draw from $N(\theta_T,P_T)$, call it $\tilde{\theta}_T$ Conditional on $\tilde{\theta}_T$, (A7)-(A8) give us $\theta_{T-1|T}$ and $P_{T-1|T}$, thus allowing us to draw $\tilde{\theta}_{T-1}$ from $N(\theta_{T-1|T},P_{T-1|T})$, and so on until t=1.

(b) Drawing the elements of H_t Conditional on Y^T, θ^T, and α, the orthogonalised innovations $u_t\equiv A(Y_t$-$X_t^{\prime}\theta_t)$, with Var(u_t)=H_t, are observable. Following Cogley and Sargent (2002), we then sample the $h_{i,t}$'s by applying the univariate algorithm of Jacquier, Polson, and Rossi (2004) element by element.[32]

(c) Drawing the hyperparameters Conditional on Y^T, θ^T, H^T, and α, the innovations to θ_t and to the $h_{i,t}$'s are observable, which allows us to draw the hyperparameters—the elements of Q and the σ_i^2—from their respective distributions.

(d) Drawing the elements of α Finally, conditional on Y^T and θ^T the ϵ_t's are observable, satisfying

$$A\epsilon_t=u_t \quad \text{(B9)}$$

with the u_t being a vector of orthogonalized residuals with known time-variiying variance H_t. Following Cogley and Sargent (2005), we interpret (B9) as a system of unrelated regressions. The first equation in the system is given by $\epsilon_{1,t}\equiv u_{1,t}$, while the following equations can be expressed as transformed regressions as

$$\left(h_{2,t}^{-\frac{1}{2}}\epsilon_{2,t}\right)=-\alpha_{2,1}\left(h_{2,t}^{-\frac{1}{2}}\epsilon_{1,t}\right)+\left(h_{2,t}^{-\frac{1}{2}}u_{2,t}\right) \quad \text{(B10)}$$

$$\left(h_{3,t}^{-\frac{1}{2}}\epsilon_{3,t}\right)=-\alpha_{3,1}\left(h_{3,t}^{-\frac{1}{2}}\epsilon_{1,t}\right)-\alpha_{3,2}\left(h_{3,t}^{-\frac{1}{2}}\epsilon_{2,t}\right)+\left(h_{3,t}^{-\frac{1}{2}}u_{3,t}\right)$$

[32] For details, see Cogley and Sargent (2005, Appendix B.2.5).

$$\dots$$

$$\left(h^{-\frac{1}{2}}_{N(N-1)/2,t}\epsilon_{N(N-1)/2,t}\right) = -\alpha_{N(N-1)/2,1}\left(h^{-\frac{1}{2}}_{N(N-1)/2,t}\epsilon_{1,t}\right) - \dots$$

$$\dots - \alpha_{N(N-1)/2,N(N-1)/2}\left(h^{-\frac{1}{2}}_{N(N-1)/2,t}\epsilon_{N(N-1)/2,t}\right) + \left(h^{-\frac{1}{2}}_{N(N-1)/2,t}u_{N(N-1)/2,t}\right)$$

where the residuals are independent standard normal. Assuming normal priors for each equation's regression coefficients the posterior is also normal, and can be computed via equations (77) of (78) in Cogley and Sargent (2005, section B.2.4).

Summing up, the MCMC algorithm simulates the posterior distribution of the states and the hyperparameters, conditional on the data, by iterating on (a)-(d). In what follows we use a burn-in period of 10,000 iterations to converge to the ergodic distribution, and after that—following Cogley and Sargent (2005)—we sample every 10th draw of the subsequent 20,000 iterations in order to reduce the autocorrelation across draws, thus getting a sample of 2,000 draws from the ergodic distribution.

B.3 Assessing the convergence of the Markov chain to the ergodic distribution

Following Primiceri (2005), we assess the convergence of the Markov chain by inspecting the autocorrelation properties of the ergodic distribution's 2,000 draws. Specifically, we consider the draws' inefficiency factors (henceforth, IFs), defined as the inverse of the relative numerical efficiency measure of Geweke (1992),

$$RNE = (2\pi)^{-1}\frac{1}{S(0)}\int_{-\pi}^{\pi} S(\omega)d\omega \tag{20}$$

where $S(\omega)$ is the spectral density of the sequence of draws from the Gibbs sampler for the quantity of interest at the frequency ω. In what follows, we estimate the spectral densities by smoothing the periodograms[33] in the frequency domain by means of a Bartlett spectral window. Following Berkowitz and Diebold (1998), we select the bandwidth parameter automatically *via* the procedure introduced by Beltrao and Bloomfield (1987).

Figures 13 and 14 show, for the United States and the United Kingdom for the post-WWII periods, the draws' inefficiency factors for the models' hyperparameters—i.e., the free elements of the matrices Q and A—and for the states, i.e. the time-varying coefficients of the VAR (the θ_t) and the volatilities (the $h_{i,t}$'s). As the figures show, the autocorrelation of the draws is extremely low, with all the estimated IFs being around or below 3—as stressed by Primiceri (2005, Appendix B), values of the IFs below or around twenty are generally regarded as satisfactory. Analogous evidence for the other three countries is not reported here for reasons of space, but is available from the authors upon request.

[33] We compute the periodograms based on the fast-Fourier transform

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Table 1 Results based on the Stock-Watson TVP-MUB methodology: *exp*- and *sup*-Wald test statistics, simulated *p*-values, and median-unbiased estimates of λ

	exp-Wald (p-value)	$\hat{\lambda}$	*sup*-Wald (p-value)	$\hat{\lambda}$
United States				
Classical Gold Standard	15.85 (0)	0.04138	40.66 (0)	0.04483
Interwar period[a]	30.58 (0)	0.08966	68.78 (0)	0.10345
Post-WWII period	25.91 (0)	0.05172	29.41 (0.005)	0.03448
Eurozone, post-WWII[b]	25.99 (0)	0.05517	60.52 (0)	0.10345
United Kingdom, post-WWII	19.02 (0)	0.08966	28.76 (0)	0.05172
Canada, post-WWII	11.37 (0)	0.04138	30.39 (0)	0.04138
Australia, post-WWII	11.40 (0)	0.05172	30.19 (0)	0.05172

[a] Trimming set to 0.25. [b] Trimming set to 0.3.

Table 2 The marginal predictive content of the spread in the United Kingdom during the Gold Standard: overall and marginal R^2's

	Benchmark model with:			
	$Y_t \equiv y_t$		$Y_t \equiv [y_t, \pi_t, r_t]'$	
Horizon (in years)	1	2	1	2
	Frequency-domain-based estimates:			
Overall predictability	0.116	8.5E-04	0.116	4.6E-04
Marginal predictability	0.002	2.4E-04	-7.7E-05	-1.2E-04
	Time-domain-based estimates:			
Overall predictability	0.148	0.005	0.154	0.027
Marginal predictability	0.038	0.005	0.001	0.008

Table 3 The marginal predictive content of the spread in the United Kingdom during the interwar period: overall and marginal R^2's

	Benchmark model with:					
	$Y_t \equiv y_t$			$Y_t \equiv [y_t, \pi_t, r_t]'$		
Horizon (in months)	3	12	24	3	12	24
	Frequency-domain-based estimates:					
Overall predictability	0.162	0.049	0.060	0.247	0.032	0.289
Marginal predictability	0.004	-0.007	0.054	0.076	-0.021	0.212
	Time-domain-based estimates:					
Overall predictability	0.159	0.057	0.109	0.209	0.077	0.143
Marginal predictability	0.023	0.015	0.104	0.032	0.024	0.045

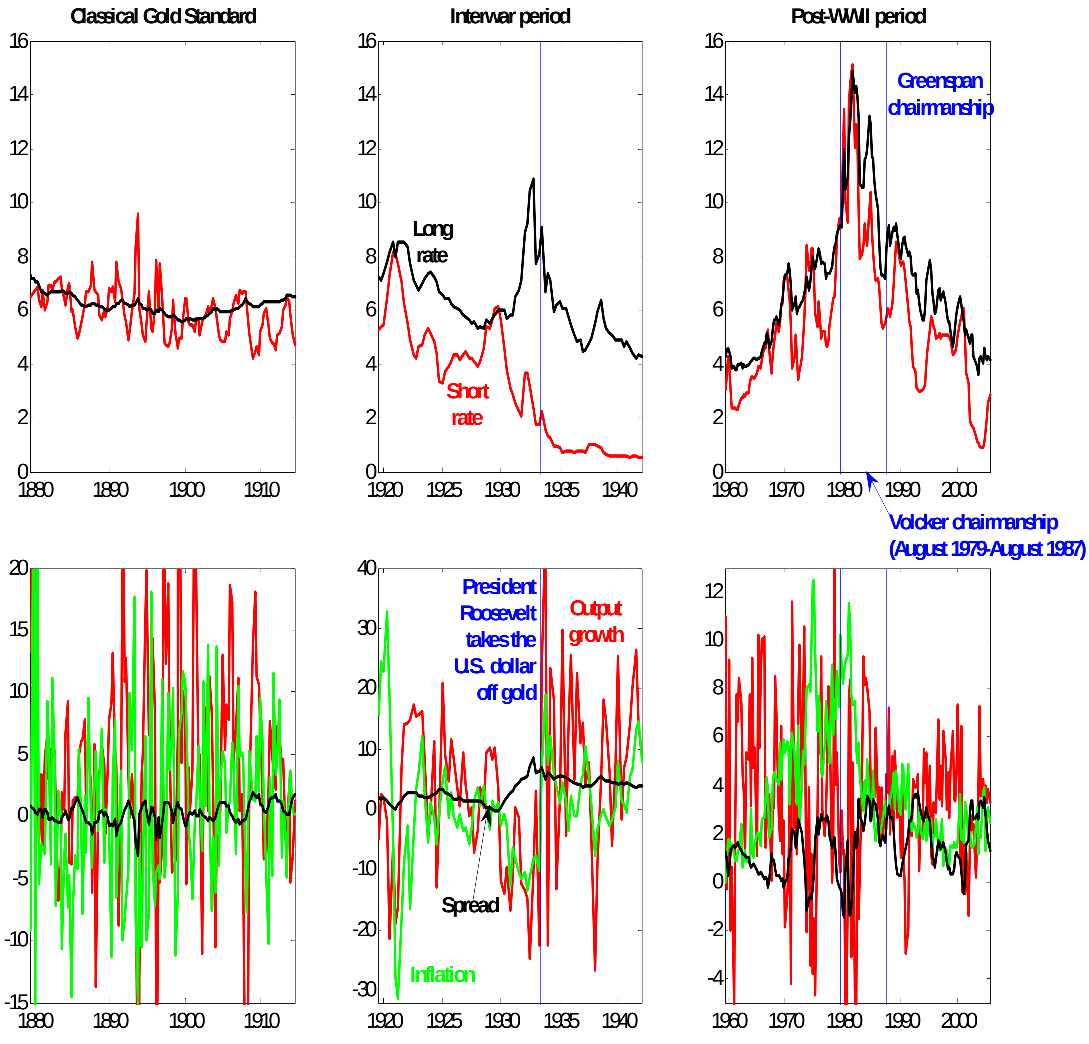

Figure 1: United States, the raw data

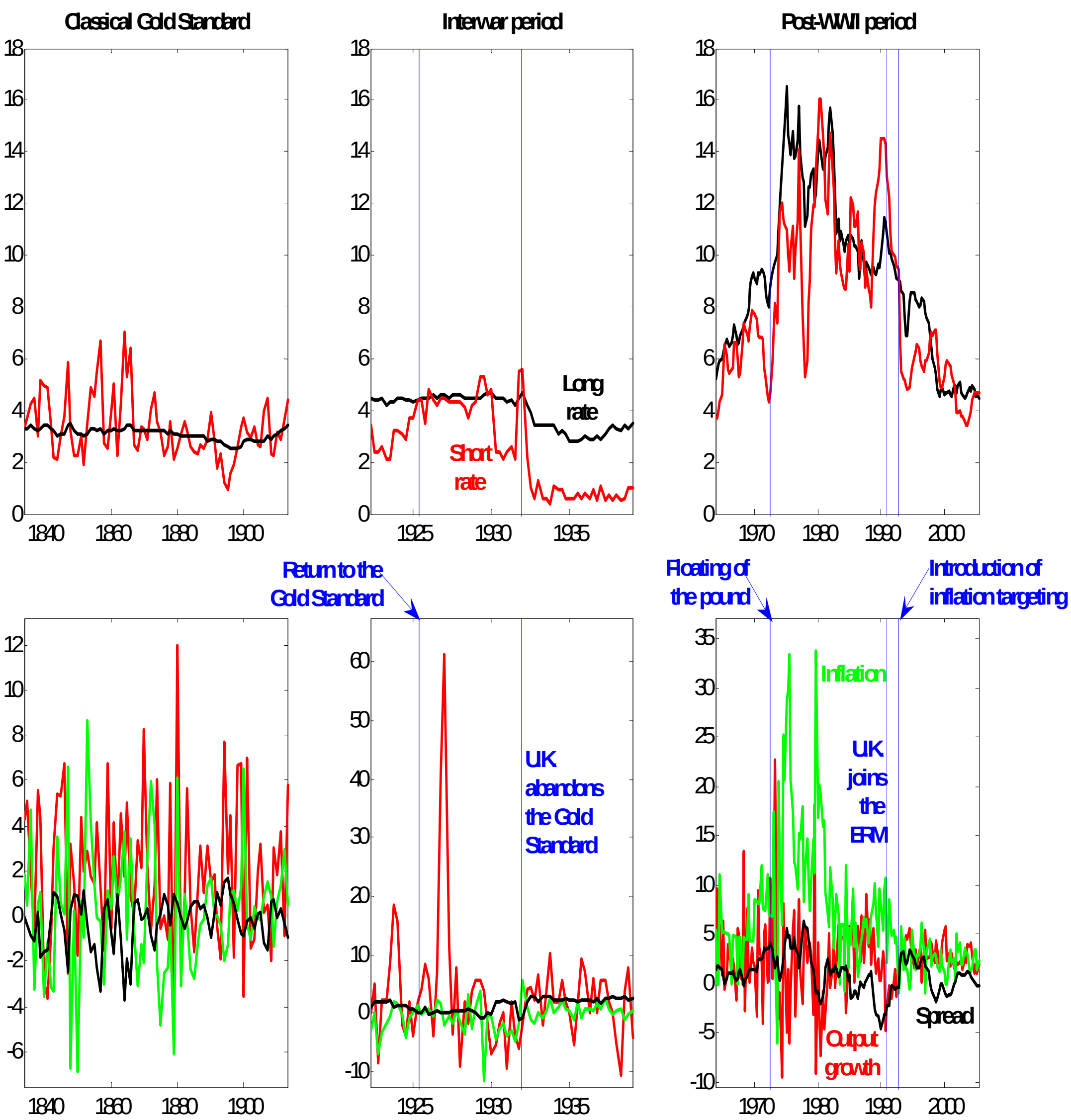

Figure 2: United Kingdom, the raw data

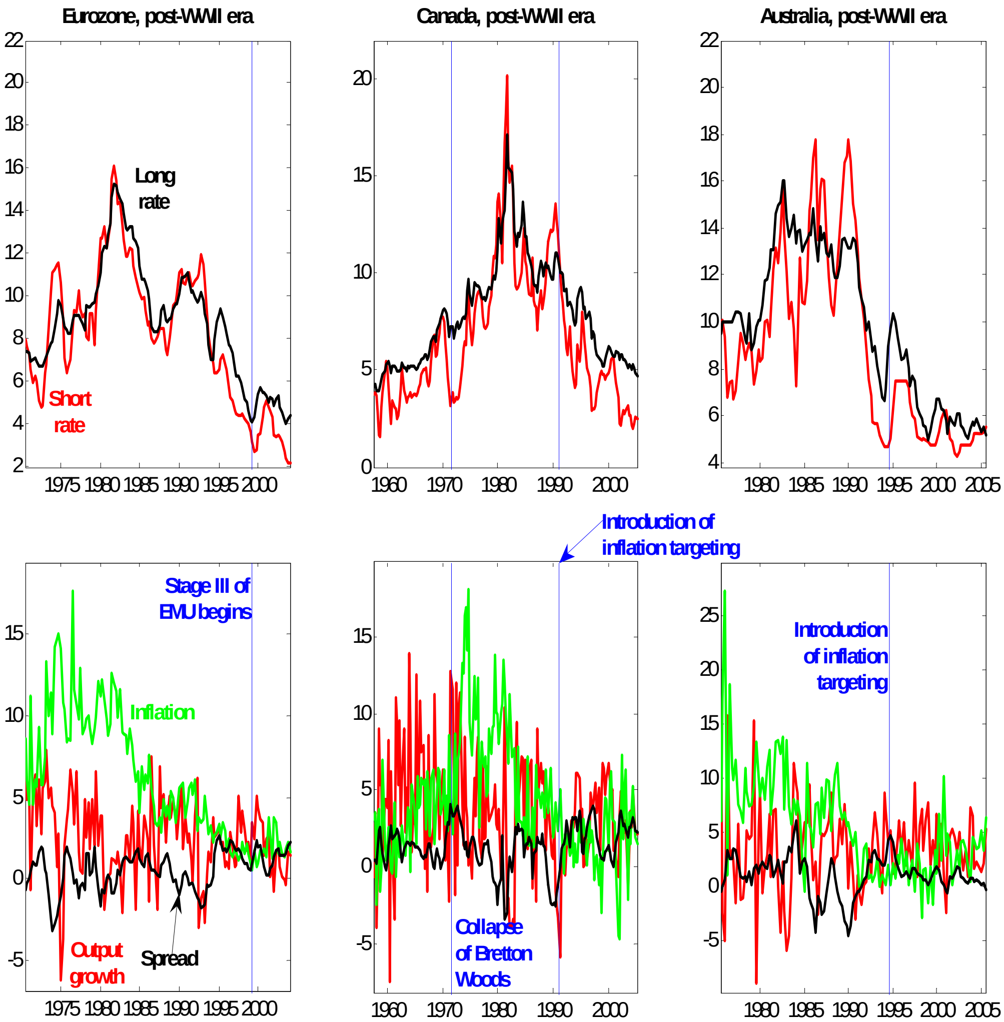

Figure 3: Eurozone, Canada, and Australia, the raw data

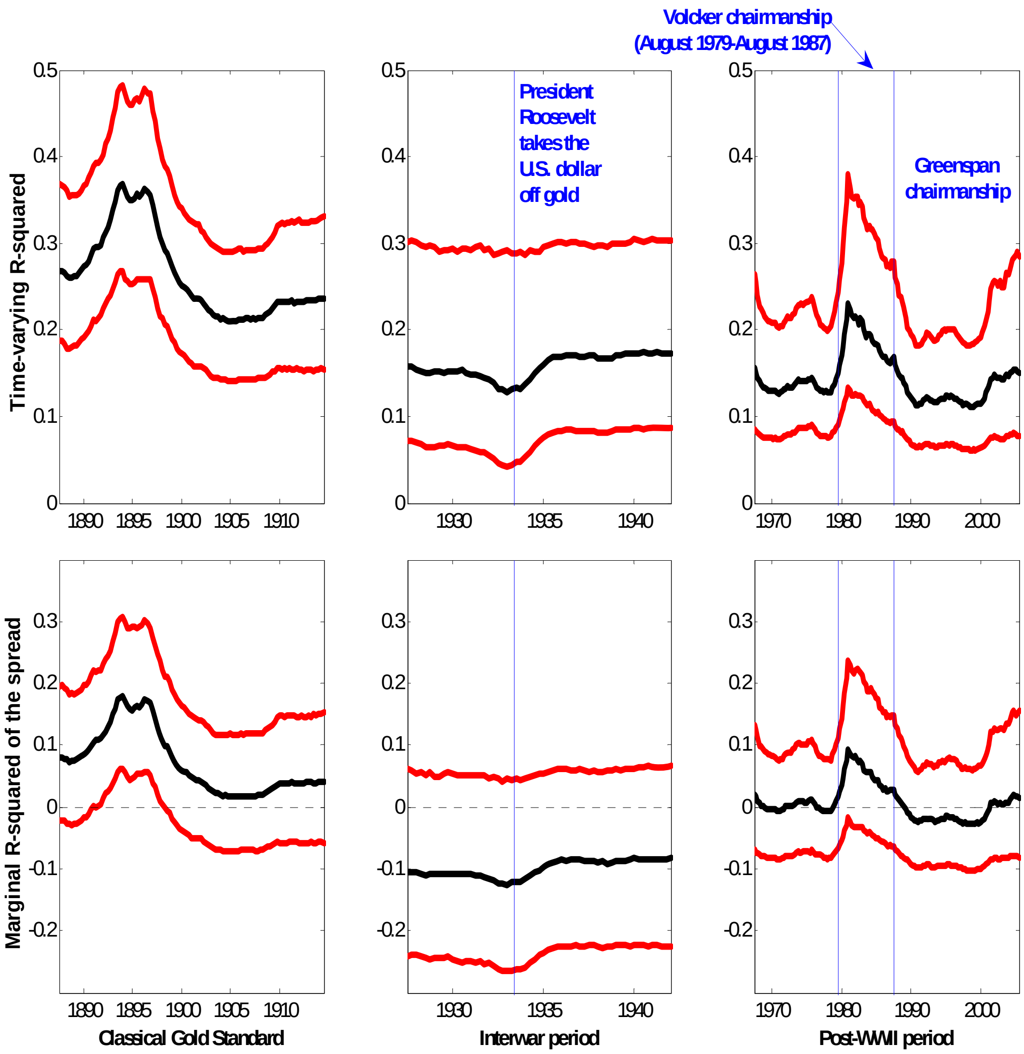

Figure 4: Measuring changes in U.S. output growth's predictability: time-varying overall R^2s and marginal predictive content of the spread, one quarter ahead (median estimates and 16th and 84th percentiles): univariate versus bivariate results

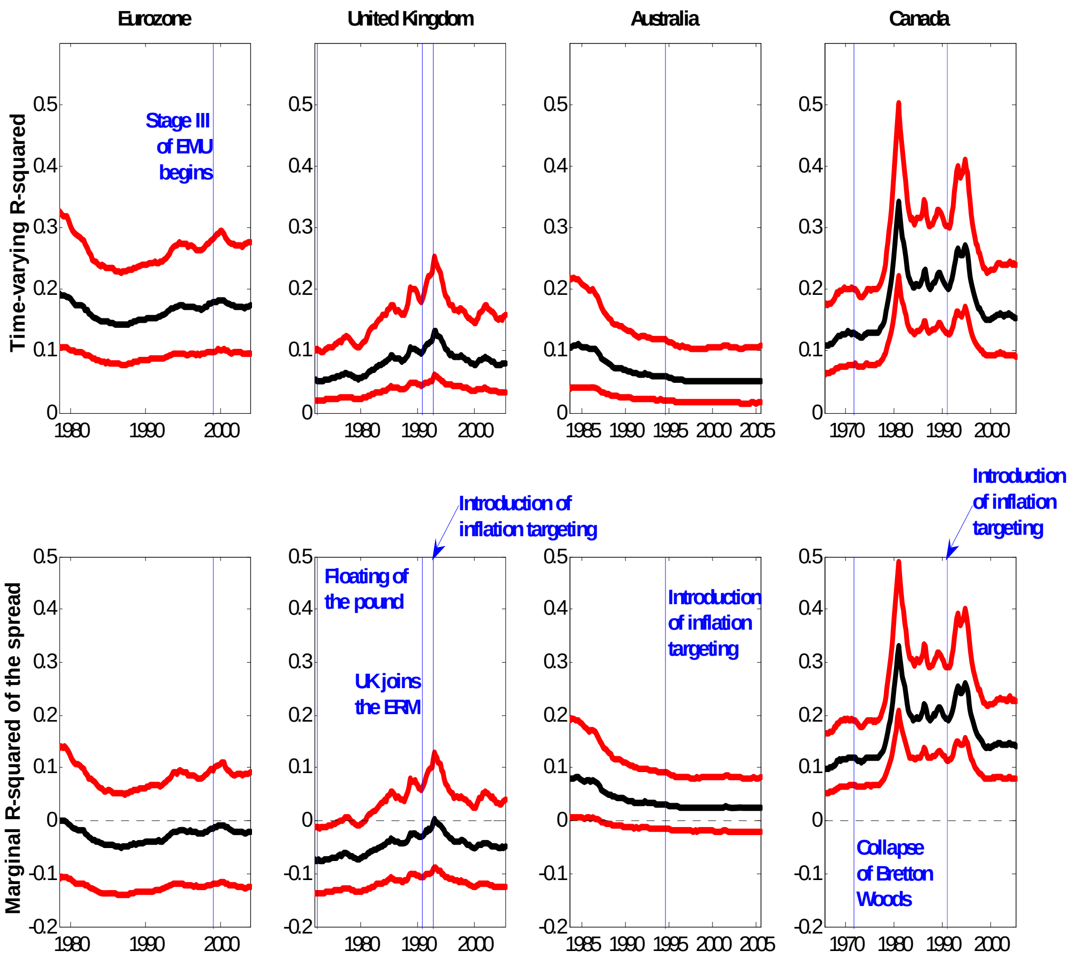

Figure 5: Measuring changes in output growth's predictability for the Eurozone, the United Kingdom, Australia, and Canada: time-varying overall R^2s, and marginal predictive content of the spread, one quarter ahead (median estimates and 16th and 84th percentiles): univariate versus bivariate results

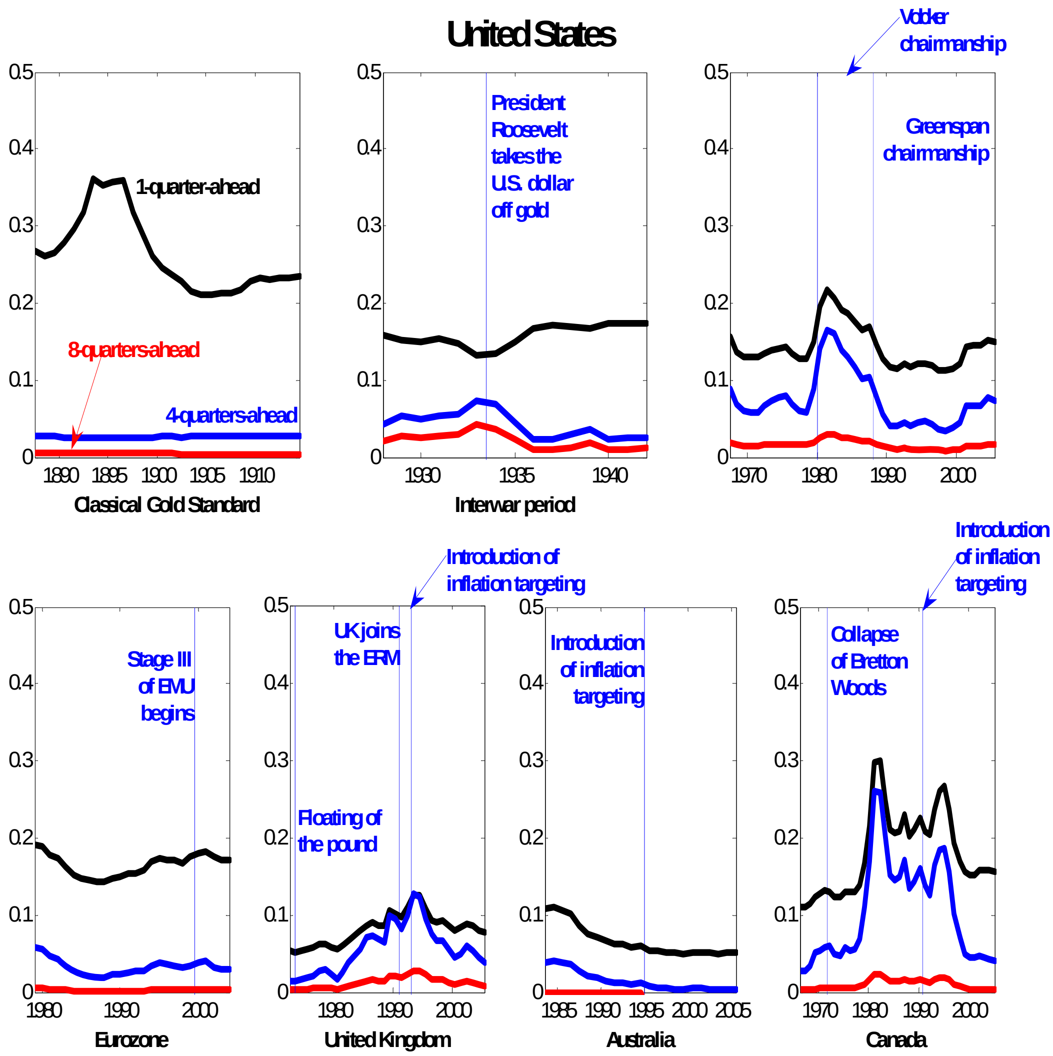

Figure 6: United States, Eurozone, United Kingdom, Australia, and Canada: overall predictability at the 1-, 4-, and 8-quarters ahead horizons (median estimates)

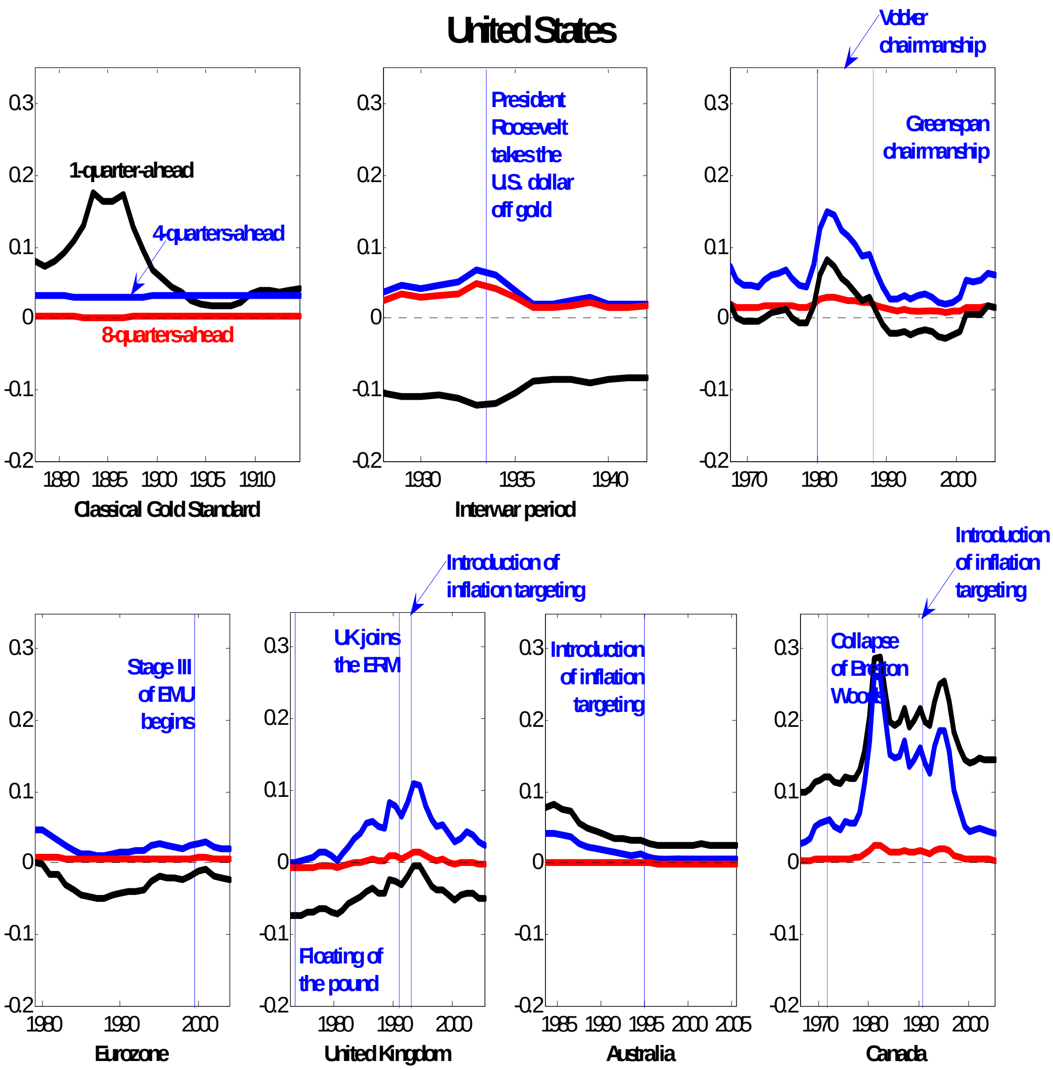

Figure 7: United States, Eurozone, United Kingdom, Australia, and Canada: marginal predictive content of the spread at the 1-, 4-, and 8-quarters ahead horizons (median estimates)

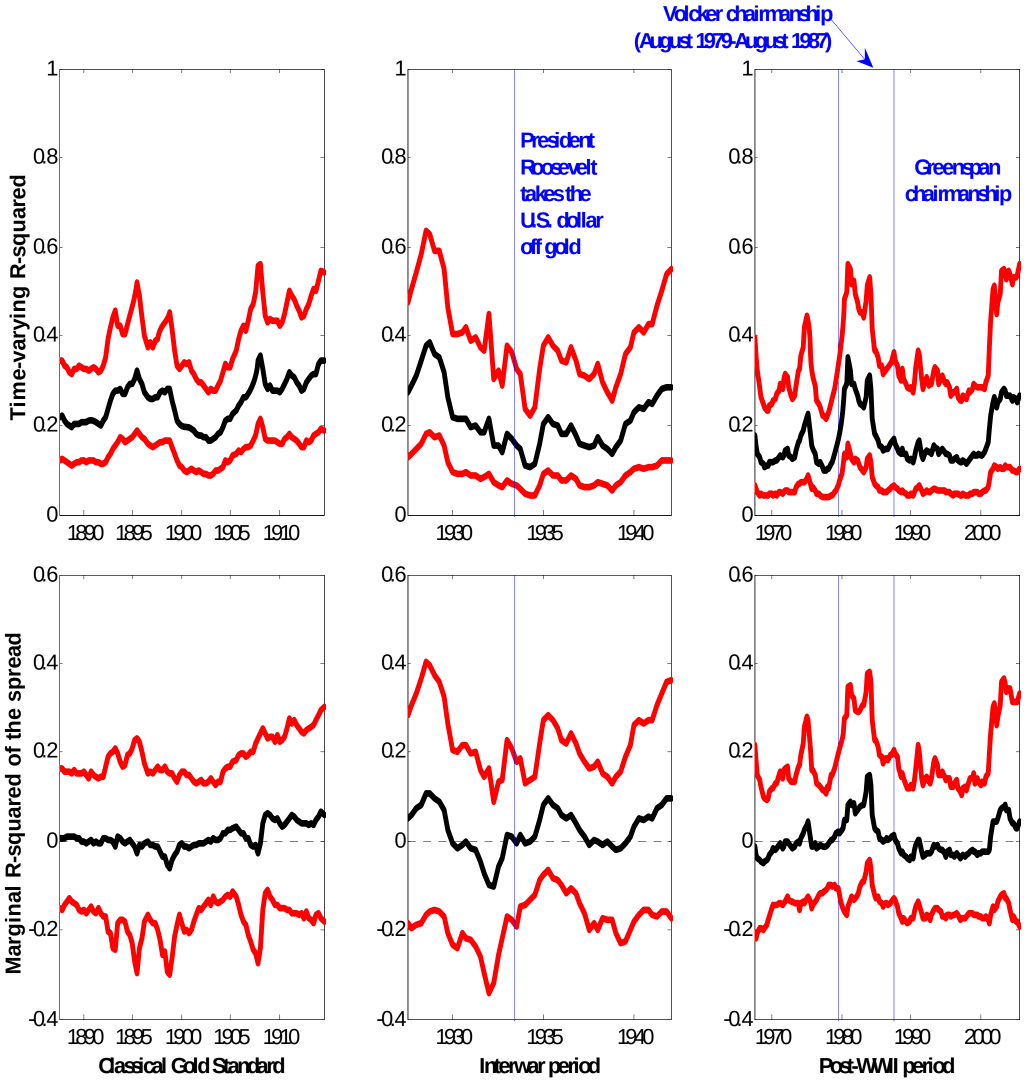

Figure 8: Measuring changes in U.S. output growth's predictability: time-varying overall R^2s, and marginal predictive content of the spread, one quarter ahead (median estimates and 16th and 84th percentiles): trivariate versus four-variate results

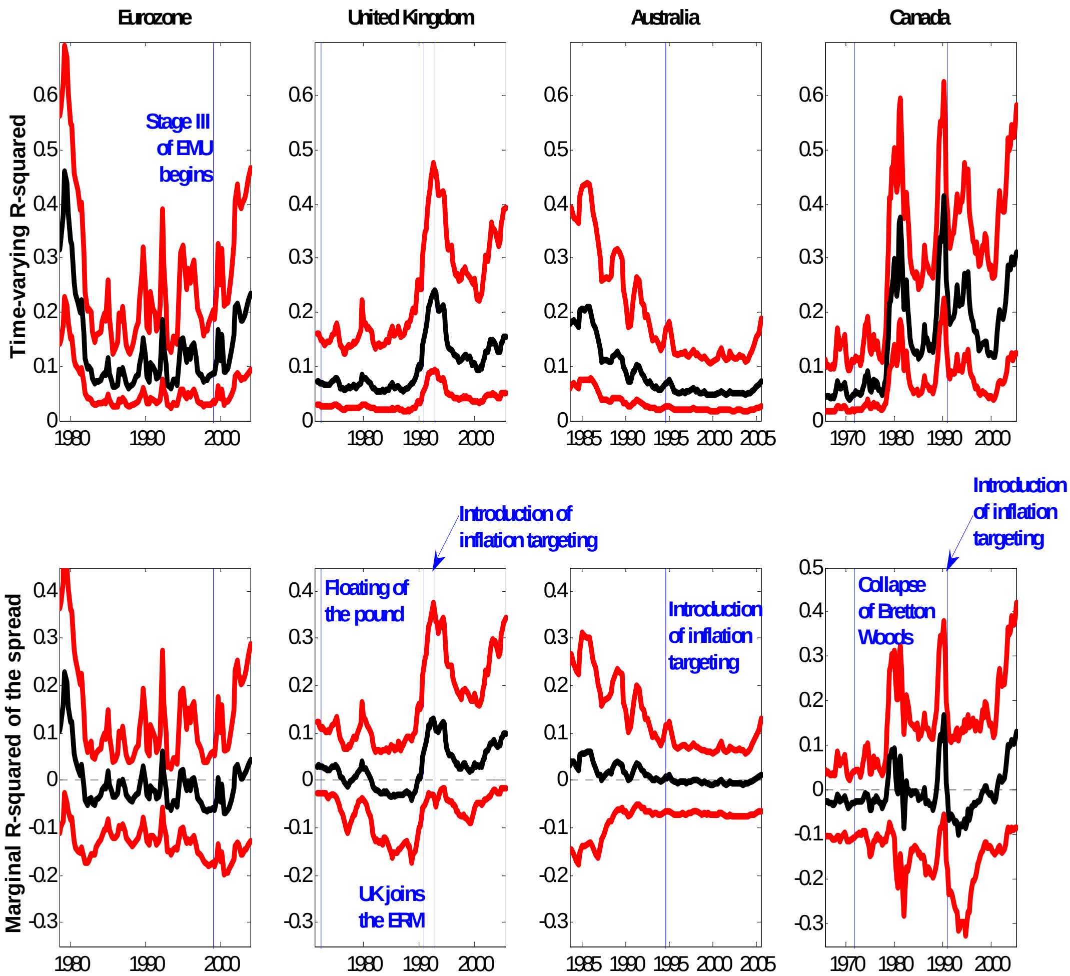

Figure 9: Measuring changes in output growth's predictability for the Eurozone, the United Kingdom, Australia, and Canada: time-varying overall R^2s, and marginal predictive content of the spread, one quarter ahead (median estimates and 16th and 84th percentiles): trivariate versus four-variate results

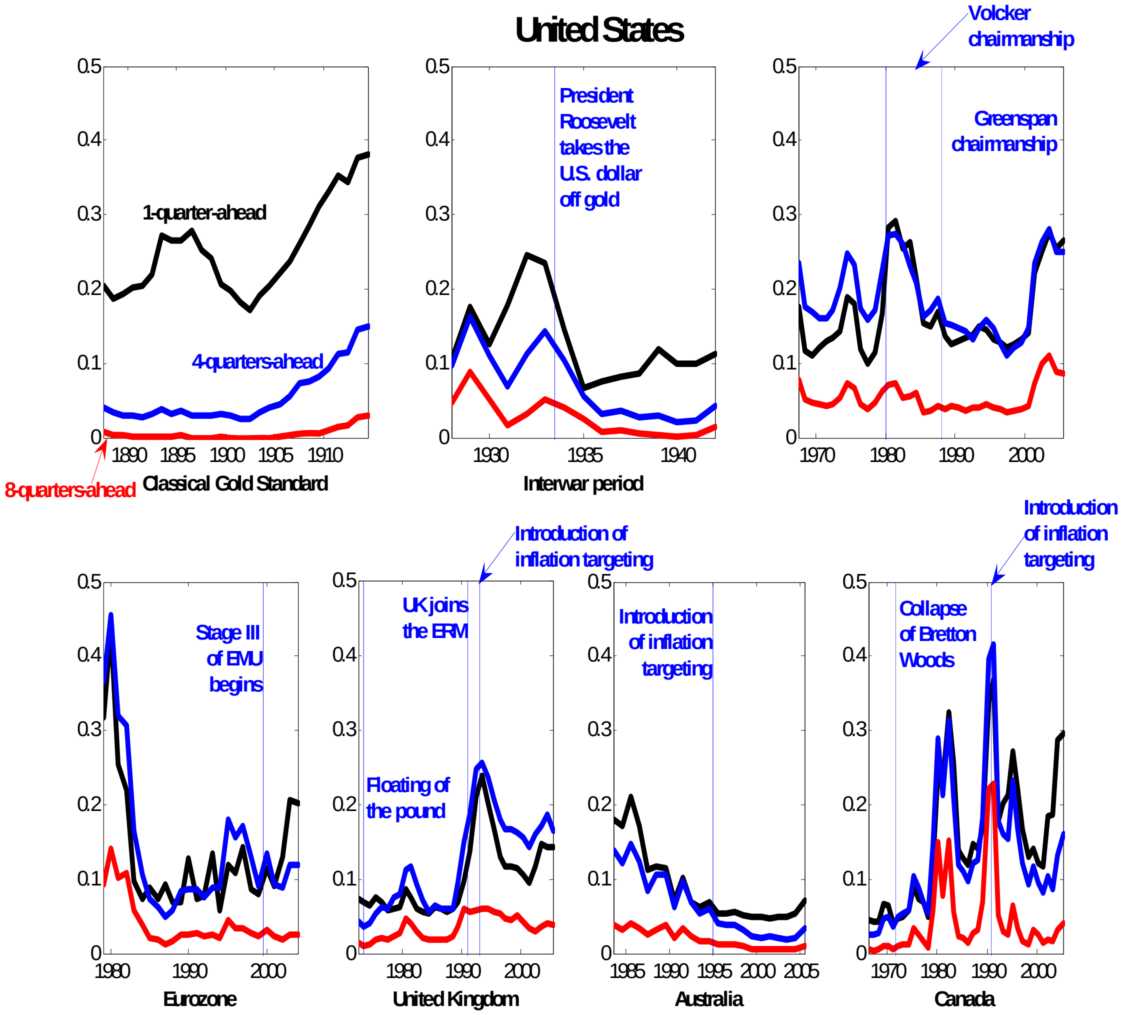

Figure 10: United States, Eurozone, United Kingdom, Australia, and Canada: overall predictability at the 1-, 4-, and 8-quarters ahead horizons (median estimates)

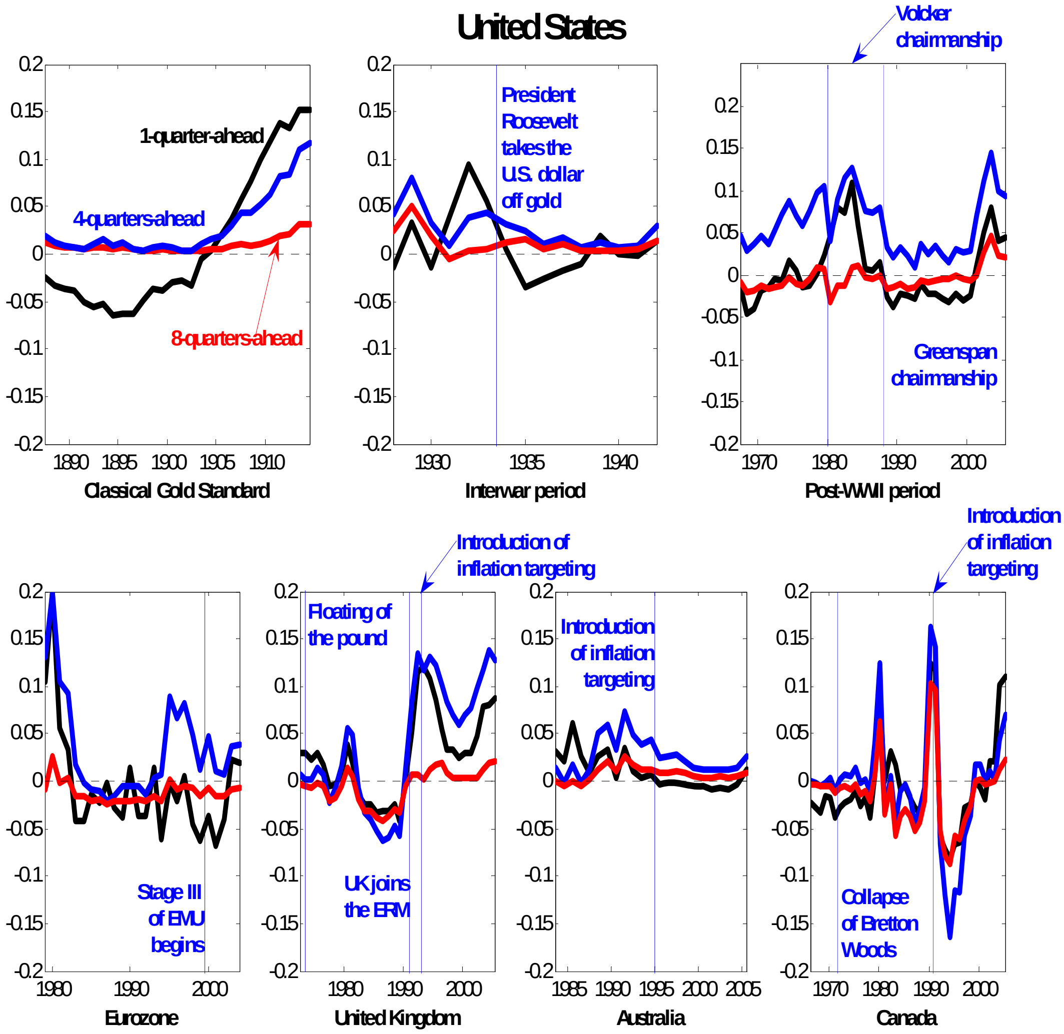

Figure 11: United States, Eurozone, United Kingdom, Australia, and Canada: marginal predictive content of the spread at the 1-, 4-, and 8-quarters ahead horizons (median estimates)

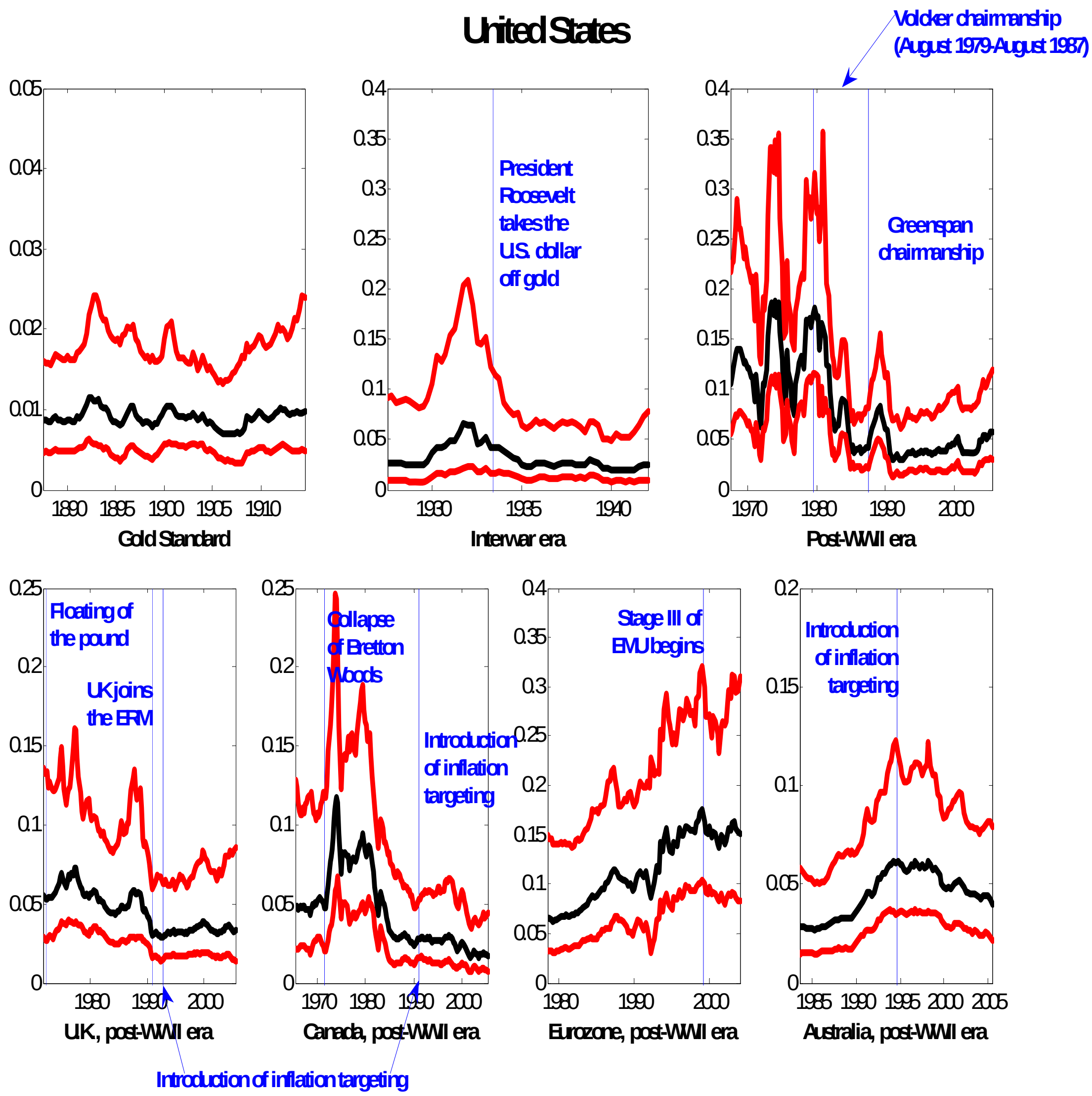

Figure 12: Measuring time-variation in inflation persistence: normalised spectrum of inflation at ω=0 (median estimates and 16th and 84th percentiles)

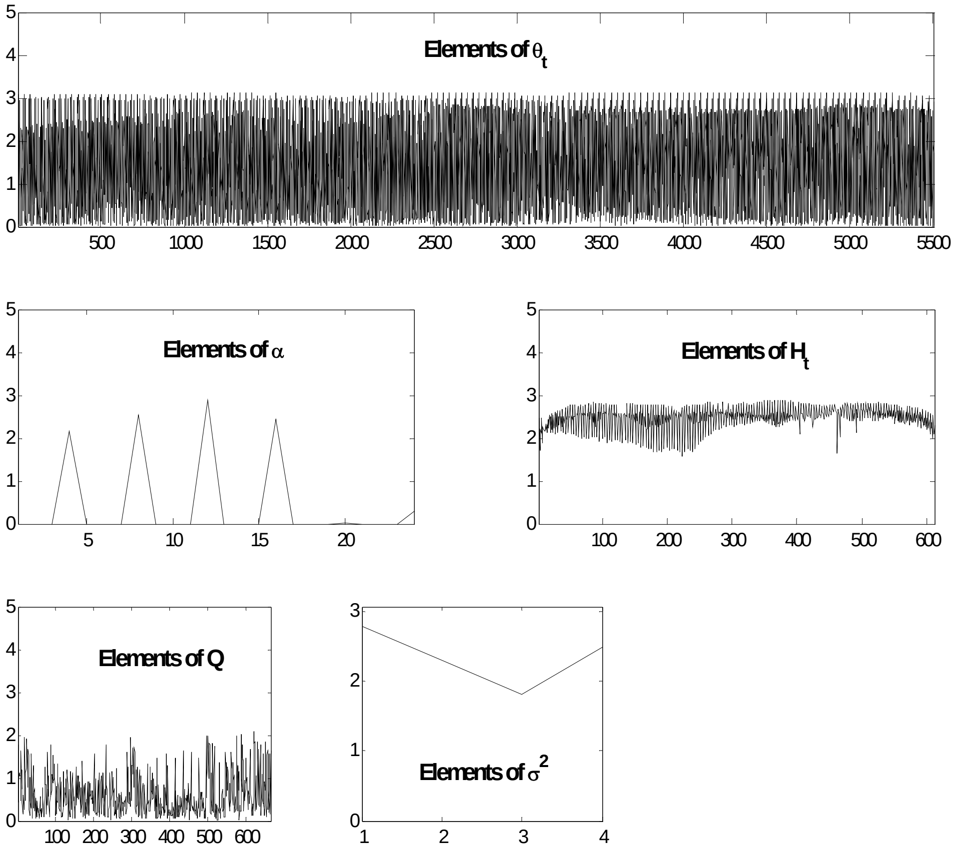

Figure 13: Checking for convergence of the Markov chain: inefficiency factors for the draws from the ergodic distribution for the hyperparameters and the states (United States, post-WWII era, VAR with the spread)

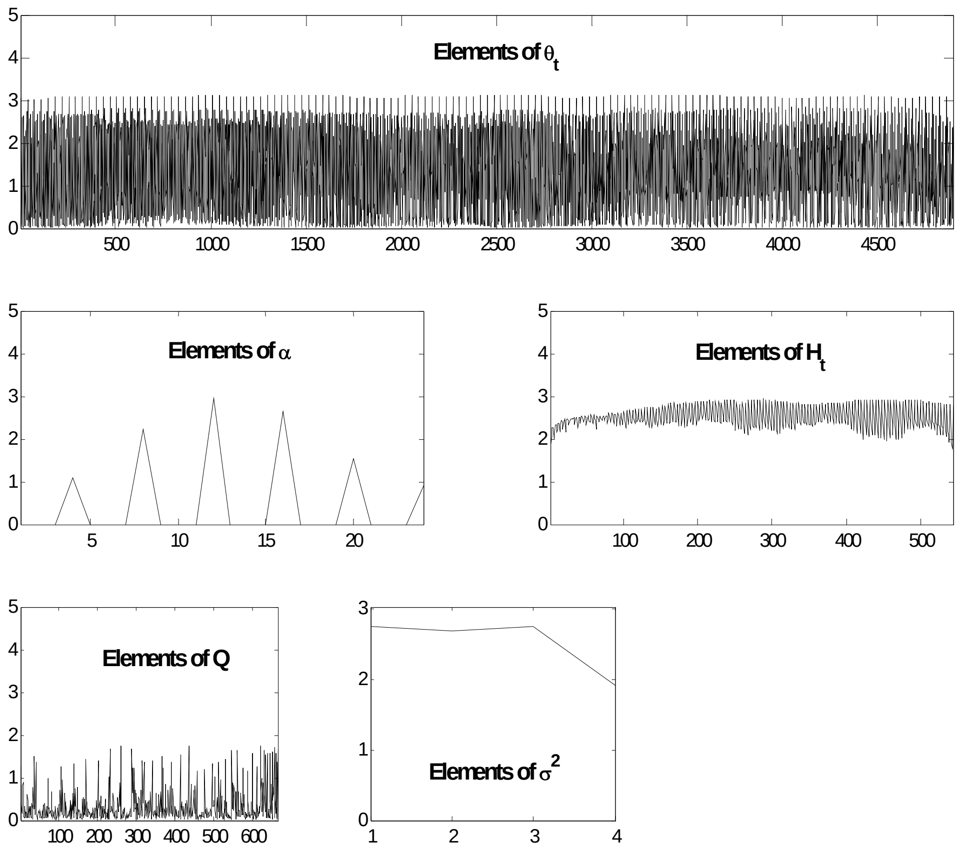

Figure 14: Checking for convergence of the Markov chain: inefficiency factors for the draws from the ergodic distribution for the hyperparameters and the states (United Kingdom, post-WWII era, VAR with the spread)

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