



**FACULTEIT ECONOMIE  
EN BEDRIJFSKUNDE**

**TWEEKERKENSTRAAT 2  
B-9000 GENT**

**Tel. : 32 - (0)9 - 264.34.61  
Fax. : 32 - (0)9 - 264.35.92**

## **WORKING PAPER**

### **Macroeconomic Regimes**

**Lieven Baele**

**Geert Bekaert**

**Seonghoon Cho**

**Koen Inghelbrecht**

**Antonio Moreno**

December 2013

2013/870

D/2013/7012/41

# Macroeconomic Regimes\*

Lieven Baele<sup>1</sup>      Geert Bekaert<sup>2</sup>      Seonghoon Cho<sup>3</sup>

Koen Inghelbrecht<sup>4</sup>      Antonio Moreno<sup>5</sup>

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## Abstract

We estimate a New-Keynesian macro model accommodating regime-switching behavior in monetary policy and in macro shocks. Key to our estimation strategy is the use of survey-based expectations for inflation and output. Output and inflation shocks shift to the low volatility regime around 1985 and 1990, respectively. However, we also identify multiple shifts between accommodating and active monetary policy regimes, which play an as important role as shock volatility in driving the volatility of the macro variables. We provide new estimates of the onset and demise of the Great Moderation and quantify the relative role played by macro-shocks and monetary policy. The estimated rational expectations model exhibits indeterminacy in the mean-square stability sense, mainly because monetary policy is excessively passive.

JEL Classification: E31, E32, E52, E58, C42, C53

Keywords: Markov-Switching (MS) DSGE models, Survey Expectations, Great Moderation, Monetary Policy, Determinacy in MS DSGE models

\* We appreciate the comments and suggestions of seminar participants at Fordham University, Columbia University, the University of Navarra, Korea University, the American Economic Society Meetings (2012, Chicago), the Society of Economic Dynamics (2011, Ghent), and the European Economic Association (2011, Oslo).

<sup>1</sup> Finance Department, CentER, and Netspar, Tilburg University. Email: Lieven.Baele@uvt.nl

<sup>2</sup> Graduate School of Business, Columbia University, and NBER. Email: gb241@columbia.edu

<sup>3</sup> School of Economics, Yonsei University. Email: sc719@yonsei.ac.kr. Fax: 82-2-393-1158. (corresponding author)

<sup>4</sup> Department Financial Economics, Ghent University, and Finance Department, University College Ghent. Email: Koen.Inghelbrecht@hogent.be

<sup>5</sup> Department of Economics, University of Navarra. Email: antmoreno@unav.es

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We estimate a New-Keynesian macro model accommodating regime-switching behavior in monetary policy and in macro shocks. Key to our estimation strategy is the use of survey-based expectations for inflation and output. Output and inflation shocks shift to the low volatility regime around 1985 and 1990, respectively. However, we also identify multiple shifts between accommodating and active monetary policy regimes, which play an as important role as shock volatility in driving the volatility of the macro variables. We provide new estimates of the onset and demise of the Great Moderation and quantify the relative role played by macro-shocks and monetary policy. The estimated rational expectations model exhibits indeterminacy in the mean-square stability sense, mainly because monetary policy is excessively passive.

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# 1 Introduction

The Great Moderation, the reduction in volatility (standard deviation) observed in most macro variables since the mid-1980s, makes it difficult to explain macroeconomic dynamics in the US over the last 40 years within a linear homoskedastic framework. There is still no consensus on whether the Great Moderation represents a structural break or rather a persistent but temporary change in regime. The causes also remain the subject of much debate. Was the Great Moderation the result of a reduction in the volatility of economic shocks, or was it brought about by a change in the propagation of shocks, for instance through a more aggressive monetary policy? Articles in favor of the “shock explanation” include McConnell and Perez-Quiros (2000), Sims and Zha (2006), Liu, Waggoner, and Zha (2011); articles in favor of the policy channel include Clarida, Galí, and Gertler (1999) and Galí and Gambetti (2008). Nevertheless, there is empirical evidence of both changes in the variance of economic shocks (Sims and Zha (2006)) and persistent changes in monetary policy (see Cogley and Sargent (2005), Boivin (2006), and Lubik and Schorfheide (2004)), necessitating an empirical framework that can accommodate both.

In this article, we estimate a standard New-Keynesian model accommodating regime changes in systematic monetary policy, in the variance of discretionary monetary policy shocks and in the variance of economic shocks. Whereas the model implies the presence of recurring regimes, it can also produce near permanent changes in regime. With the structural model, we can revisit the timing of the onset of the Great Moderation, and it so happens, also its demise. Moreover, we can then trace the sources of changes in the volatility of macroeconomic outcomes to changes in the volatility of demand, supply and discretionary monetary policy shocks, and to changes in systematic monetary policy. We find that output and inflation shocks moved to a lower variability regime in 1985 and 1990, respectively, but move back to the higher variability regime towards the end of 2008. Systematic monetary policy became more active after 1980, whereas discretionary monetary policy shocks were much less frequent after 1985<sup>1</sup>. The aggressive lowering of interest rates in the 2000-2005 period preceding the recent financial crisis is characterized as an activist regime. Put together, we identify the 1980-2007 period as a period with substantially lower output and inflation variability. From several perspectives, including counterfactual analysis, monetary policy was a critical driver of the Great Moderation.

While we retain the elegance of the theoretical Rational Expectations model, we make use of survey forecasts for inflation and GDP in the estimation. Ang, Bekaert, and Wei (2007) show that survey expectations beat any other model in forecasting future inflation out of sample. The use of survey forecasts not only brings additional information to bear on a complex estimation

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<sup>1</sup>Throughout the article we use active or activist policy to indicate the monetary policy regime where the interest rate reacts to expected inflation more than one to one, in contrast to passive monetary policy.

problem, but also simplifies the identification of the regimes under certain assumptions. In the extant literature, survey forecasts have mostly been used to provide alternative estimates of the Phillips curve (see Roberts (1995) and Adam and Padula (2011)). Instead, we study the role of survey expectations in shaping macroeconomic dynamics in the context of a standard NK model, accommodating regime switches.

While current medium-scale Dynamic Stochastic General Equilibrium (DSGE) models typically feature more variables and richer dynamics (see, for instance, Smets and Wouters (2007), Del Negro, Schorfheide, Smets, and Wouters (2007)), we deliberately focus on a small scale New-Keynesian model with an output gap, inflation, and interest rate equation for several reasons. First of all, this is the first attempt to estimate a small-scale DSGE model with survey-based expectations, which by themselves comprise very valuable information about a large set of variables. As a result, it is both instructive and relevant to focus on a relatively simple benchmark which also facilitates comparing estimation results with previous studies. Second, the model is rich enough to capture the time-varying role of both monetary policy and the key shocks shaping the Great Moderation in terms of output and inflation. Medium-scale models incorporating capital and labor explicitly may account for output fluctuations better than our model, but we conjecture that the identification of inflation dynamics, monetary policy, and the Great Moderation would not be greatly affected. Third, the estimation of even a stylized model with a realistic number of regimes remains actually very complex. Part of our contribution is to embed survey forecasts in the estimation and to obtain a Markov-Switching Rational Expectations (MSRE) Equilibrium, applying recent results in Cho (2013).

There has been recent progress in DSGE models incorporating regime-switching and time variation of structural parameters, including Bikbov and Chernov (2013) and Fernández-Villaverde, Guerrón-Quintana, and Rubio-Ramírez (2010). These articles use very different identification strategies, and do not make use of survey expectations. Apart from our modeling differences, which we discuss further in the model section, our analysis of the rational expectations equilibrium in a Markov-switching New-Keynesian model extends Davig and Leeper (2007) to an empirically more realistic setting. Our paper is closely related to Bianchi (2013), who also estimates a regime-switching New-Keynesian model. His model is a medium-scale DSGE model which differentiates the effects of macro shocks on consumption and investment. In his model, all macro shocks switch simultaneously, whereas we allow shocks to switch independently. As a result, our model displays many more different regimes (16) than his (4). Because the origin of supply, demand, and monetary policy shocks is by definition very different, we view our specification as more realistic. As Bianchi (2013) does, we find a stabilizing switch towards active monetary policy in the early 80s, but he only identifies one more switch in monetary policy towards the end of the sample whereas we identify several MP switches. Liu, Waggoner, and Zha (2011) also estimate a New-Keynesian

model with switches in shocks and the inflation target, but do not accommodate switches in policy response coefficients, which we identify as key to explain historical U.S. macro dynamics.

None of the aforementioned studies analyzes determinacy, an important characteristic of rational expectations models. For example, Lubik and Schorfheide (2004) document indeterminacy in the pre-Volcker period and discuss the estimation biases arising when indeterminate equilibria are excluded. Applying the methodology developed by Farmer, Waggoner, and Zha (2011), we find the estimated New-Keynesian model to be indeterminate in the mean-square stability sense. Davig and Leeper (2007) and Farmer, Waggoner, and Zha (2009) have previously shown that a temporarily passive monetary policy can be admissible as a part of a determinate equilibrium in simple calibrated MSRE models. However, in our more complex model featuring endogenous persistence, the actual policy stance in the passive regime for the U.S. economy during the 1968-2008 period is estimated to be excessively passive relative to the active regime, thereby causing indeterminacy. The recent return to a passive regime also contributed to the end of the Great Moderation. We then examine what policy parameter configurations would ensure a determinate equilibrium.

Section 2 describes the New-Keynesian model, detailing the role of regime-switching and expectations formation. Section 3 discusses the data and estimation method. Section 4 presents the empirical results, emphasizing the parameter estimates and the identified regimes. Section 5 concludes.

## 2 The Model

### 2.1 The basic New-Keynesian model

While our methodology is more generally useful, we focus attention on the following three-variable-three-equation New-Keynesian macro model, a benchmark of much recent monetary policy and macroeconomic analysis:

$$\pi_t = \delta E_t \pi_{t+1} + (1 - \delta) \pi_{t-1} + \lambda y_t + \varepsilon_{\pi,t}, \quad \varepsilon_{\pi,t} \sim N(0, \sigma_\pi^2) \quad (1)$$

$$y_t = \mu E_t y_{t+1} + (1 - \mu) y_{t-1} - \phi(i_t - E_t \pi_{t+1}) + \varepsilon_{y,t}, \quad \varepsilon_{y,t} \sim N(0, \sigma_y^2) \quad (2)$$

$$i_t = \rho i_{t-1} + (1 - \rho) [\beta E_t \pi_{t+1} + \gamma y_t] + \varepsilon_{i,t}, \quad \varepsilon_{i,t} \sim N(0, \sigma_i^2) \quad (3)$$

where  $\pi_t$  is the inflation rate,  $y_t$  is the output gap and  $i_t$  is the nominal interest rate.  $E_t$  is the conditional expectations operator. The three equations are subject to aggregate supply (AS), aggregate demand (IS) and monetary policy shocks, respectively. We denote these shocks by  $\varepsilon_{\pi,t}$  (AS-shock),  $\varepsilon_{y,t}$  (IS-shock), and  $\varepsilon_{i,t}$  (monetary policy shock). The  $\delta$  and  $\mu$  parameters represent the degree of forward-looking behavior in the AS equation (reflecting firm behavior) and IS equation

(reflecting consumer behavior), respectively (see Woodford (2003)). If they are not equal to one the model features endogenous persistence. The  $\phi$  parameter measures the impact of changes in real interest rates on output and  $\lambda$  the effect of output on inflation. The monetary policy reaction function is a forward-looking Taylor rule with smoothing parameter  $\rho$ . While policy rules featuring contemporaneous rather than expected inflation are still popular (see e.g. Fernández-Villaverde, Guerrón-Quintana, and Rubio-Ramírez (2010)), it is well accepted that policy makers consider expected measures of inflation in their policy decisions (see Bernanke (2010), Boivin and Gianoni (2006)). Policy should not react to temporary shocks that affect the contemporaneous rate of inflation, but not the future path of inflation.

The model is a simple example of a Dynamic Stochastic General Equilibrium (DSGE) macro model, characterized by a set of difference equations where today's decisions are a function of expected future macro variables as well as lags of the endogenous variables. These equations represent the log-linearized first-order conditions of the optimizing problems faced by a representative agent, firms, and the monetary authority. In matrix form, the model can be expressed as:

$$AX_t = BE_t X_{t+1} + DX_{t-1} + \varepsilon_t, \quad \varepsilon_t \sim N(0, \Sigma) \quad (4)$$

where  $X_t$  is the vector of macro variables and  $\varepsilon_t$  is the vector of structural macro shocks.  $A$ ,  $B$ , and  $D$  are matrices of structural parameters and  $\Sigma$  is the diagonal variance matrix of  $\varepsilon_t$ . Throughout this article, we focus on a rational expectations equilibrium (REE, henceforth) that depends only on the minimum state variables following McCallum (1983), also referred to as a fundamental solution. The solution to model (4) then follows a VAR(1) law of motion:

$$X_t = \Omega X_{t-1} + \Gamma \varepsilon_t \quad (5)$$

where  $\Omega$  and  $\Gamma$  are highly non-linear functions of the structural parameters, which can be solved following Klein (2000), Sims (2002), or Cho and Moreno (2011). We postpone discussion of the characterization of the rational expectations equilibria to Section 2.3.

Macro models often have a hard time fitting non-linear macro dynamics. While there are many potential reasons for this, we focus on two. First, there is considerable evidence of parameter instability. As noted by Clarida, Galí, and Gertler (1999), the monetary authority may have learned over time to react more aggressively to inflation deviations from target in order to tame output and inflation fluctuations, leading to instability in the systematic monetary policy parameters. In addition, the Great Moderation literature (McConnell and Perez-Quiros (2000), Cogley and Sargent (2005), Fernández-Villaverde and Rubio-Ramírez (2007), and Sims and Zha (2006)) shows that the output shocks identified in both reduced-form and structural models are heteroskedastic, displaying a pronounced decline after the mid 1980s. As a result, econometricians have tried to accommodate

these parameter changes through subsample analysis (Clarida, Galí, and Gertler (1999), Moreno (2004), and Boivin and Giannoni (2006)), time varying structural parameter and volatility estimation (Kim and Nelson (2006), Fernández-Villaverde and Rubio-Ramírez (2007), Ang, Boivin, Dong, and Loo-Kung (2010)) or through regime-switching models (Bikbov and Chernov (2013) and Sims and Zha (2006)). We incorporate regime-switching behavior in both systematic monetary policy and the variances of the structural shocks. The other parameters are assumed time invariant because they arise from micro-founded models.

Second, the rational expectations assumption may constrain the ability of the current generation of macro models to characterize macro dynamics. Chief among these shortcomings is the fact that agents only employ the variables used to construct the model in forming expectations of future macro variables. Given that most macro models only use a limited number of variables, the information sets used by RE agents seem to be unrealistically constrained<sup>2</sup>. There are a number of potential avenues to overcome this problem. The generalized method of moments (GMM) allows researchers to condition the estimation of model parameters on information sets which include additional variables to those implied by the model (see, for instance, Clarida, Galí, and Gertler (1999)). Boivin and Giannoni (2008) estimate a DSGE RE macro model, enhancing the information set available to agents for decision making purposes with a large number of macro variables governed by a factor structure. Bekaert, Cho, and Moreno (2010), Bikbov and Chernov (2013), and Rudebusch and Wu (2008) use term structure data to help identify a New-Keynesian macro model. The work of Bikbov and Chernov (2013) is most closely related to ours, as they also allow regime shifts in the shock variances and systematic monetary policy. However, their identification strategy is very different, as they use term structure data and an exogenous pricing kernel (inconsistent with the IS curve) to price the term structure.

Instead, we use survey-based expectations (SBE) to help identify the parameters of a DSGE macro model. SBE reflect the direct answers of a large number of economic agents to questions about the expected future path of macroeconomic variables. Unlike RE, SBE are thus not model conditioned and naturally reflect the different perceptions of economic agents based on a potentially very rich information set. Recently, several authors (Roberts (1995), Adam and Padula (2011) and Nunes (2010)) have estimated New-Keynesian Phillips curves using SBE. The results

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<sup>2</sup>Moreover, RE imply that all agents have a perfect knowledge of the model and only adjust their expectations in reaction to the model dynamics in order to reach the equilibrium, leaving no room for any alternative perceptions or mechanisms which in practice would likely alter their decisions. According to Solow (2004), Phelps (2007) and Woodford (2013), this tight endogeneity of the RE framework may impair its ability to explain macro dynamics. On the theoretical side, De Grauwe (2008) develops a DSGE model where agents exhibit bounded rationality, whereas Sims (2005) introduces the rational inattention concept, relaxing some of the RE assumptions. In addition, Onatski and Stock (2002), among others, develop techniques to perform policy analysis in the presence of model, parameter and shock uncertainty around a reference model, thus leaving some room for macro realizations to deviate from a benchmark model with perfectly known parameters and forcing processes.



of these efforts have overall been positive, as the estimate of the important Phillips curve parameter, linking inflation to the output gap, becomes statistically significant under SBE, in contrast to the results produced by most RE models. Nevertheless, the use of SBE in DSGE macro models has been limited to date and restricted to single-equation estimation. Of course, there is much skepticism about SBE: agents may not be truth-telling or may omit important information in forming forecasts of future macro variables. However, Ang, Bekaert, and Wei (2007) show that SBE of inflation predict inflation out-of-sample better than a large number of the standard structural and reduced-form inflation models proposed in the literature.<sup>3</sup> Consequently, SBE likely contain important information about future macro variables. We show below that incorporating SBE greatly facilitates the computation of the likelihood function and thus the identification of the regime shifts.

## 2.2 Introducing regime switches

We postulate the presence of 4 regime variables, to model regime shifts in the nature of systematic monetary policy and in the variances of the structural shocks. The first variable  $s_t^{mp}$  switches  $\beta$  and  $\gamma$  in equation (3), which represent the systematic monetary policy parameters. The second variable  $s_t^\pi$  shifts the volatility of the aggregate supply shocks. The third variable  $s_t^y$  shifts the volatility of the IS shocks. The fourth variable  $s_t^i$  affects the volatility of the monetary policy shock. These variables can take on two values and follow Markov chains with constant transition probabilities in the Hamilton (1989) tradition. The agents are assumed to know the regime at each point in time so that learning issues are dispensed with. In particular, agents rationally account for potential future regime shifts in monetary policy when taking expectations. We assume that the regime variables are independent. For future reference, let  $S_t = (s_t^{mp}, s_t^\pi, s_t^y, s_t^i)$ .

The regime-dependent volatility model for the three shocks in equation (4) simply allows for two different values of the conditional variance, as a function of the regime variable. For example, for the AS equation, we have:

$$Var(\varepsilon_{\pi,t} | X_{t-1}, S_t) = \sigma_\pi^2(s_t^\pi) = \exp(\alpha_{\pi,0} + \alpha_{\pi,1}s_t^\pi) \quad (6)$$

with  $s_t^\pi = 1, 2$  and the exponential function guaranteeing non-negative volatilities. We adopt the convention that the variance in regime 1 is higher than the variance of regime 2 for each structural shock:  $\sigma_\pi^2(s_t^\pi = 1) > \sigma_\pi^2(s_t^\pi = 2)$ ,  $\sigma_y^2(s_t^y = 1) > \sigma_y^2(s_t^y = 2)$ ,  $\sigma_i^2(s_t^i = 1) > \sigma_i^2(s_t^i = 2)$ .

The regime variable  $s_t^{mp}$  accommodates potential persistent shifts in the systematic policy parameters  $\beta$  and  $\gamma$ . In particular, we expect to find an activist regime with  $\beta$  larger than 1 and a

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<sup>3</sup>Boivin (2006), for instance, uses the Greenbook forecasts employed before each FOMC meeting by the Fed in order to identify changes in its stance against inflation. These forecasts include information from a wide range of sources, including forecasters' opinions.

passive regime with  $\beta$  smaller than 1. A number of economists (Clarida, Galí, and Gertler (1999), Boivin and Giannoni (2006)) suggest that  $\beta$  experienced a structural break around 1980, with  $\beta$  being lower than 1 before and larger thereafter. While we find such a model *ex ante* implausible, it can still be approximated by our regime-switching model if the regimes are very persistent with very small transition probabilities. Nevertheless, in our model, a switch to a new regime is never viewed as permanent. If regime classification yields a passive regime 100% of the time before 1980, and an activist regime 100% of the time afterwards, the permanent break hypothesis surely gains credence relative to a model of persistent but non-permanent changes in policy. It is also possible that the influential 1979-1982 Volcker period affects inference substantially. Was this period the first switch into a more active regime or is it best viewed as a period of discretionary contractionary policy? By letting the variable  $s_t^i$  affect the variability of the monetary policy shock, we also accommodate the latter possibility.

Incorporating the regime variables, equation (4) becomes:

$$A(S_t)X_t = B(S_t)E_t X_{t+1} + DX_{t-1} + \varepsilon_t, \quad \varepsilon_t \sim N(0, \Sigma(S_t)) \quad (7)$$

where  $A(S_t)$  and  $B(S_t)$  capture the regime-switching behavior of the central bank and  $\Sigma(S_t)$  governs the time-varying variances of the structural shocks. With regimes affecting both systematic monetary policy and the variance of shocks, we can use the model to revisit the question of what drove down inflation and output growth variability during the 1980s and 1990s: was it policy or luck (see e.g. Stock and Watson (2002) and Blanchard and Simon (2001))? A large literature has examined this issue from both reduced-form (Cogley and Sargent (2005), McConnell and Perez-Quiros (2000), Sims and Zha (2006)) and structural (Moreno (2004), Lubik and Schorfheide (2004), Boivin and Giannoni (2006), and Inoue and Rossi (2011)) perspectives. Disagreement remains. For instance, Benati and Surico (2009) show that the results of Sims and Zha (2006), suggesting a prominent role for heteroskedasticity, may be biased against finding a role for policy changes. The combination of a structural New-Keynesian model with regime shifts in both monetary policy parameters and shock variables can provide novel evidence on the sources of macroeconomic variability.

Our model fits into a rapidly growing body of work incorporating policy changes and/or heteroskedasticity into New-Keynesian models. Part of this literature is more theoretical in nature, considering issues of equilibrium existence and stability, in models that are not likely to be empirically successful. We discuss this important literature in Section 2.3. The empirical literature on DSGEs with time-varying parameter and shock distribution is very recent. Some authors, such as Fernández-Villaverde and Rubio-Ramírez (2007) and Justiniano and Primiceri (2008), postulate heteroskedastic variances and fixed structural parameters in their DSGEs, whereas Davig and

Doh (2013) develop a New-Keynesian model with regime-switching parameters but constant shock variances. Fernández-Villaverde, Guerrón-Quintana, and Rubio-Ramírez (2010), Bianchi (2013) and Bikbov and Chernov (2013) allow for time variation in both the structural shock variances and the systematic part of their RE New-Keynesian macro models, and are thus closest to our framework. Bianchi (2013) uses only one regime variable to accommodate heteroskedasticity. We show below that this is overly restrictive. Our use of SBE also allows for a much simpler estimation method than is possible in Bianchi (2013).

### 2.3 The rational expectations equilibrium under regime-switching

A linear rational expectations model (4) is said to be determinate if it has a unique and stable (non-explosive) equilibrium, which takes the form of a fundamental REE as in equation (5). In case of indeterminacy, the models generally have multiple fundamental and non-fundamental (“sunspot”) equilibria. It is now well-understood that a violation of the Taylor principle, typically identified as  $\beta$  being less than 1 in equation (3), leads to indeterminate equilibria in the prototypical New-Keynesian model. Intuitively, raising the short-term nominal interest rate less than one for one to an increase in (expected) inflation actually lowers the real rate, fueling inflation even more through output gap expansion and the Phillips curve mechanism. However, the US data seem to suggest a structural break in  $\beta$ , with  $\beta$  lower than 1 (“passive policy”) before 1980 and higher than 1 (“active policy”) afterwards (Clarida, Galí, and Gertler (1999), Boivin and Giannoni (2006)). From the perspective of a standard New-Keynesian model, this implies that the propagation system was not uniquely determined before 1980 and/or that non-fundamental (sunspot) equilibria may have played a role before 1980 (see Lubik and Schorfheide (2004)).

Recently, Davig and Leeper (2007) generalized the Taylor principle to a baseline New-Keynesian macro model with regime-switching in monetary policy, which is nested in the model of equation (7). Specifically, they show that the model can have a unique stable equilibrium even when the central bank is temporarily passive as long as there is a positive probability that the passive regime switches to the active regime, and the structural shocks are bounded. Consequently, a Markov-switching rational expectations model (MSRE for short), apart from being more economically reasonable than a permanent break model, offers the potential to explain US macro-dynamics, even before 1980, in the context of a model with a unique and stable equilibrium.

It is therefore important to fully characterize the stability and determinacy of the equilibrium for the US economy. To do so, we must define and verify a stability concept (to rule out unstable solutions) and identify fundamental solutions to the model. Following Farmer, Waggoner, and Zha (2010), we adopt mean-square stability as the primary concept of stability in this paper. Hence, we require the first and second moments of  $X_t$  to be finite. The bounded stability concept that Davig

and Leeper (2007) and Benhabib (2009) use essentially requires bounded random variables, and is not suitable in our framework for two reasons. Determinacy conditions under bounded stability have not been established for models with predetermined variables, and the support of structural shocks in our model is unbounded as they follow normal distributions.

Following Farmer, Waggoner, and Zha (2009, 2011), we express the general solution to our model (7) as a sum of a fundamental solution and a non-fundamental (sunspot) component:

$$X_t = \Omega(S_t)X_{t-1} + \Gamma(S_t)\varepsilon_t + u_t, \quad (8)$$

$$s.t. u_t = F(S_t)E_t u_{t+1} \quad (9)$$

where the first two components in (8) represent a fundamental solution given by equation (7) and  $u_t$  is a sunspot component. Note that the state variables in this model are the vector of lagged endogenous variables,  $X_{t-1}$ , the vector of the exogenous variables,  $\varepsilon_t$  and the current set of regimes  $S_t$ . The restrictions that  $\Omega(S_t)$ ,  $\Gamma(S_t)$  and  $F(S_t)$  must satisfy in a rational expectations equilibrium can be easily derived by plugging equation (8) into equation (7). They are given explicitly in Appendix A. Determinacy then requires two conditions: the uniqueness of the stable fundamental solution and the non-existence of stable sunspot components.

Because the general MSRE model with lagged endogenous variables lacks an eigensystem, establishing determinacy conditions is far from trivial, and we cannot rely on the extant literature. For example, Farmer, Waggoner, and Zha (2011) propose a method to identify model solutions using a numerical procedure. However, since the number of fundamental solutions is unknown, such a procedure cannot really establish the first determinacy condition. Furthermore, Farmer, Waggoner, and Zha (2009) propose a condition for the second determinacy requirement, but it is only valid in models without lagged variables and therefore does not apply to our model either. We therefore rely on relatively new results in Cho (2013) who generalizes the “forward method” introduced by Cho and Moreno (2011) for linear models. The forward solution for a linear RE model results from solving the model recursively forward. The forward solution is the unique fundamental solution that satisfies no-bubble (or transversality) condition; the condition that makes the expectations of the present value of future endogenous variables converge to zero. Consequently, the forward solution selects an economically reasonable fundamental equilibrium and delivers the numerical solution in one step. Importantly, Cho (2013) shows that this logic carries over to MSRE models and develops very tractable determinacy conditions in the mean-square stability sense for general MSRE models with predetermined variables, which we rely upon here. Appendix A contains technical details about the methodology.

To develop some intuition, consider an  $n$ -dimensional linear RE model, a special case of our model in the absence of regime-switching, so that all the coefficient matrices of the model and the

solution are constant. This linear RE model is known to have  $2n$  generalized eigenvalues and it is determinate if there exist exactly  $n$  stable roots (see Klein (2000), for instance). McCallum (2007) shows that the  $n$  roots of  $\Omega$  in equation (8) and the reciprocals of the roots of the associated  $F$  in equation (9) constitute those  $2n$  generalized eigenvalues. Using this observation, determinacy can be equivalently stated as follows: the linear RE model is determinate if there exists an  $\Omega$  and its associated  $F$  such that

$$r(\Omega) < 1 \text{ and } r(F) \leq 1, \quad (10)$$

where  $r(\cdot)$  is the spectral radius, the maximum absolute eigenvalue of the argument matrix.<sup>4</sup> The latter condition has a straightforward interpretation from  $u_t = FE_t u_{t+1}$  in equation (9):  $r(F) \leq 1$  implies that the second determinacy condition holds so that there is no stable sunspot component  $u_t$  (the expected sunspot is explosively related to the current sunspot as the inverse of  $F$  has unstable eigenvalues). This condition in conjunction with the first condition regarding  $\Omega$  then ensures that there is unique stable fundamental solution, hence the model is determinate.

In a MSRE set-up, these conditions will have to take into account that there will be transitions between different regimes and thus between different coefficient matrices. Focussing on a simple model without lagged state variables, Farmer, Waggoner, and Zha (2009) show that the determinacy conditions involve transition probabilities and the second moments of the variables. Cho (2013) derives the determinacy conditions for more general MSRE models, which are analogous to the conditions in equation (10). In particular, our model (7) is determinate if there exists a solution of the form (8)-(9) such that

$$r(\bar{D}_\Omega) < 1 \text{ and } r(D_F) \leq 1, \quad (11)$$

where  $\bar{D}_\Omega$  and  $D_F$  are the transition probability weighted matrices defined as<sup>5</sup>:

$$\bar{D}_\Omega = \begin{bmatrix} p_{11}\Omega(1) \otimes \Omega(1) & p_{21}\Omega(2) \otimes \Omega(2) \\ p_{12}\Omega(1) \otimes \Omega(1) & p_{22}\Omega(2) \otimes \Omega(2) \end{bmatrix}, \quad D_F = \begin{bmatrix} p_{11}F(1) \otimes F(1) & p_{12}F(1) \otimes F(1) \\ p_{21}F(2) \otimes F(2) & p_{22}F(2) \otimes F(2) \end{bmatrix}. \quad (12)$$

with  $\Omega(i)$ ,  $F(i)$ , for  $i = 1, 2$ , denoting the coefficient matrices associated with regime  $i$ . Recall that in equations (8) and (9) the state  $S_t$  only depends on the monetary policy regime  $s_t^{mp}$ , so the two states 1 and 2 represent the active and passive regimes, respectively. Therefore the probabilities in equation (12), represent transition probabilities between active and passive policy regimes,  $p_{ij} = P[S_{t,mp} = i | S_{t-1,mp} = j]$ . Therefore, to check for determinacy, the matrices in equation (12) must be computed and their spectral radii checked.

<sup>4</sup>The first condition in (10) implies that there are  $n$  stable generalized eigenvalues and the latter condition implies that the remaining  $n$  generalized eigenvalues are unstable.

<sup>5</sup>Notice that in the absence of regime switching, the argument matrices collapse to  $\bar{D}_\Omega = \Omega^2$  and  $D_F = F^2$ . But, since  $r(\Omega^2) < 1$  if and only if  $r(\Omega) < 1$ , and the same is true for  $D_F$ , the conditions in (11) are equivalent to those in (10). Therefore, our determinacy conditions are a natural extension of those for linear models to MSRE models.

Cho (2013) simplifies this process making use of the so-called forward solution. As in the linear RE model,  $r(D_F) \leq 1$  implies the non-existence of stable sunspot components. He shows that when checking this condition for the forward solution, its violation implies that all the other fundamental solutions are unstable. Therefore, stability of the forward solution, i.e.,  $r(\bar{D}_\Omega) < 1$  directly implies determinacy and the forward solution is the determinate solution. While the determinacy conditions in linear (equation 10) and MSRE (equation 11) models appear analogous, the actual derivations and proofs for the MSRE case are somewhat involved and we relegate further technical details to Appendix A.

Building on these recent results, it is straightforward to verify whether a given model has stable fundamental solutions. We rely on these techniques to define a reasonable compact parameter space for our estimation problem, in which it is likely that a stable RE equilibrium exists. To do so (and to aid our practical estimation), we conduct an extensive study of the existence of RE equilibria for different parameter configurations. The analysis is described in more detail in Appendix B, but we provide a short summary of the major findings here. Essentially, we conduct a grid search over an extensive parameter range, and verify whether we can characterize the set of parameters for which a fundamental forward solution exists. This proved a non-trivial task and no simple characterization is possible. However, the most critical parameters in driving the existence of a RE equilibrium clearly are  $(\delta, \mu, \beta_1, \beta_2)$ . Recall that  $\beta_1 > \beta_2$ , identifying the first regime as the “active” regime. Not surprisingly, given Davig and Leeper’s work, an equilibrium can still exist with  $\beta_2$  smaller than 1, and  $\beta_1$  larger than 1. Values of  $\mu$  and  $\delta$  smaller than 0.5 will lead to non-existence, but an equilibrium may exist if only one of the two is smaller than 0.5 (and the other one relatively high).

We use this information to consider a restricted parameter space for the estimation (see more below). Nevertheless, estimating the model in equation (7) with a relatively large number of regime variables remains difficult. In order to construct the likelihood function, we must not only integrate across all combinations of potential (unobserved) regimes, but also numerically compute the highly non-linear reduced-form coefficient matrices  $(\Omega(S_t)$  and  $\Gamma(S_t))$  for all combinations of potential regimes. We circumvent this problem and simultaneously bring additional information to bear on the estimation by incorporating survey forecasts, as we show in the next subsection.

## 2.4 Introducing survey expectations

Undoubtedly, the information used by professional forecasters greatly exceeds the information set spanned by the variables present in the simple model in equations (1)-(3). Given that survey expectations outperform empirical and theoretical models predicting inflation, they can also prove useful in estimating macroeconomic parameters and dynamics. To incorporate SBE into the model,

we assume that survey expectations of inflation and output obey the following law of motion:

$$\pi_t^f = \alpha E_t \pi_{t+1} + (1 - \alpha) \pi_{t-1}^f + w_t^\pi \quad (13)$$

$$y_t^f = \alpha E_t y_{t+1} + (1 - \alpha) y_{t-1}^f + w_t^y \quad (14)$$

with  $w_t^\pi \sim N(0, \sigma_f^\pi)$  and  $w_t^y \sim N(0, \sigma_f^y)$ . Consequently, survey expectations potentially react to true rational expectations one for one if  $\alpha$  equals 1, but may also slowly adjust to true rational expectations and depend on past survey expectations. This is reminiscent of Mankiw and Reis (2002)'s model of the Phillips curve, in which information disseminates slowly throughout the population.

In our model, we combine the determination of SBE with the regime-switching counterparts of equations (1)-(3). That is, we retain the assumption of rational expectations, and simply use additional information to identify both the structural parameters and the regimes in a 5 variable system. Nevertheless, the estimation remains complex as we still need to solve the rational expectations equilibrium at each step in the optimization and for all possible regime combinations. If we let the variance of the shocks in equations (9)-(10) go to zero, so that SBE are an exact function of past SBE and current RE, we can greatly simplify estimation. In this case, we can infer the RE of inflation and output from equations (13) and (14) and substitute them into the main model equations to obtain:

$$\pi_t = \frac{\delta}{\alpha} (\pi_t^f - (1 - \alpha) \pi_{t-1}^f) + (1 - \delta) \pi_{t-1} + \lambda y_t + \varepsilon_{\pi,t}, \quad \varepsilon_{\pi,t} \sim N(0, \sigma_\pi^2(s_t^\pi)) \quad (15)$$

$$y_t = \frac{\mu}{\alpha} (y_t^f - (1 - \alpha) y_{t-1}^f) + (1 - \mu) y_{t-1} - \phi i_t + \frac{\phi}{\alpha} (\pi_t^f - (1 - \alpha) \pi_{t-1}^f) + \varepsilon_{y,t}, \quad (16)$$

$$\varepsilon_{y,t} \sim N(0, \sigma_y^2(s_t^y))$$

$$i_t = \rho i_{t-1} + (1 - \rho) \left[ \frac{\beta(s_t^{mp})}{\alpha} (\pi_t^f - (1 - \alpha) \pi_{t-1}^f) + \gamma(s_t^{mp}) y_t \right] + \varepsilon_{i,t}, \quad (17)$$

$$\varepsilon_{i,t} \sim N(0, \sigma_i^2(s_t^i))$$

Notice that when  $\alpha = 1$ , the RE are assumed equivalent with SBE. We ask the data to gauge the wedge between those two expectations. The parameter  $\alpha$  generally measures the relative weight of RE and past SBE in expectation formation for professional forecasters.

Let  $X_t^f = [\pi_t^f \ y_t^f]'$ . In matrix form, the regime-switching New-Keynesian model becomes:

$$A(S_t) X_t = B(S_t) X_t^f + D(S_t) X_{t-1}^f + F X_{t-1} + \varepsilon_t, \quad \varepsilon_t \sim N(0, \Sigma(S_t)) \quad (18)$$

with:

$$A(S_t) = \begin{bmatrix} 1 & -\lambda & 0 \\ 0 & 1 & \phi \\ 0 & -(1-\rho)\gamma(s_t^{mp}) & 1 \end{bmatrix}, \quad B(S_t) = \begin{bmatrix} \frac{\delta}{\alpha} & 0 \\ \frac{\phi}{\alpha} & \frac{\mu}{\alpha} \\ \frac{(1-\rho)}{\alpha}\beta(s_t^{mp}) & 0 \end{bmatrix},$$

$$D(S_t) = \begin{bmatrix} -\frac{\delta(1-\alpha)}{\alpha} & 0 \\ -\frac{\phi(1-\alpha)}{\alpha} & -\frac{\mu(1-\alpha)}{\alpha} \\ -\frac{(1-\rho)(1-\alpha)}{\alpha}\beta(s_t^{mp}) & 0 \end{bmatrix}, \quad F = \begin{bmatrix} (1-\delta) & 0 & 0 \\ 0 & (1-\mu) & 0 \\ 0 & 0 & \rho \end{bmatrix},$$

and conditional on  $\alpha \neq 0$ ,

$$\Sigma(S_t) = \begin{bmatrix} \sigma_{AS}(s_t^\pi) & 0 & 0 \\ 0 & \sigma_{IS}(s_t^y) & 0 \\ 0 & 0 & \sigma_{MP}(s_t^i) \end{bmatrix}.$$

This model leads to the following reduced-form:

$$X_t = \Omega_1(S_t)X_{t-1}^f + \Omega_2(S_t)X_t^f + \Omega_3(S_t)X_{t-1} + \Gamma(S_t)\varepsilon_t, \quad \varepsilon_t \sim N(0, \Sigma(S_t)), \quad (19)$$

with  $\Omega_1(S_t) = A(S_t)^{-1}B(S_t)$ ,  $\Omega_2(S_t) = A(S_t)^{-1}D(S_t)$ ,  $\Omega_3(S_t) = A(S_t)^{-1}F$  and  $\Gamma(S_t) = A(S_t)^{-1}$ . A major advantage of this approach is that the matrices determining the law of motion of  $X_t$  are simple analytical functions of the structural parameters, thus making the likelihood function much easier to compute, simplifying estimation. There is no need to compute the RE equilibrium at each step in the optimization of the likelihood, and the regimes can be inferred as in the standard reduced-form multivariate models (see Hamilton (1989) and Sims and Zha (2006)). Importantly, SBE adds new information, absent in the variables and structure of the New-Keynesian model, to aid parameter estimation.

### 3 Data and Estimation

The model requires analogs for five variables: inflation, the output gap, the short-term interest rate, and survey-based estimates of expected inflation and the expected output gap. Inflation is measured as the log-difference of the chain-type Gross Domestic Product (GDP) deflator price index. The output measure is real GDP and we employ a quadratic trend to measure potential output. The output gap is then defined in the usual fashion as the percentage deviation of output relative to trend. We report results for the output gap measure computed using a quadratic trend. We retrieve both the GDP and GDP deflator data from the Bureau of Economic Analysis database. Expected inflation is the median survey responses of expected GDP inflation over the next quarter.



To construct the expected output gap, we use current GDP, the predicted trend, and expected GDP growth over the next quarter. We again use the median survey response to proxy for expected GDP growth. Both expected inflation and output are from the Survey of Professional Forecasters (SPF) published by the Federal Reserve Bank of Philadelphia. Finally, the short-term interest rate is the 3-month treasury bill (secondary market rate). The data frequency is quarterly and our sample period goes from the fourth quarter of 1968 to the second quarter of 2008. Appendix C has more details on the data and the variables construction.

The model in (19) is estimated via limited information maximum likelihood, given that we do not use the  $\pi_t^f$  and  $y_t^f$  equations. The information set  $I_{t-1}$  consists of all the available information up to time  $t-1$ :  $I_{t-1} = \{Q_{t-1}, Q_{t-2}, \dots, Q_0\}$ , where  $Q_t = [X_t \ X_t^f]'$ . The full dataset is thus  $\tilde{Q}_T = [Q_T, Q_{T-1}, \dots, Q_0]$ . We denote the parameters to be estimated as  $\theta$ , so that the aim is to maximize the density function  $f(\tilde{X}_T; \theta)$ . While agents in the economy observe the regime variables,  $S_t$ , the econometrician does not and only has data on  $Q_t$ . Therefore, we maximize the likelihood function for  $\tilde{X}_T$ , integrating out the dependence on  $S_t$ , as is typical in the regime-switching literature<sup>6</sup>. In particular, note that the regime variable  $S_t$  can take on 16 values ( $2^4$ ). We can rewrite the conditional likelihood at  $t$  as:  $f(X_t|I_{t-1}) = \sum_{i=1}^{16} P(S_t = i|I_{t-1}) f(X_t|I_{t-1}, S_t = i)$ . This decomposition allows evaluating the conditional density using equation (19). The regime dependent likelihoods are weighted by the so-called “ex-ante” regime probabilities, which can be easily created recursively, as described in Hamilton (1994). After identifying the parameters, the econometrician can make inferences about the regimes by computing the “smoothed” regime probabilities, which are the probability of the regime given full sample information  $I_T = \tilde{Q}_T$ . We use the well-known recursive algorithm developed in Kim (1994) and described in Hamilton (1994) to compute these probabilities. A well-identified regime switching model should produce smoothed regime probabilities that are either close to zero or close to one (see e.g. Ang and Bekaert (2002)). With two possible regimes, a smoothed regime probability of 0.5 indicates the econometrician has failed to identify the regime.

We would like the estimation to produce parameters for which a fundamental rational expectations equilibrium exists. To do so, we proceed in two steps. First, we use the analysis of Section 2.3 to construct a compact parameter space that attempts to exclude regions where REEs are unlikely to exist. Because of the non-convexity of the set, we use a rather wide parameter space (details are available upon request), that encompasses the parameter values yielding a REE. Second, at each step in the optimization, we verify whether the forward solution exists. If not, the likelihood function is penalized, steering optimization away from such regions in the parameter space.

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<sup>6</sup>We sacrifice full efficiency by ignoring  $f(X_t^f|I_{t-1}; \theta)$  in the estimation. Technically, this requires assuming  $f(X_t^f|S_t, I_{t-1}; \theta) = f(X_t^f|I_{t-1}; \theta)$ . While not very palatable at first, in our model, the regimes can in principle be identified without using survey data, so that the assumption is implicitly valid.

Appendix D describes the different specification tests that we perform on the residuals of the model. First, for each equation, we test the hypotheses of a zero mean and zero serial correlation (up to two lags) of the residuals (the “mean test”); unit mean and zero serial correlation (two lags) for the squared standardized residuals (the “variance test”); zero skewness, and appropriate kurtosis. In performing these tests, we recognize that the test statistics may be biased in small samples, especially if the data generating process is as non-linear as the model is above. Therefore, we use critical values from a small Monte Carlo analysis also described in Appendix D. Second, the economic model should also capture the correlation between the various variables. We test for each residual whether its joint covariances with all other residuals are indeed zero. We also perform a joint test for all covariances. As in the first set of tests, we obtain critical values from a small Monte Carlo analysis.

Table 1 reports Monte Carlo  $p$ -values of all these tests for our main model, on the left hand side. The residual levels and variances are well behaved, with the exception of the output gap, where the test uncovers some remaining autocorrelation in the residuals. The regime-switching model is able to capture most skewness and kurtosis in the data, only failing the zero skewness test for inflation. The model’s weakest point appears to be the fit of covariances between the three shocks. The last two lines in Table 1 reveal that the model fails to fully capture the correlation structure between the various economic variables.

## 4 Empirical Results

### 4.1 Parameter estimates

Table 2 presents the parameter estimates of the Regime-Switching DSGE New-Keynesian macro model yielding a stable fundamental RE equilibrium, as described in Section 2.3. It also shows a number of statistical tests of parameter equality. All parameters have the right sign and are statistically significant, but we did constrain the  $\phi$  coefficient to a positive value of 0.1. As is common in maximum likelihood estimation of this class of New-Keynesian models, unconstrained estimation yields either negative or very small and insignificant estimates of  $\phi$  (see Ireland (2001), Fuhrer and Rudebusch (2004) and Cho and Moreno (2006))<sup>7</sup>.

In the AS equation,  $\delta$  is 0.425, implying a similar weight on the forward-looking and endogenous persistence terms. The IS equation is more forward looking, since  $\mu$  is 0.675. Given the small standard errors of these parameters, our estimation reveals strong evidence in favor of endogenous persistence. Moreover, a Wald test strongly rejects the hypothesis that the degree of forward lookingness in the AS and IS equations can be captured with one coefficient (Rudebusch (2001);

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<sup>7</sup>Our results are qualitatively similar when we set  $\phi = 0.01$  or  $\phi = 0.20$ .

Gurkaynak, Sack, and Swanson (2005)).

The Phillips curve parameter  $\lambda$  is large at 0.102, implying a strong transmission mechanism from output to inflation and thus a strong monetary policy transmission mechanism. Previous estimations of rational expectations models fail to obtain reasonable and significant estimates of  $\lambda$  with quarterly data (Fuhrer and Moore (1995)). Some alternative estimations have yielded significant estimates, such as Galí and Gertler (1999) who use a measure for marginal cost replacing the output gap; Bekaert, Cho, and Moreno (2010) who identify a natural rate of output process from term structure data; or Roberts (1995) and Adam and Padula (2011) who use SBE but in a single equation context with fixed regimes. However, our estimate is even larger than the coefficients reported in these articles. We conjecture that the introduction of slow moving SBE of inflation generates additional correlation between (expected) inflation and the output gap.

Regarding the monetary policy rule, the interest rate persistence is large, 0.834, in agreement with most studies in the literature (Clarida, Galí, and Gertler (1999), Bekaert, Cho, and Moreno (2010), among others). Our estimation allows for regime switches in the key monetary policy parameters,  $\beta$ , the response to expected inflation, and  $\gamma$ , the response to the output gap. In the “activist” regime, which is the first regime,  $\beta$  is 2.312, well above 1 statistically, whereas in the passive regime,  $\beta$  is 0.598, significantly below 1. Thus, our estimation clearly identifies a sharp economic and statistical difference in the response to inflation across monetary policy regimes. In their single equation monetary policy rule estimation, Davig and Leeper (2005) also estimate a significant difference between  $\beta$ 's across regimes, but of a smaller magnitude than our estimates. The contemporaneous articles of Bikbov and Chernov (2013) and Bianchi (2013), estimating Markov-switching RE New-Keynesian models, also identify a large difference in  $\beta$  across regimes. The interest rate response to the output gap,  $\gamma$ , is higher than in the aforementioned estimations (1.187 and 0.687, respectively), and it is larger in the more “activist” regime, although not in a statistically significant way.

Finally,  $\alpha$ , the parameter governing the law of motion for the survey-based expectations, is 0.986, meaning that SBE adjust almost completely to RE. We examine below whether this finding is the result of imposing rational expectations on the estimation. Because the other parameters are directly related to the identification of the regimes, we discuss them in the next sub-section.

## 4.2 Macroeconomic regimes

The key output of our model is the identification of macroeconomic regimes. The volatility parameters imply strong evidence of time-varying variances in macroeconomic shocks. For the output gap and inflation shocks, volatility in the high volatility regime is around double that in the low volatility regime. However, for interest rates, the high volatility regime features volatility that is

about 6 times as high as in quiet times, suggesting a potentially important role for discretionary monetary policy. Because interest rates are measured in quarterly percent, the volatility of interest rate shocks in the low volatility state is very small (0.04%), implying a strict commitment to the monetary policy rule.

The transition probability coefficients imply overall quite persistent regimes. For inflation, the expected duration of the high variance regime is very high at 100 quarters, but the low variance regime is persistent as well. Output gap regimes are somewhat less persistent, with the high variance regime expected to last about 27 quarters, while discretionary interest rate regimes are much less persistent, with the high interest rate variability regime expected to last about 8 quarters. Accommodating monetary policy regimes last on average longer than activist regimes, which are short-lived lasting on average 7 quarters.

These transition probabilities are important inputs in the identification of the time path of the regimes. Figure 1 plots the smoothed probabilities for the four independent regime variables. Panel A shows the smoothed probabilities of respectively the high inflation shock volatility regime and the high output shock volatility regime. Note that the regime probabilities tend to be either close to one or zero, indicating adequate regime identification. We observe a sudden drop in output shock volatility starting in 1981 and fully materializing in 1985. The decreased volatility persists until 2007, coinciding with the onset of the credit crisis. The variability of inflation shocks starts to decrease later, with the smoothed probability going below 0.5 at the beginning of 1986, and going toward zero just before the 1990 recession. Signs of a reversal in the low variability regime are already visible in 2003, with its probability reaching less than 50 percent in the third quarter of 2006 already. Our evidence in favor of a switch towards a higher variability regime is stronger and its timing earlier than in Bikbov and Chernov (2013).

Panel B shows the smoothed probabilities of respectively the active monetary policy regime in which the Fed aggressively stabilizes inflation, and the high volatility regime for interest rate shocks. The high interest rate shock volatility regime occurs quite frequently and is always on during recessions, including during the 1980-1982 Volcker period. This implies that in times of recession, the Fed is more willing to deviate from the interest rate rule. Bikbov and Chernov (2013) also categorize the Volcker period as a period of discretionary monetary policy. Unlike their results, we also find systematic monetary policy to be activist during this period. Interestingly, our model shows that activist monetary policy spells generally became more frequent from 1980 onwards. We identify the 1993-2000 period as an accommodating monetary policy stance. Because this period is characterized by relatively low inflation, a passive monetary policy stance implies relatively high interest rates. One interpretation is that inflation expectations were firmly anchored, due to the more aggressive stance of the Fed during the previous decade. In addition, the possibility of switching back to the stabilizing regime, as captured by our regime-switching DSGE, may also

anchor inflation expectations. Notice that this regime identification is quite different from the permanent shift in monetary policy around 1980, put forward in earlier studies such as Clarida, Galí, and Gertler (1999) and Lubik and Schorfheide (2004), but consistent with contemporaneous results in Fernández-Villaverde, Guerrón-Quintana, and Rubio-Ramírez (2010).

In 2000 there is a switch to the activist regime, as interest rates rapidly declined following the beginning of the 2000 recession, while inflation stayed low. Hence, according to our analysis, interest rates in the first 5 years of the previous decade were lower than what was prescribed by the Taylor rule (see Taylor (2009)). Bernanke (2010) ascribes this to the “jobless recovery” experienced at the time, but some may surmise that this aggressive monetary policy was one of the root causes of the recent credit crisis (see Rajan (2006), Bekaert, Hoerova, and Lo Duca (2013)). The recent credit crisis starting in 2007 is preceded by a passive monetary policy regime which, given the low inflation environment, implies that interest rates increased. In the beginning of the credit crunch, our model identifies a switch towards an (expansionary) discretionary monetary policy, whereas the probability of a systematic stabilizing policy also increases, leading to a sharp decline of interest rates.

### 4.3 Stability and Determinacy under Rational Expectations

We now compute the forward solution of the model to determine a fundamental solution consistent with the transversality condition, and examine determinacy under rational expectations. The forward solution has the form of equation (8) and the coefficient matrices  $\Omega$  and  $\Gamma$  are given by:

$$\Omega(s_t^{mp} = 1) = \begin{bmatrix} 0.884 & 0.067 & -0.198 \\ -0.061 & 0.391 & -0.424 \\ 0.272 & 0.102 & 0.610 \end{bmatrix}, \quad \Omega(s_t^{mp} = 2) = \begin{bmatrix} 1.184 & 0.093 & -0.626 \\ 0.480 & 0.444 & -1.161 \\ 0.186 & 0.062 & 0.583 \end{bmatrix}$$

$$\Gamma(s_t^{mp} = 1) = \begin{bmatrix} 1.537 & 0.206 & -0.238 \\ -0.106 & 1.204 & -0.510 \\ 0.474 & 0.312 & 0.732 \end{bmatrix}, \quad \Gamma(s_t^{mp} = 2) = \begin{bmatrix} 2.060 & 0.286 & -0.751 \\ 0.834 & 1.366 & -1.393 \\ 0.323 & 0.190 & 0.699 \end{bmatrix}$$

Note that the volatility regime variables do not affect these coefficient matrices.  $\Gamma(s_t^{mp})$  governs the initial responses of the structural shocks to the variables. For instance, inflation and the output gap fall following a contractionary monetary policy shock (see third column of each  $\Gamma$ ). In the case of a positive inflation shock, if the initial stance of monetary policy is active ( $s_t^{mp} = 1$ ), the output gap falls and inflation rises. However, when the policy is passive, the central bank raises the

nominal interest rate less than one for one, reducing the real interest rate. This actually raises the output gap as the  $(2, 1)$ -th component of  $\Gamma(s_t^{mp} = 2)$  is positive.

While the long run Taylor principle argument indicates that a passive monetary policy can be admissible as a determinate equilibrium, our estimated system may be indeterminate as the parameter  $\beta_2$  in the passive regime is 0.598, significantly less than 1. We employ the numerical search method of Farmer, Waggoner, and Zha (2011) and indeed, find that multiple stable solutions exist to our estimated model, leading to indeterminacy. This implies that monetary policy could have ensured determinacy, had it been less passive for our sample period. Hence, it is important to quantify the degree of passiveness admissible for determinacy. For this task, we resort to Cho (2013), who develops very tractable determinacy conditions for MSRE models (see Section 2.3, equation (11)).

We now analyze what combinations of policy parameters,  $\beta_1$  and  $\beta_2$ , satisfy these determinacy conditions, holding other parameters fixed. Figure 2 plots our determinacy and indeterminacy regions. Clearly, the policy stance in our MSRE model can be temporarily passive, and still yield a determinate equilibrium; however, it cannot be too passive. Indeed,  $r(\bar{D}_\Omega) = 0.775$  and  $r(D_F) = 1.25$  at our parameter estimates, implying that our equilibrium is outside the determinacy region. Recall that the passive policy stance prevailed in the pre-Volcker era and for more than half of the post-Volcker regime. Reflecting this fact, our estimate of  $\beta_2$  is low, namely 0.598, putting the model in the indeterminacy region. To ensure determinacy,  $\beta_2$  should be greater than 0.936. Several articles have identified spells of passive monetary policy before (Fernández-Villaverde, Guerrón-Quintana, and Rubio-Ramírez (2010), Bianchi (2013)) but our article is the first to characterize determinacy and show that recent passive policy stances result in an overall indeterminate MSRE equilibrium for the US economy. When varying other estimated parameter values, the determinacy region is not much affected, except when we vary  $\rho$ . When  $\rho$  becomes relatively large, determinacy requires both policy regimes to be active if one regime is too active relative to the other.

#### 4.4 Impulse responses

There are three independent structural shocks in the model (see  $\varepsilon_t$  in equation (19)). A nice feature of our model is that the impulse responses are regime-dependent, and should differ across regimes. Because agents are assumed to know the regime, we compute the impulse responses using an information set that incorporates both data and the regime; they follow from calculating  $E[X_{t+k}|I_t, s_t^{mp} = i]$ , for  $i = 1, 2$ . Appendix E describes a simple procedure to compute these impulse responses recursively. Note that this computation takes into account the expectations of agents regarding future switches in the monetary policy regime.

Figures 3 to 5 produce these regime dependent impulse responses of all three macro-variables

to one-standard deviation shocks, focusing on, respectively, AS, IS and monetary policy shocks. In each figure, there are three panels corresponding to the three macro-variables. We show 4 different impulse responses, depending on the monetary policy regime and the shock volatility regime. While the volatility regimes only affect the initial size of the shock, the relative magnitude of the impulse responses helps us interpret macroeconomic dynamics in different time periods.

Figure 3 focuses on AS shocks. This is of considerable interest as there is a lively debate on whether the stagflations of the seventies were partially policy driven. The figure shows that following an AS shock, inflation is highest in the high inflation shock volatility - passive monetary policy regime, as was observed in the 1970s, and lowest in the low inflation shock volatility - activist monetary policy regime, as observed from 1985 to 1993. It is especially activist monetary policy that contributes to a lower inflation response. Investigating output gap responses, a positive AS shock drives down the output gap in a protracted way under an activist monetary policy response, because the real interest rate increases. However, the output gap increases when monetary policy is accommodating as then the real interest rate decreases following a positive AS shock. However, after about 6-7 quarters, the output gap is lower under an accommodating regime than it is under an activist regime. The effect of AS shocks on nominal interest rates is also strikingly regime-dependent. Except for the initial periods, the accommodating regime yields higher nominal interest rate responses than the activist regime. This is because under accommodating monetary policy, it takes time for inflation to decrease - both through the direct effect of monetary policy and through expectations -, so that interest rates must be kept high for a long time. The regime-dependent responses therefore provide simultaneously an interesting interpretation of the historical record on the macroeconomic response to the negative aggregate supply shocks in the seventies and a counter-factual analysis. The accommodating policy regime implied (excessively) high interest rates, high inflation, and a substantial long term loss in output. The responses under an activist regime show that an aggressive Fed could have likely lowered the magnitude of the inflation response, reduced inflation volatility, kept interest rates overall lower and avoided the longer-term output loss, at the cost of a short-term loss over the first 5 quarters.

Figure 4 shows the responses to the IS shock. The inflation responses are similar across monetary policy regimes, but move over a wider range under the accommodating regime. In that regime, inflation rates move substantially below their mean during some periods, simultaneously with the interest rate undershooting its mean. The output gap responses are also quite similar across regimes. The similarity of the responses may have something to do with the fact that monetary policy reacts similarly to demand shocks across both regimes. While Panel C shows that the interest response to a demand shock is higher in the activist regime, the response differences are both in absolute and relative terms multiple times smaller than the responses to supply shocks, observed in Figure 3.

Finally, Figure 5 shows the responses to the monetary policy shock. Clearly, the activist monetary policy regime implies (much) more stable inflation and output dynamics than the passive regime. The macroeconomic volatility under the accommodating regime is especially dramatic when the interest rate shock is in the high volatility regime (recall that the interest rate shock volatility is multiple times higher in that case). A contractionary monetary policy shock lowers inflation and the output gap in both regimes, but, as the third panel shows, this is not only accommodated with less macroeconomic but also less interest rate volatility in the activist regime.

## 4.5 Macro-variability and its Sources

US economic history has witnessed profound changes in the volatility of macroeconomic variables over time, as evidenced by the literature on the Great Moderation. In the context of our model, this time variation in macroeconomic variability is driven by changing regimes in the variability of macroeconomic shocks (driven by  $s_t^\pi, s_t^y, s_t^i$ ) and regime dependent feedback parameters, which also depend on the monetary policy regime,  $s_t^{mp}$ . In this section, we derive the unconditional and regime-dependent variances of our macro variables, and provide different decompositions to shed light on the sources of macroeconomic variability.

### 4.5.1 A Variance Decomposition

The regime variable  $S_t$  contains 16 different regimes, as each of the four independent regimes,  $s_t^{mp}$ ,  $s_t^\pi$ ,  $s_t^y$  and  $s_t^i$  has two states. Appendix F shows in detail how to compute the unconditional variance as a sum of regime-dependent variances:

$$\text{Var}(X_t) = \sum_{i=1}^S \text{Var}(X_t | S_t = i) \cdot P_i \quad (20)$$

where  $P_i = \Pr(S_t = i)$  is the unconditional, ergodic regime probability, and  $S = 16$ . Appendix F also derives closed-form expressions for the regime-dependent variances. We then compute the contribution of a particular regime to the total variance as:

$$r_x(S_t = i) = \frac{\text{Var}(x_t | S_t = i) P_i}{\text{Var}(x_t)} \quad (21)$$

where  $x_t$  represents  $\pi_t$ ,  $y_t$  or  $i_t$ .

Table 3 reports these ratios together with the long run, ergodic distribution ( $P_i$ ). For instance, the regime combination of an active monetary policy and high shock volatility across all three equations contributes 1.24, 1.98 and 3.17% to the total variance of inflation, the output gap and the interest rate, respectively. The regimes contributing the most to the unconditional variance reflect



passive monetary policy, the high variability regime for inflation shocks and the low variability regime for output shocks. The latter is true because the low variability regime for output occurs more frequently than the high variability regime (69.81% versus 30.19% in fact), whereas the opposite is true for inflation shocks, where the high variability regime occurs 68.97% of the time and also for interest rate shocks where the high variability regime occurs 59.47% of the time.

The most noticeable result is that in all cases, the contribution to total variance of any variable is much smaller under the active monetary regime than it is under the passive regime. For instance, when the economy is in the high volatility regime for all shocks, the active regime contributes only 1.98% to the total variance of the output gap, whereas the passive regime contributes 14.31%, about 7.23 times more. Of course, the contribution could simply be low because the active regime has a much lower probability of occurring. In the high volatility regimes, the ergodic probability of the active regime is 3.23% while it is 9.16% under the passive regime, about three times higher. Therefore, even after controlling for differences in ergodic probabilities, the volatility of the output gap under the active regime is much smaller than that under the passive regime. This is generally true for all regime combinations and all the macro-variables.

To see this more explicitly, Table 4 shows variance ratios for the various regimes,  $Var(x_t|S_t = i)/Var(x_t)$ , that is the variance in that particular regime relative to the unconditional variance. Strikingly, the variance ratio for output and inflation variability in the active regime when all the shocks are in the high variability regime is lower than the variance ratio for the output and inflation variability in the passive regime when all the shocks are in the low variability regime. This suggests that the monetary policy regime has a rather important impact on macro-variability and perhaps an impact that exceeds the impact of the variability of macro shocks.

To compare the relative effect on variability of shocks versus policy, the last line shows the ratio of the variance in a regime where all macro shocks are in the high variability regime versus the variance of a regime where all the macro shocks are in the low variability regime. These ratios obviously depend on the macro variable and the policy regime, but their range is rather narrow varying between 2.30 and 2.87. To compare this to the effect of monetary policy, Table 5 shows the ratio of the passive versus active variances, controlling for the shock variability regimes. It is obvious that policy has a relatively larger effect on output and inflation variances than do macro shocks. In terms of variances, the passive monetary policy regime leads to variances of inflation and the output gap that are about two to five times as large as their variances in the active regime.

## 4.5.2 The Great Moderation

The above computations can also help us identify the start and the end of the Great Moderation. In our model, the Great Moderation is a period in which the time-varying variance is substantially below its unconditional counterpart. Recall that the variance can take on 16 different values as there

are 16 regimes. At each point of time, agents in the economy know the regime (and hence the variance), but we can only estimate the probabilities of different regimes occurring using the data. We therefore estimate the variance at each point of time as the sum of the regime-dependent variances weighted by their associated time-varying smoothed regime probabilities using full sample information. That is,

$$\widehat{Var}(X_t) = \sum_{i=1}^S Var[X_t|S_t = i]P[S_t = i|I_T] \quad (22)$$

If regime classification is perfect (that is, the smoothed probabilities are zero or 1), the summation simply selects one of the 16 regime-dependent variances. Because regime-classification in a system with 16 regimes is unlikely to be perfect, using smoothed regime probabilities to average the regime-dependent variances seems reasonable.

Figure 6 graphs the ratio of an estimate of the time-varying variance relative to the unconditional variance for inflation, the output gap and interest rates. Visually, the graph clearly identifies the Great Moderation lasting from the third quarter of 1980 to the third quarter of 2007, with inflation and output variability being substantially below the 1 line, often even being less than 50% of the unconditional variance. Do note that there are short episodes during the Great Moderation where inflation and particularly output variability briefly spike up.

Our previous computations suggest that policy played a rather important role in the Great Moderation. For example, it is striking that we identify the Great Moderation to start *before* the shock variabilities move to a lower variability regime. This is, of course, due to a switch from a passive to active monetary policy regime around 1980. To visualize the effect of policy on macro-variances, we run a counterfactual analysis. In Figure 7, we graph a volatility ratio, namely the standard deviation of the three macro variables, conditional on the monetary policy regime always being in the passive regime versus the actual time-varying volatility, that is, the square root of the variance computed in equation (22). When computing the counterfactual volatility, the underlying variance computation transfers mass from states where  $s_t^{mp} = 1$  to the corresponding state (and its variance) where  $s_t^{mp} = 2$ . Figure 8 does the opposite computation, it computes the volatility assuming the monetary policy regime is always activist, and graphs the ratio of the actual over the activist volatility.<sup>8</sup>

With these two graphs in hand, we can reinterpret the historical evolution of macro-volatility as generated by our model. In the seventies, macro-volatility was around twice as high as it could

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<sup>8</sup>Specifically, we define the counterfactual probability measure of permanently passive monetary policy regime as  $\hat{P}(S_t = i|Passive, I_T)$  where  $\hat{P}(s_t^{mp} = 1, j, k, l|I_T) = 0$  and  $\hat{P}(s_t^{mp} = 2, j, k, l|I_T) = P(s_t^{mp} = 1, j, k, l|I_T) + P(s_t^{mp} = 2, j, k, l|I_T)$  for all  $s_t^\pi = j, s_t^y = k, s_t^i = l, j, k, l = 1, 2$ . Using this probability measure, we can define the time-varying variance of the policy being always passive as  $\widehat{Var}[X_t|Passive]$ . The counterfactual activist probability measure and activist variance can also be defined analogously. Figure 7 and 8 depict respectively  $\frac{\sqrt{\widehat{Var}[X_t|Passive]}}{\sqrt{\widehat{Var}(X_t)}}$  and  $\frac{\sqrt{\widehat{Var}(X_t)}}{\sqrt{\widehat{Var}[X_t|Active]}}$ .

have been, had monetary policy been active (see Figure 8). From 1981 to 1993, active monetary policy managed to reduce macro-volatility substantially - it would have been 50% to 200% higher otherwise (Figure 7). The relatively subdued macro-variability after 1993 to around 2000 was due to low variability in the macro shocks, as monetary policy was passive. Of course, as we have argued before, the earlier aggressive policy stance may have helped anchor expectations during a rather mild macroeconomic climate. Taking our model literally, monetary policy could have further reduced macro-volatility by continuing to be aggressive. Because inflation was low at that time, an active monetary policy would have meant *lower* interest rates. The jump in counterfactual volatility around 2000 in Figure 7 is the more dramatic of the two graphs. In other words, if monetary policy had remained passive, macro-volatilities would have increased substantially. Bernanke's (2010) speech explicitly discusses this episode as the Federal Reserve reacting aggressively to a deflation scare, reducing the interest rate way below what a standard Taylor rule would predict. The period also witnessed a number of macroeconomic shocks that could have caused macro-volatility to increase and augmented recession risk, such as the events of September 11, 2001.

#### **4.6 Rational expectations versus survey expectations**

Our estimation imposes a parameter space that ensures the existence of a fundamental rational expectations equilibrium. What happens if this assumption is relaxed? Table 6 shows the results for the unconstrained estimation. In Table 1, the right-hand side panel also produces specification tests for this model. The model only performs marginally better than the constrained model. Moreover, the resulting estimates imply explosive dynamics for the RE model. Nevertheless, it is noteworthy that the parameter estimates are very similar to those obtained in the constrained estimation. The only significant difference is that  $\mu$ , the forward-looking parameter in the IS equation, is now significantly smaller, 0.331, relative to 0.675 before. This is similar to the values obtained by Fuhrer and Rudebusch (2004), in their systematic single equation estimation in a fixed regime context. Bekaert, Cho, and Moreno (2010) also estimate a lower value for  $\mu$ , namely 0.422, but this is coupled with a high estimate for the degree of forward-looking behavior in the AS equation ( $\delta$  in our model). As we have verified through simulation exercises, the combination of low  $\delta$  and low  $\mu$ —maintaining standard values for other parameters—implies the non-existence of a stable RE equilibrium, both in a fixed regime and in a multiple regime context. In economic terms, stable RE dynamics require AS and IS equations with a sufficient degree of forward looking behavior, such that shocks are rapidly absorbed.

Finally,  $\alpha$ , the parameter governing the law of motion for the survey-based expectations, is 0.410 in the unconstrained case, whereas it was 0.986 in the constrained estimation. This is an important difference. When we enforce a stable RE, RE appear indistinguishable from SBE, whereas

in the unconstrained estimation, SBE slowly adjust to RE, being heavily influenced by past expectations. In fact,  $\alpha$  is statistically indistinguishable from 0.5, implying that rational expectations and past survey-based expectations obtain similar relative weights in the expectations formation process. In other words, viewed through the lens of this macroeconomic model, survey expectations only slowly adjust to rational expectations, being heavily influenced by past expectations. This is consistent with Mankiw, Reis, and Wolfers (2004), who show that the adjustment of SBE to the macro environment is gradual. Conversely, the dependence on rational expectations is highly significant, implying that survey expectations likely convey much information, useful in estimating macroeconomic parameters and dynamics.

Figure 9 shows the regime probabilities for the unconstrained model, which should be compared to Figure 1 for the RE model. Focusing first on Panel B, the monetary policy regime identification, both for systematic and discretionary policy is very similar, qualitatively and quantitatively, to that in the constrained estimation. In Panel A, we observe some differences in terms of output shock regime identification. First, the high output volatility prevails from the beginning of the sample, whereas in the constrained estimation this regime appears more gradually. In addition, the Great Moderation in terms of output volatility shocks starts abruptly around 1986, which is a few years later than in the constrained estimation. Second, the low volatility output shock regime already ends in 2000, much earlier than in the constrained optimization. These differences can be easily understood examining the transition probabilities of the IS shock regime variable across estimations (see Tables 2 and 6). The unconstrained estimation shows much more persistence in the high variance regime and less persistence in the low volatility regime than the constrained estimation.

To sum up, when we relax the assumption of RE, we find an  $\alpha$  that is statistically different from 1, implying SBE that load heavily on past SBE. Nevertheless, the parameter estimates, model dynamics and regime identification are similar in this model to what they were in the RE equilibrium.

## 5 Conclusions

In this article, we identified macroeconomic regimes through the lens of a simple New-Keynesian model accommodating regime switches in macroeconomic shocks and systematic monetary policy. We demonstrate that monetary policy has witnessed several spells of activist policy, which have become more frequent post 1980. Nevertheless, we do not see a permanent switch from accommodating to activist policy around 1980, but rather occasional switches back and forth between the two regimes. One reason is that the data suggest an important and time-varying role for discretionary monetary policy. For example, the Volcker period is characterized by both activist systematic policy and discretionary active policy. We also document important changes in the variances

of output and volatility shocks. It is no surprise that we find strong evidence of a “shock variability moderation” occurring around 1985 for output, whereas for inflation the timing is somewhere between 1985 and 1990. What is new is that we find strong evidence of this volatility reduction having ended, for output at the onset of the recent economic crisis (more precisely in 2007), for inflation, earlier in 2005. The variability of shocks is not the only determinant of macro-variability however. Our model implies that the effect of monetary policy regimes on macro variability is relatively larger than the effect of the variability of shocks. When we investigate the time path of the overall variability of inflation and the output gap, we find that the Great Moderation starts around 1980 and ends in about 2007. During that period, a predominantly active monetary policy and low variability economic shocks combined to make output and inflation substantially less variable than unconditional averages would suggest.

Estimating a rational expectations New-Keynesian model with regime switches is difficult from a numerical perspective. Our innovation was to expand the information set with survey expectations on inflation and output growth. By formulating a simple law of motion for these expectations as a function of the true rational expectations, we could greatly simplify the likelihood construction. Constraining the parameter space to those parameters that yield a stable rational expectations equilibrium, we find survey expectations to be almost equivalent to rational expectations. However, when we relax these constraints, we find survey expectations to only gradually adjust to rational expectations and the parameters to be outside the rational expectations equilibrium space. Fortunately, the identification of regimes remains similar to that obtained in the rational expectations model, except that the Great Moderation in terms of output volatility ends much earlier (in 2000!) when identified from the unconstrained model.

There are two possible interpretations to these different estimation results. One possibility is that agents truly have rational expectations but that our New-Keynesian model is misspecified. Perhaps, we need a more intricate natural rate of output process or we must add investment equations as in Smets and Wouters (2007) to better fit the data. We did experiment with slightly more complex specifications (e.g. alternative characterizations of the monetary policy rule, alternative values for state-dependent transition probabilities, correlated regimes) within the confines of the stylized New-Keynesian model, finding little improvement in fit, and no noteworthy new results. Perhaps some of the parameters we now assume to be time-invariant may also be unstable. For example, a number of recent articles including Benati (2008), Hofmann, Peersman, and Straub (2012), and Liu, Waggoner, and Zha (2011)) have raised the possibility of an unstable AS equation, for instance because the degree of price and/or wage indexation changes through time.<sup>9</sup> An alternative possibility is that the assumption of rational expectations is too rigid, and we must build a model that

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<sup>9</sup>Some preliminary analysis did not reveal any evidence in favor of switches in  $\delta$ , the parameter governing the degree of forward looking behavior in the AS equation.

accommodates the presence of agents with not fully rational expectations. In any case, we hope this article stimulates the use of survey expectations in building and estimating macroeconomic models.

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Table 1: Specification Tests on Model Residuals

This table reports Monte-Carlo p-values for the different specification tests described in Appendix D. For both the Rational Expectations and Unconstrained Model, the univariate tests test for a zero mean, no second order autocorrelation, zero skewness, and no excess kurtosis in the standardized residuals of the output, inflation, and interest rate equations. The bottom panel reports Monte-Carlo p-values for a test of zero covariances of the factor shocks of one state variable with the factor shocks of the other two state variables, as well as a joint test that all covariances are equal to zero.

	<b>Rational Expectations Model</b>			<b>Unconstrained Model</b>		
<b>Univariate Tests</b>	Output	Inflation	Short Rate	Output	Inflation	Short Rate
<i>Mean Test</i>						
Zero mean	0.560	0.855	0.905	0.377	0.559	0.807
Autocorrelation	0.000	0.930	0.926	0.000	0.119	0.839
Joint	0.000	0.985	0.976	0.000	0.250	0.933
<i>Variance Test</i>						
Unit Variance	0.771	0.684	0.907	0.294	0.552	0.743
Autocorrelation	0.057	0.739	0.485	0.907	0.502	0.504
Joint	0.382	0.845	0.451	0.771	0.536	0.208
<i>Test on Higher Moments</i>						
Zero Skewness	0.221	0.033	0.281	0.365	0.450	0.433
Zero Excess Kurtosis	0.861	0.251	0.643	0.853	0.480	0.694
<b>Covariance Tests</b>						
Covar shocks with other	0.000	0.000	0.090	0.001	0.000	0.038
Joint	0.000			0.000		

Table 2: Parameter Estimates Rational Expectations Model

This table reports the estimation results of the Rational Expectations Model with independent regimes in respectively the monetary policy parameters  $\beta$  and  $\gamma$ , the volatility of inflation shocks  $\epsilon_{AS_t}$ , the volatility of output shocks  $\epsilon_{IS_t}$ , and the volatility of monetary policy shocks  $\epsilon_{MP_t}$ , as outlined in Section 2. The regime-switching variables are respectively denoted as  $s_t^{mp}$ ,  $s_t^\pi$ ,  $s_t^y$ , and  $s_t^i$ . Panel 1 reports the parameters of the AS and IS equation. Panel 2 reports the monetary policy parameters. Panel 3 shows the regime-switching volatilities of respectively the inflation shocks ( $\sigma_{AS}(s_t^\pi)$ ), the output shocks ( $\sigma_{IS}(s_t^y)$ ), and the interest rate shocks ( $\sigma_{MP}(s_t^i)$ ) (on a quarterly basis). Panel 4 reports the transition probabilities for the four independent regime-switching variables. Panel 5 reports the alpha parameter. Standard errors are reported between parentheses. For the regime-switching parameters, we also report p-values for a Wald test of equality across regimes between square brackets.

<i>1. Output Gap and Inflation Parameters</i>					
$\delta$	$\lambda$		$\mu$	$\phi$	
0.425	0.102		0.675	0.100	
(0.065)	(0.044)		(0.030)	-	
<i>2. Monetary Policy Parameters</i>					
$\rho$	$\beta(s_t^{mp} = 1)$	$\beta(s_t^{mp} = 2)$	$\gamma(s_t^{mp} = 1)$	$\gamma(s_t^{mp} = 2)$	
0.834	2.312	0.598	1.187	0.687	
(0.022)	(0.182)	(0.140)	(0.414)	(0.111)	
	[0.001]		[0.217]		
<i>3. Volatilities</i>					
$\sigma_{AS}(s_t^\pi = 1)$	$\sigma_{AS}(s_t^\pi = 2)$	$\sigma_{IS}(s_t^y = 1)$	$\sigma_{IS}(s_t^y = 2)$	$\sigma_{MP}(s_t^i = 1)$	$\sigma_{MP}(s_t^i = 2)$
0.334	0.162	0.142	0.072	0.249	0.041
(0.051)	(0.044)	(0.031)	(0.007)	(0.021)	(0.007)
	[0.000]		[0.000]		[0.000]
<i>4. Transition Probabilities</i>				<i>5. Alpha</i>	
$P_{11}^{mp}$	$P_{22}^{mp}$	$P_{11}^\pi$	$P_{22}^\pi$	$\alpha$	
0.878	0.957	0.991	0.980	0.986	
(0.108)	(0.036)	(0.012)	(0.028)	(0.020)	
$P_{11}^y$	$P_{22}^y$	$P(s_t^i)$	$Q(s_t^i)$		
0.963	0.984	0.893	0.843		
(0.047)	(0.015)	(0.045)	(0.057)		

Table 3: Variance Decompositions and Ergodic Distribution for all Regimes

This table reports the ergodic distribution and the variance decomposition results. In the first and sixth columns, A and P stand for ‘active’ and ‘passive’ monetary policy regimes. H and L stand for ‘high’ and ‘low’ volatility regimes for supply, demand and monetary policy shocks, respectively. The 2nd and 7th columns show the probability of each regime in the ergodic distribution, which measures the unconditional probability of each regime combination. The 3rd through 5th, and 8th through 10th columns report the ratio of the variance of each variable conditional on a regime combination to its total variance, in percent. That is, all columns add up to 100.

$(s_t^{mp}, s_t^\pi, s_t^y, s_t^i)$	$P_i$	$r_\pi$	$r_y$	$r_i$	$(s_t^{mp}, s_t^\pi, s_t^y, s_t^i)$	$P_i$	$r_\pi$	$r_y$	$r_i$
(A, H, H, H),	3.23	1.24	1.98	3.17	(P, H, H, H)	9.16	14.24	14.31	12.08
(A, H, H, L)	2.20	0.79	1.23	1.96	(P, H, H, L)	6.24	8.61	6.52	7.51
(A, H, L, H)	7.46	2.86	4.35	7.26	(P, H, L, H)	21.17	32.89	32.18	27.83
(A, H, L, L)	5.09	1.82	2.68	4.50	(P, H, L, L)	14.43	19.88	14.46	17.30
(A, L, H, H)	1.45	0.24	0.50	0.71	(P, L, H, H)	4.12	3.25	4.56	2.86
(A, L, H, L)	0.99	0.14	0.28	0.40	(P, L, H, L)	2.81	1.72	1.65	1.62
(A, L, L, H)	3.36	0.55	1.05	1.62	(P, L, L, H)	9.53	7.49	10.13	6.56
(A, L, L, L)	2.29	0.32	0.59	0.90	(P, L, L, L)	6.49	3.96	3.54	3.72

Table 4: Ratio of Regime-Dependent Variances and Unconditional Variances in each regime

This table shows the ratio of the variance of a given variable ( $\pi$ : inflation,  $y$ : output gap and  $i$ : interest rate) conditional on a given regime to the unconditional variance of that variable implied by the model. Each regime combines systematic monetary policy (A: active, P: passive) and regime shock size for the three shocks (H: high, L: low). In the last line, we divide the ratio of the all-high-shock regime by the all-low-shock regime, for both active and passive monetary policy regimes.

$(s_t^{mp}, s_t^\pi, s_t^y, s_t^i)$	$r_\pi$	$r_y$	$r_i$	$(s_t^{mp}, s_t^\pi, s_t^y, s_t^i)$	$r_\pi$	$r_y$	$r_i$
(A, H, H, H),	0.38	0.61	0.98	(P, H, H, H)	1.56	1.56	1.32
(A, H, H, L)	0.36	0.56	0.89	(P, H, H, L)	1.38	1.05	1.20
(A, H, L, H)	0.38	0.58	0.97	(P, H, L, H)	1.55	1.52	1.31
(A, H, L, L)	0.36	0.52	0.88	(P, H, L, L)	1.38	1.00	1.20
(A, L, H, H)	0.17	0.34	0.49	(P, L, H, H)	0.79	1.11	0.69
(A, L, H, L)	0.14	0.29	0.40	(P, L, H, L)	0.61	0.59	0.58
(A, L, L, H)	0.16	0.31	0.48	(P, L, L, H)	0.79	1.06	0.69
(A, L, L, L)	0.14	0.26	0.39	(P, L, L, L)	0.61	0.55	0.57
(H,H,H) vs (L,L,L)	2.76	2.39	2.49	(H,H,H) vs (L,L,L)	2.55	2.87	2.30

Table 5: Ratio of Passive and Active Regime-Dependent Variances

This table shows the ratio between the regime dependent variances conditional on the passive monetary policy regime and those conditional on the active monetary policy regime, for all variables.

$(s_t^{mp}, s_t^\pi, s_t^y, s_t^i)$	$r_\pi$	$r_y$	$r_i$
(P, H, H, H)/(A, H, H, H)	4.05	2.55	1.34
(P, H, H, L)/(A, H, H, L)	3.85	1.87	1.35
(P, H, L, H)/(A, H, L, H)	4.05	2.61	1.35
(P, H, L, L)/(A, H, L, L)	3.85	1.90	1.35
(P, L, H, H)/(A, L, H, H)	4.76	3.23	1.41
(P, L, H, L)/(A, L, H, L)	4.38	2.05	1.44
(P, L, L, H)/(A, L, L, H)	4.77	3.41	1.43
(P, L, L, L)/(A, L, L, L)	4.39	2.13	1.45

Table 6: Estimation Results for Unconstrained Regime-Switching Macro Model

This table reports the estimation results of the unrestricted New-Keynesian Model with independent regimes in respectively the monetary policy parameters  $\beta$  and  $\gamma$ , the volatility of inflation shocks  $\varepsilon_{AS,t}$ , the volatility of output shocks  $\varepsilon_{IS,t}$ , and the volatility of monetary policy shocks  $\varepsilon_{MP,t}$ , as outlined in Section 2. The regime-switching variables are respectively denoted as  $s_t^{mp}$ ,  $s_t^\pi$ ,  $s_t^y$ , and  $s_t^i$ . Panel 1 reports the parameters of the AS and IS equation. Panel 2 reports the monetary policy parameters. Panel 3 shows the regime-switching volatilities of respectively the inflation shocks ( $\sigma_{AS}(s_t^\pi)$ ), the output shocks ( $\sigma_{IS}(s_t^y)$ ), and the interest rate shocks ( $\sigma_{MP}(s_t^i)$ ) (on a quarterly basis). Panel 4 reports the transition probabilities for the four independent regime-switching variables. Panel 5 reports the alpha parameter. Standard errors are reported between parentheses. For the regime-switching parameters, we also report p-values for a Wald test of equality across regimes between square brackets.

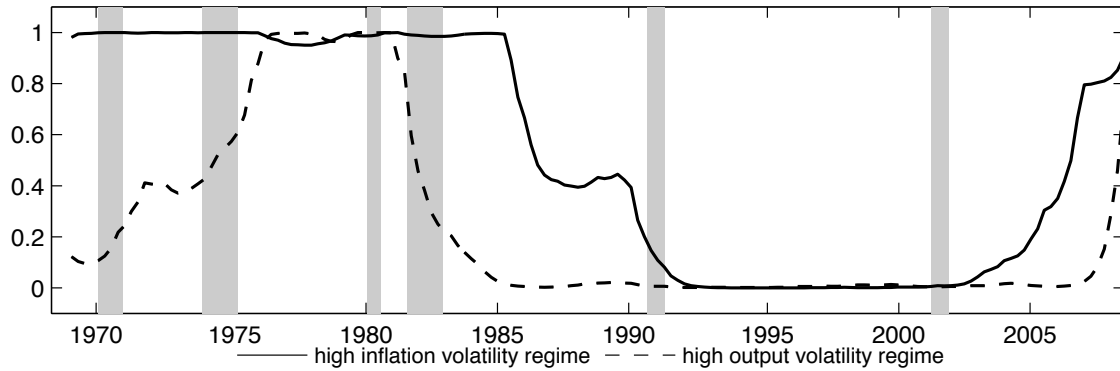
<i>1. Output Gap and Inflation Parameters</i>					
$\delta$	$\lambda$		$\mu$	$\phi$	
0.351	0.076		0.331	0.100	
(0.070)	(0.031)		(0.048)	-	
<i>2. Monetary Policy Parameters</i>					
$\rho$	$\beta(s_t^{mp} = 1)$	$\beta(s_t^{mp} = 2)$	$\gamma(s_t^{mp} = 1)$	$\gamma(s_t^{mp} = 2)$	
0.871	2.164	0.210	1.335	0.748	
(0.020)	(0.250)	(0.192)	(0.328)	(0.138)	
<i>3. Volatilities</i>					
$\sigma_{AS}(s_t^\pi = 1)$	$\sigma_{AS}(s_t^\pi = 2)$	$\sigma_{IS}(s_t^y = 1)$	$\sigma_{IS}(s_t^y = 2)$	$\sigma_{MP}(s_t^i = 1)$	$\sigma_{MP}(s_t^i = 2)$
0.316	0.137	0.092	0.050	0.253	0.038
(0.034)	(0.026)	(0.010)	(0.023)	(0.040)	(0.004)
<i>4. Transition Probabilities</i>				<i>5. Alpha</i>	
$P_{11}^{mp}$	$P_{22}^{mp}$	$P_{11}^\pi$	$P_{22}^\pi$	$\alpha$	
0.841	0.936	0.990	0.973	0.410	
(0.095)	(0.032)	(0.011)	(0.024)	(0.066)	
$P_{11}^y$	$P_{22}^y$	$P_{11}^i$	$P_{22}^i$		
0.973	0.959	0.891	0.837		
(0.080)	(0.116)	(0.057)	(0.061)		



Figure 1: Smoothed Regime Probabilities (Rational Expectations Model)

This figure shows the smoothed probabilities of the regimes in the general regime-switching New-Keynesian Macro model with four independent regime variables. Panel A shows the smoothed probabilities of respectively the high inflation shock volatility regime and the high output shock volatility regime. Panel B shows the smoothed probabilities of respectively the active monetary policy regime, and the high interest rate shock volatility regime. NBER recessions are shaded gray.

Panel A: High Inflation and Output Volatility Regimes



Panel B: Active Monetary Policy and High Interest Rate Volatility Regimes

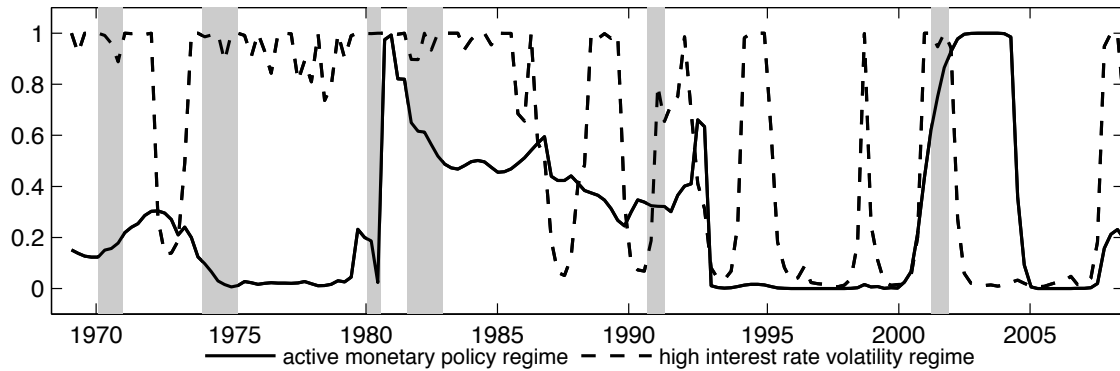


Figure 2: Determinacy and Indeterminacy Regions under Rational Expectations

This figure shows the determinacy/indeterminacy regions of our MS DSGE implied by different values of the two regime-dependent interest rate responses to expected inflation. The remaining parameters are set at their estimated values.

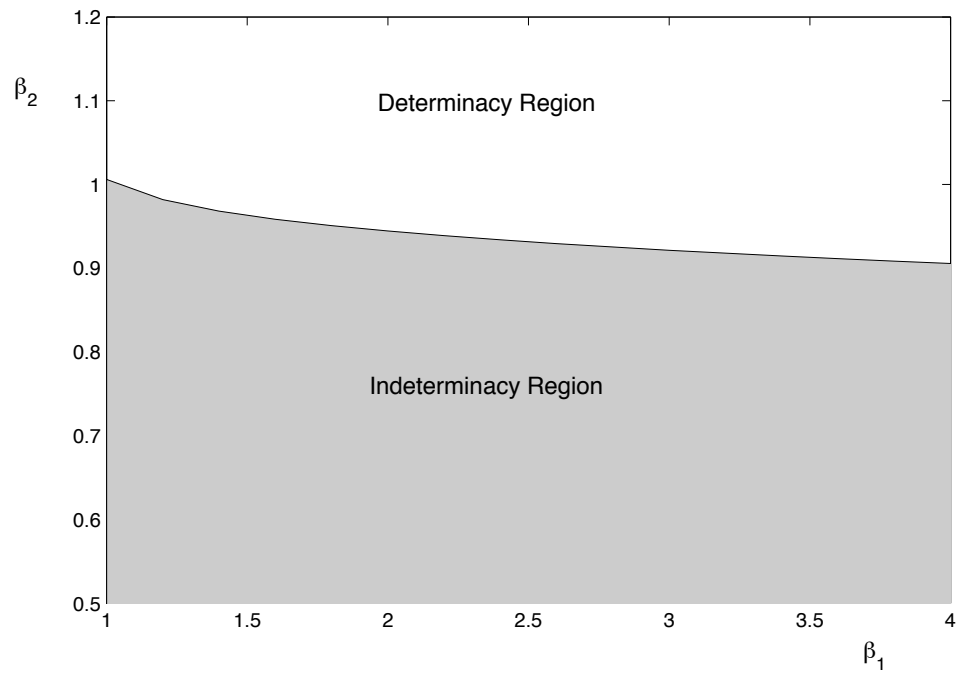


Figure 3: Impulse Responses to AS Shocks

This figure shows the impulse responses of inflation, the output gap and the short-term interest rate to the AS shock implied by our structural regime switching model. Each panel plots the responses of each variable dependent on a given monetary policy and shock regime.  $MP_1$  ( $MP_2$ ) represents the estimated stabilizing (accommodating) monetary policy reaction function, whereas  $\sigma_1$  ( $\sigma_2$ ) is the estimated high (low) shock.

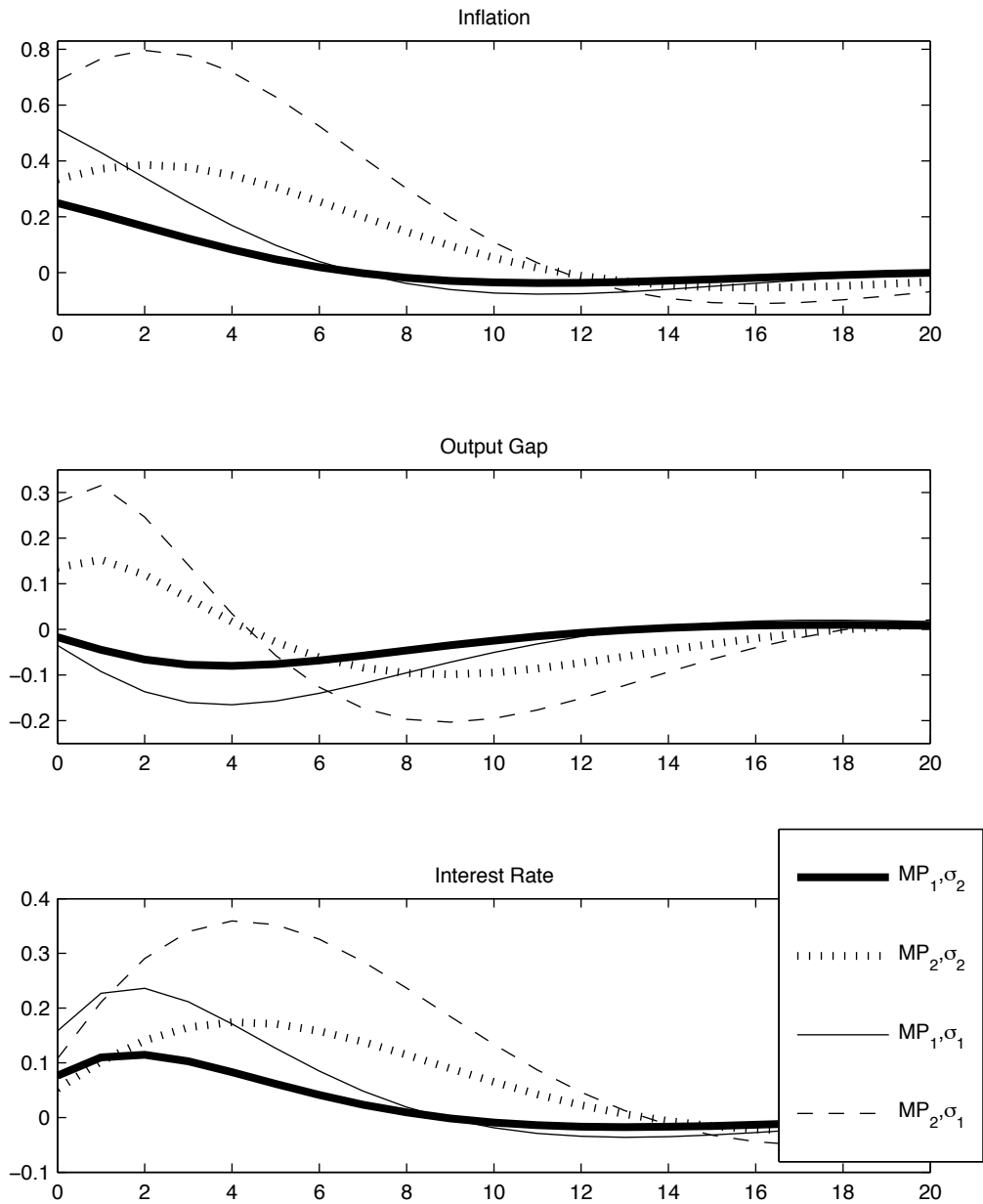


Figure 4: Impulse Responses to IS Shocks

This figure shows the impulse responses of inflation, the output gap and the short-term interest rate to the IS shock implied by our structural regime switching model. Each panel plots the responses of each variable dependent on a given monetary policy and shock regime.  $MP_1$  ( $MP_2$ ) represents the estimated stabilizing (accommodating) monetary policy reaction function, whereas  $\sigma_1$  ( $\sigma_2$ ) is the estimated high (low) shock.

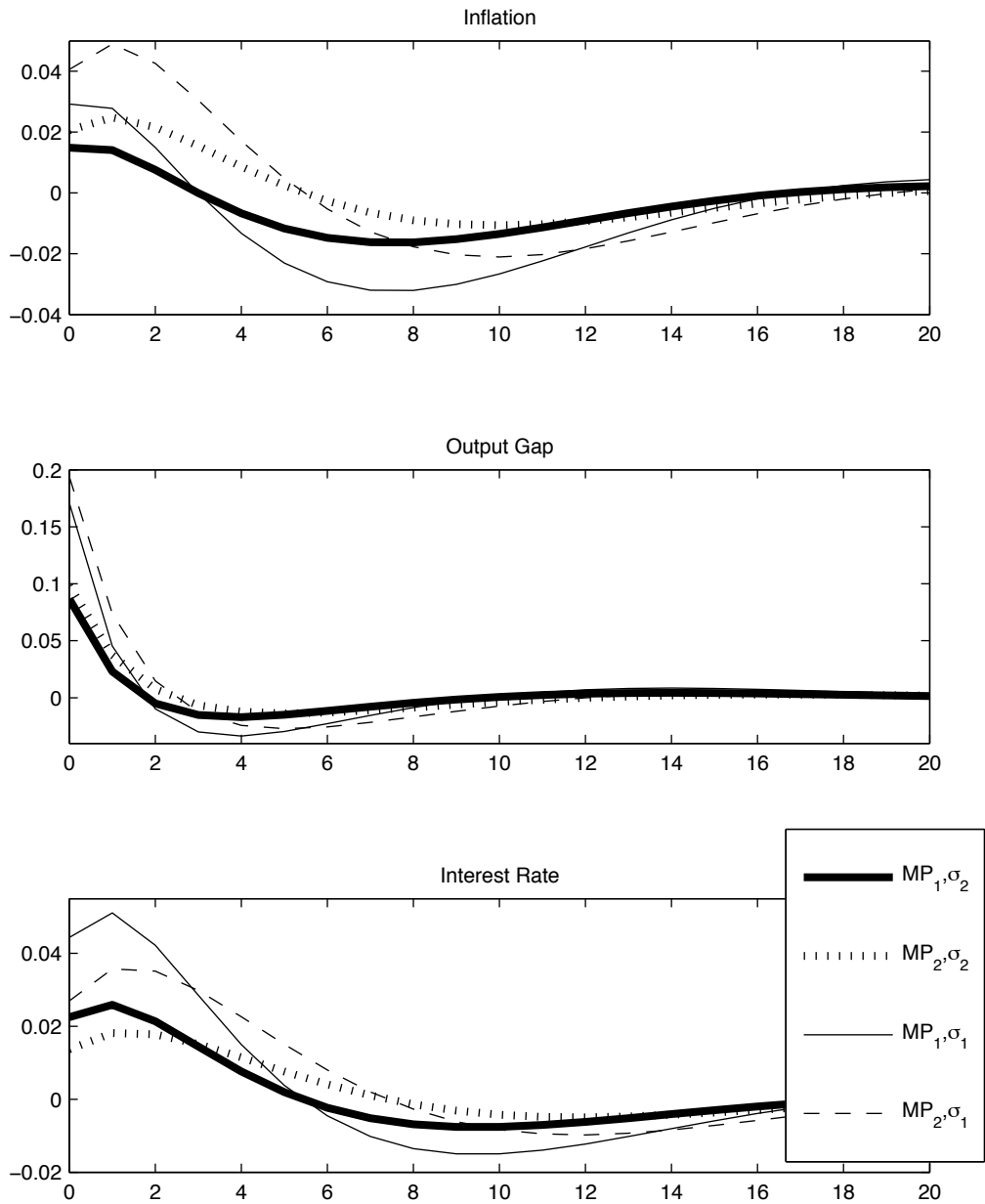


Figure 5: Impulse Responses to MP Shocks

This figure shows the impulse responses of inflation, the output gap and the short-term interest rate to the MP shock implied by our structural regime switching model. Each panel plots the responses of each variable dependent on a given monetary policy and shock regime.  $MP_1$  ( $MP_2$ ) represents the estimated stabilizing (accommodating) monetary policy reaction function, whereas  $\sigma_1$  ( $\sigma_2$ ) is the estimated high (low) shock.

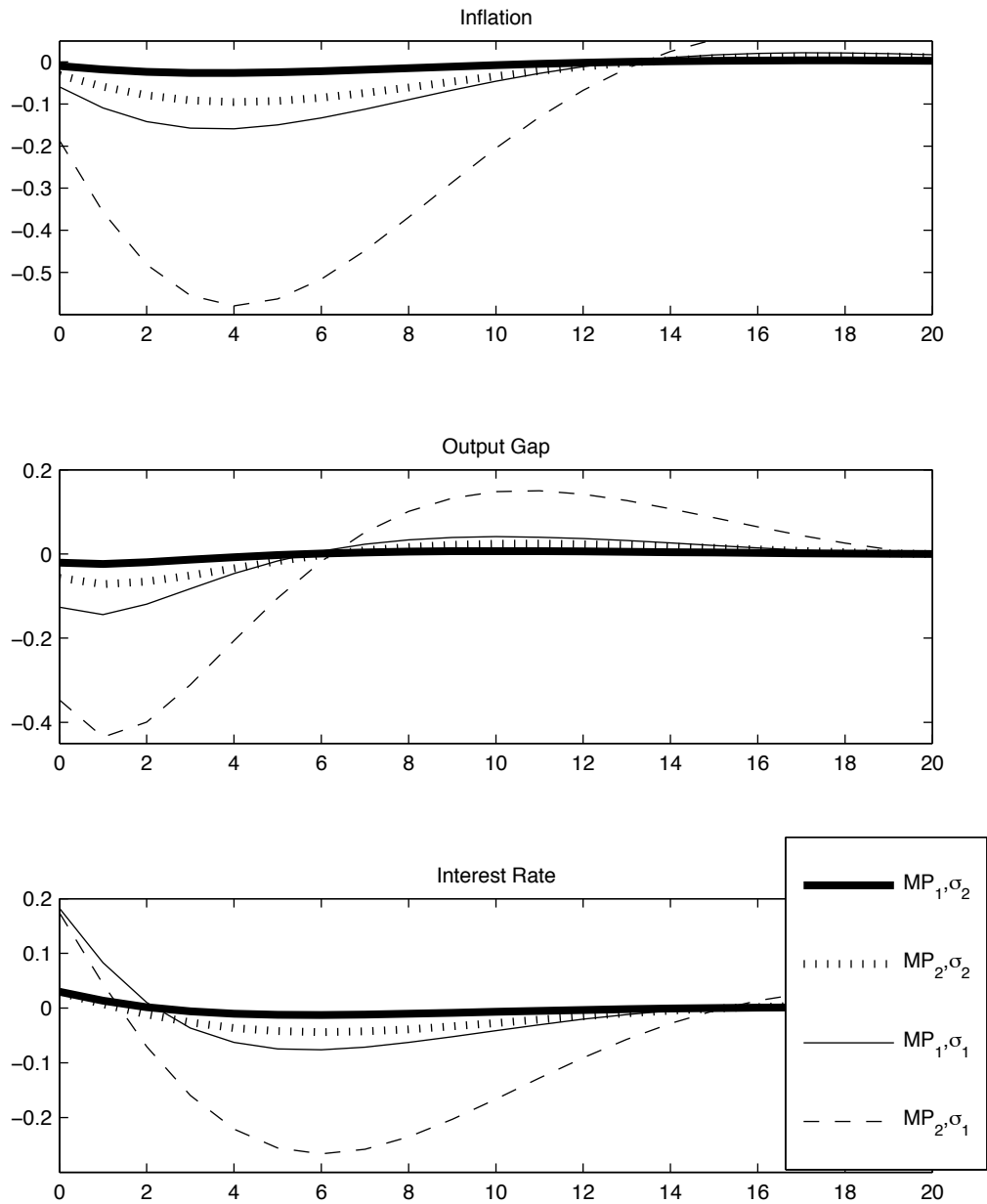


Figure 6: Time-Varying Variances relative to Unconditional Variances

This figure plots the ratio of the time-varying variance, computed using the smoothed regime probabilities as described in equation (19), to the unconditional variance for the inflation, output gap and interest rate series.

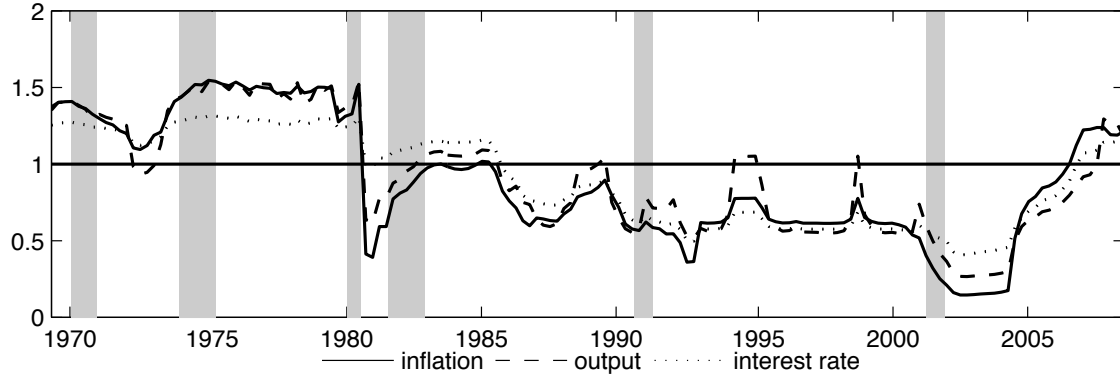


Figure 7: Contribution of Monetary Policy to the Volatility  $[\sqrt{\widehat{Var}[X_t|Passive]}/\sqrt{\widehat{Var}(X_t)}]$

This figure plots the volatility ratio between the counterfactual volatility of the three macro variables, conditional on the monetary policy regime always being in the passive regime, and their time-varying volatility (calculated using the smoothed probabilities).

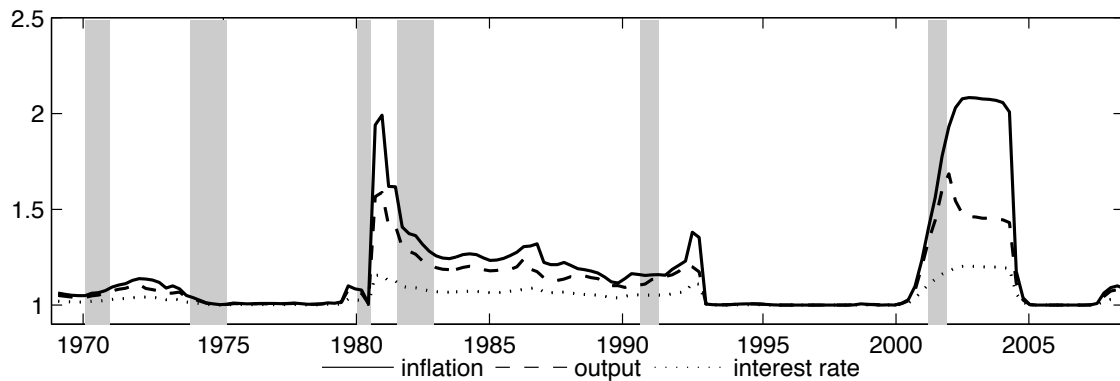


Figure 8: Contribution of Monetary Policy to the Volatility  $[\sqrt{\widehat{\text{Var}}(X_t)} / \sqrt{\widehat{\text{Var}}[X_t|Active]}]$

This figure plots the volatility ratio between the time-varying volatility (calculated using the smoothed probabilities) of the three macro variables and their counterfactual volatility, conditional on the monetary policy regime always being in the active regime.

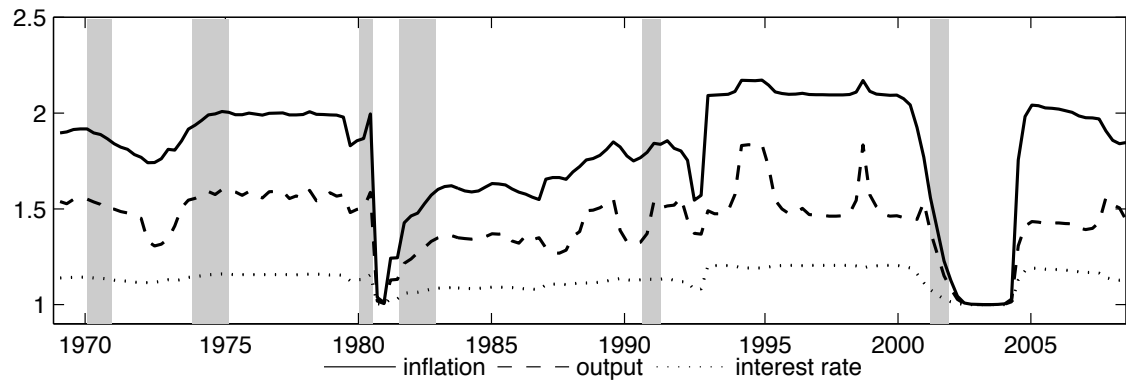
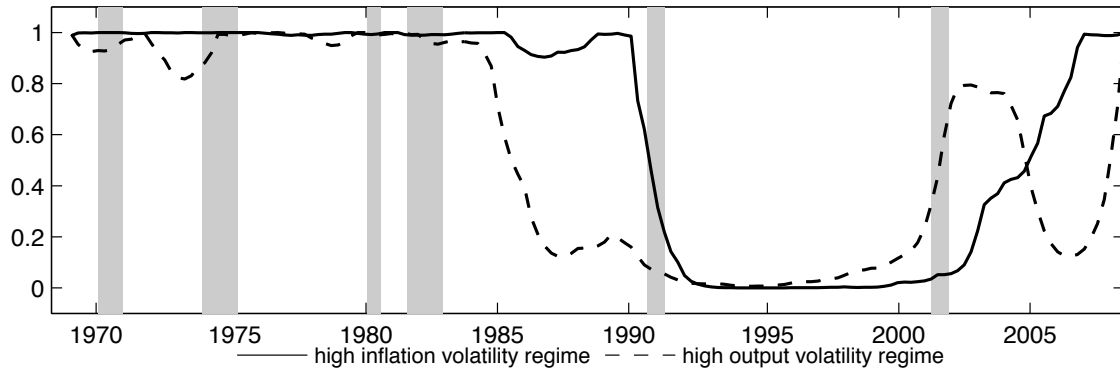


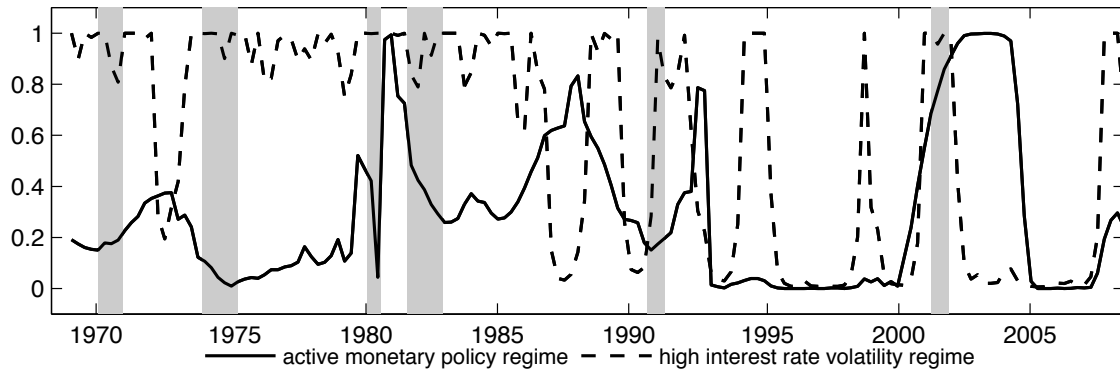
Figure 9: Smoothed Regime Probabilities (Unconstrained Model)

This figure shows the smoothed probabilities of the regimes in the unrestricted regime-switching New-Keynesian Macro model with four independent regime variables. Panel A shows the smoothed probabilities of respectively the high inflation shock volatility regime and the high output shock volatility regime. Panel B shows the smoothed probabilities of respectively the active monetary policy regime, and the high interest rate shock volatility regime. NBER recessions are shaded gray.

Panel A: High Inflation and Output Volatility Regimes



Panel B: Active Monetary Policy and High Interest Rate Volatility Regimes





# Appendix

## A Forward Solution and Determinacy

This appendix outlines the actual computation of the forward solution and sketches the technical details regarding the determinacy conditions. For expositional purposes, let us rewrite the model in equation (7) as:

$$X_t = \hat{A}(S_t)E_t X_{t+1} + \hat{B}(S_t)X_{t-1} + \hat{C}(S_t)\varepsilon_t, \quad \varepsilon_t \sim N(0, \Sigma(S_t)), \quad (\text{A-1})$$

where  $\hat{A}(S_t) = A(S_t)^{-1}B(S_t)$ ,  $\hat{B}(S_t) = A(S_t)^{-1}D$  and  $\hat{C}(S_t) = A(S_t)^{-1}$ . Recall the general form of the rational expectations solution in equations (8) and (9), which we reproduced here for convenience:

$$X_t = \Omega(S_t)X_{t-1} + \Gamma(S_t)\varepsilon_t + u_t, \quad (\text{A-2})$$

$$s.t. \quad u_t = F(S_t)E_t u_{t+1}. \quad (\text{A-3})$$

Now plug the general solution in (A-2) into the model in (A-1) and match the coefficient matrices in (A-2) and (A-3), to obtain the following restrictions on  $\Omega(S_t)$ ,  $\Gamma(S_t)$  and  $F(S_t)$ :

$$\Omega(S_t) = [I_n - \hat{A}(S_t)E_t \Omega(S_{t+1})]^{-1} \hat{B}(S_t), \quad (\text{A-4})$$

$$\Gamma(S_t) = [I_n - \hat{A}(S_t)E_t \Omega(S_{t+1})]^{-1} \hat{C}(S_t), \quad (\text{A-5})$$

$$F(S_t) = [I_n - \hat{A}(S_t)E_t \Omega(S_{t+1})]^{-1} \hat{A}(S_t). \quad (\text{A-6})$$

These conditions must be satisfied by any RE equilibrium. Following Cho (2013), one can solve the model (A-1) forward recursively. To do so, first define the coefficient matrices of the model as the initial values of the recursion with  $\Omega_1(S_t) = \hat{B}(S_t)$ ,  $\Gamma_1(S_t) = \hat{C}(S_t)$ ,  $M_1(S_t) = F_1(S_t) = \hat{A}(S_t)$ . Now lead the model one period forward, premultiply by  $\hat{A}(S_t)$  and take conditional expectations to obtain:

$$\hat{A}(S_t)E_t X_{t+1} = \hat{A}(S_t)E_t [M_1(S_{t+1})X_{t+2}] + \hat{A}(S_t)E_t \Omega_1(S_{t+1})X_t, \quad (\text{A-7})$$

which exploits the law of iterated expectations. Plugging this expectational term into (A-1) and rearranging, yields a relation between the current endogenous variables, expectations of the future endogenous variables at time  $t + 2$ , and the current state variables. By repeating these steps

recursively, we can derive the following forward representation of the model:

$$X_t = E_t[M_k(S_t, \dots, S_{t+k-1})X_{t+k}] + \Omega_k(S_t)X_{t-1} + \Gamma_k(S_t)\varepsilon_t, \quad (\text{A-8})$$

where the coefficient matrices are defined as:

$$\Omega_k(S_t) = [I_n - \hat{A}(S_t)E_t\Omega_{k-1}(S_{t+1})]^{-1}\hat{B}(S_t), \quad (\text{A-9})$$

$$\Gamma_k(S_t) = [I_n - \hat{A}(S_t)E_t\Omega_{k-1}(S_{t+1})]^{-1}\hat{C}(S_t), \quad (\text{A-10})$$

$$F_k(S_t) = [I_n - \hat{A}(S_t)E_t\Omega_{k-1}(S_{t+1})]^{-1}\hat{A}(S_t), \quad (\text{A-11})$$

$$M_k(S_t, \dots, S_{t+k-1}) = F_k(S_t)M_{k-1}(S_{t+1}, \dots, S_{t+k-1}). \quad (\text{A-12})$$

for  $k \geq 2$ . Let  $\Omega^*(S_t) = \lim_{k \rightarrow \infty} \Omega_k(S_t)$ ,  $\Gamma^*(S_t) = \lim_{k \rightarrow \infty} \Gamma_k(S_t)$  and  $F^*(S_t) = \lim_{k \rightarrow \infty} F_k(S_t)$ . Furthermore, consider a forward solution characterized by  $\lim_{k \rightarrow \infty} E_t[M_k(S_t, \dots, S_{t+k-1})X_{t+k}] = 0$ , a condition Cho and Moreno (2011) call the "no-bubble" condition:

$$X_t = \Omega^*(S_t)X_{t-1} + \Gamma^*(S_t)\varepsilon_t, \quad (\text{A-13})$$

If the coefficient matrices indeed converge as  $k$  tends to infinity, this solution in the limit fulfils the RE restrictions in (A-4) through (A-6). It is therefore a fundamental solution of the form (A-2), which is referred to as the forward solution. In the absence of regime-switching, this forward solution is stable if  $r(\Omega^*) < 1$ . In the case of regime-switching, the forward solution is mean-square stable if  $r(\bar{D}_{\Omega^*}) < 1$ , where  $\bar{D}_{\Omega^*}$  is defined in the main text.

Now we examine the determinacy conditions that there exists a unique mean-square stable fundamental solution and there is no mean-square stable sunspot component associated with that solution. The general solution associated with the forward solution is given by:

$$X_t = \Omega^*(S_t)X_{t-1} + \Gamma^*(S_t)\varepsilon_t + u_t, \quad (\text{A-14})$$

$$s.t. \ u_t = F^*(S_t)E_t u_{t+1} \quad (\text{A-15})$$

First, we need to obtain a condition under which there is no mean-square stable sunspot component satisfying (A-15). In the absence of regime-switching, it is well-known that the condition  $r(F^*) \leq 1$  ensures that any  $u_t$  process is unstable. For the regime-switching case, Cho (2013) derives an analogous result:  $r(D_{F^*}) \leq 1$  implies that there is no mean-square stable sunspot component, where  $D_{F^*}$  is a transition probability weighted matrix defined in the main text. This is a powerful result because a single matrix condition establishes non-existence of mean-square stable sunspot

component. For the actual derivation, we refer to Cho (2013).

Second, for the uniqueness of the stable fundamental solution, we need to introduce an important property of the forward solution. The requirement that the expectational term in equation (A-8) converges to zeros in the limit can be thought of as a transversality condition. This expectational term can be recursively evaluated with a fundamental solution  $X_t = \Omega(S_t)X_{t-1} + \Gamma(S_t)\varepsilon_t$  as  $L_k(S_t)X_t$  where  $L_k(S_t) = F_k(S_t)E_t[L_{k-1}(S_{t+1})\Omega(S_{t+1})]$  and  $L_0(S_{t+1}) = I_n$ . In the limiting case of (A-8),  $L(S_t) = \lim_{k \rightarrow \infty} L_k(S_t) = 0_{n \times 1}$  if evaluated with the forward solution by construction. If  $L(S_t)$  is evaluated with any other fundamental solution, then  $L(S_t) \neq 0_{n \times 1}$  in (A-8) because otherwise, it is a contradiction to the supposition that the solution differs from the forward solution. That is, for all other solutions,

$$L(S_t) = F^*(S_t)E_t[L(S_{t+1})\Omega(S_{t+1})] \neq 0_{n \times 1}. \quad (\text{A-16})$$

In fact, Cho (2013) shows that  $r(D_{F^*}) \leq 1$  implies  $r(\bar{D}_\Omega) \geq 1$  for all  $\Omega$ , that is, all other solutions are mean-square unstable. This is another important result because we do not need to solve for all other fundamental solutions, the number of which is unknown. Therefore, mean-square stability of the forward solution  $r(\bar{D}_{\Omega^*}) < 1$  together with  $r(D_{F^*}) \leq 1$  establishes determinacy.

## B Identifying the Parameter Space of Existence of Fundamental REEs

We characterize the parameter configuration for which a RE equilibrium exists numerically through a combination of grid search and randomized parameter choices. The parameters of the standard deviations and the transition probabilities do not matter for the existence of the fundamental REE. Hence, it is sufficient to consider the parameters in the AS, IS and MP equations, plus the transition probabilities of the monetary policy regimes:  $\delta$ ,  $\lambda$ ,  $\mu$ ,  $\phi$ ,  $\rho$ ,  $\beta_1$ ,  $\beta_2$ ,  $\gamma_1$ ,  $\gamma_2$ ,  $P_{11}^{mp}$  and  $P_{22}^{mp}$  with a restriction  $\beta_1 > \beta_2$  where  $\beta_i = \beta(s_t^{mp} = i)$ ,  $\gamma_i = \gamma(s_t^{mp} = i)$  for  $i = 1, 2$ , and  $P_{ij}^{mp} = \Pr(s_{t+1}^{mp} = j | s_t^{mp} = i)$  for  $i, j = 1$  and  $2$ . This table describes the parameter ranges we consider:

Parameter	Range	Parameter	Range
$\delta$	$(0, 1]$	$\lambda$	$(0, \infty)$
$\mu$	$(0, 1]$	$\phi$	$(0, \infty)$
$\rho$	$(0, 1)$		
$\beta_1$	$[1, \infty)$	$\beta_2$	$(0, \infty)$
$\gamma_1$	$(0, \infty)$	$\gamma_2$	$(0, \infty)$
$P_{11}^{mp}$	$(0, 1)$	$P_{22}^{mp}$	$(0, 1)$

Let  $\theta$  indicate a particular parameter vector and  $\Theta$  the parameter space specified in the above table. We decompose  $\Theta$  into the following disjoint subspaces  $\Theta(E)$  and  $\Theta(NE)$ .  $\Theta(E)$  is the space over which a fundamental REE exists and it has finite first moments, and  $\Theta(NE) = \Theta \setminus \Theta(E)$ . We refer to Cho (2013) for further details. Let  $\Theta(B(E))$  be the outer boundary of  $\Theta(E)$ . So the ultimate goal is to identify  $\Theta(B(E))$ .

An initial crude grid search and randomization procedure over the whole parameter space reveals that the parameters in  $\Theta(B(E))$  are inter-related in a complicated fashion and  $\Theta(B(E))$  is non-convex. Therefore, identifying and characterizing  $\Theta(B(E))$ , an 11-dimensional contour set, is a daunting task. Nevertheless, the initial procedure showed that  $\delta$ ,  $\mu$ ,  $\beta_1$  and  $\beta_2$  are the most critical parameters determining the existence of REEs. Therefore, we decompose  $\Theta$  into two subspaces  $\Theta_1$  and  $\Theta_2$  where  $\theta_1 = (\delta, \mu, \beta_1, \beta_2) \in \Theta_1$  and  $\theta_2 = (\lambda, \phi, \rho, \gamma_1, \gamma_2, P_{11}^{mp}, P_{22}^{mp}) \in \Theta_2$ . Then, for a given  $\theta_2$ , we grid-search over  $\Theta_1$  to identify  $\Theta(B(E))$ . It turns out that such a set is locally convex in  $\theta_1$  for a given  $\theta_2$ . Then we vary  $\theta_2$  to assess whether  $\Theta(B(E))$  is altered. Let's illustrate this procedure with an example:

Step 1: Fix  $\bar{\theta}_2$  at (0.05, 0.05, 0.5, 0.1, 0.1, 0.75, 0.5).

Step 2: Choose 10 values for each one of  $\theta_1 = (\delta, \mu, \beta_1, \beta_2)$ . That is, we check the existence of the forward solution for 10,000 sets of parameters, finding 8,755 sets for which the forward solution exists. We define the set  $\Theta_1(E|\bar{\theta}_2) = \{\theta_1 \mid \text{The forward solution exists at } (\theta_1, \bar{\theta}_2)\}$  and the boundary of this set is  $\Theta_1(B(E)|\bar{\theta}_2)$ .

Step 3: We change one parameter value in  $\theta_2$  and follow steps 1 and 2 to see how the set  $\Theta_1(B(E)|\bar{\theta}_2)$  changes.

It is not possible to tabulate the set  $\Theta(B(E))$  in a systematic way. Instead, we verbally describe our main findings. First, holding other parameters fixed, combinations of high  $\beta_1 (> 1)$  and low  $\beta_2 (< 1)$  form the boundary,  $\Theta(B(E))$ . Hence, a REE can exist in a model where monetary policy is temporarily passive ( $\beta_2 < 1$ ). Second,  $\Theta(B(E))$  is convex (locally) over  $\theta_1$ . Third, combinations of high  $\delta (> 0.5)$  and low  $\mu (< 0.5)$ , or vice versa, lie on the boundary. The forward solution does not exist for alternative private sector values (low  $\delta (< 0.5)$  and low  $\mu (< 0.5)$ ). Fourth, the parameter space  $\Theta(B(E))$  is convex over  $P_{11}^{mp}$  and  $P_{22}^{mp}$ , but not convex over  $(\lambda, \phi, \rho, \gamma_1, \gamma_2)$  in  $\theta_2$ . In particular, we are able to derive a lower boundary for  $P_{11}^{mp}$  (that is, parameters higher than the boundary are always in  $\Theta(E)$ ), and an upper boundary for  $P_{22}^{mp}$ .

In general, it is very difficult to compactly describe the parameter space for which the forward solution exists. The boundary  $\Theta(B(E))$  is not convex over all the parameters and the parameters are very interrelated. Nevertheless, our experiments here help us restrict the parameter space for the estimation procedure.

## C Data Appendix

Our dataset consists of economic state variables for the US. Our sample period is from the fourth quarter of 1968 to the second quarter of 2008 for a total of 159 observations. The state variables are seasonally adjusted and expressed in percentages at a quarterly basis. Below we give details on the exact data sources used and on the way the series are constructed:

1. **Output Gap ( $y$ ):** The output measure is real Gross Domestic Product (GDP), from the Bureau of Economic Analysis. The gap is computed as the percentage difference between output and its quadratic trend. The output gap is divided by four to express it at a quarterly basis.
2. **Expected Output Gap ( $y^f$ ):** The expected output gap is constructed as follows :

$$\begin{aligned} E_t [y_{t+1}] &= E_t \left[ \frac{g_t}{g_t} \left( \frac{g_{t+1}}{tr_{t+1}} - 1 \right) \right] \\ &= g_t \frac{E_t \left[ \frac{g_{t+1}}{g_t} \right]}{tr_{t+1}} - 1 \end{aligned}$$

with

$$\begin{aligned} g_t &= \text{level of real GDP at time } t \\ tr_t &= \text{(quadratic) trend value of real GDP at time } t. \end{aligned}$$

We use survey-based expectations of real GDP (level) for the current and next quarter to measure  $E_t \left[ \frac{g_{t+1}}{g_t} \right]$ . The source is the Survey of Professional Forecasters (SPF).

3. **Inflation ( $\pi$ ):** Percentage difference in the Gross Domestic Product Chain-type Price Index, from the U.S. Department of Commerce: Bureau of Economic Analysis.
4. **Expected Inflation ( $\pi^f$ ):** Median survey response of expected growth in the GDP deflator over the next quarter, from the Survey of Professional Forecasters (SPF).
5. **Nominal Risk-free Rate ( $i$ ):** 3-Month Treasury Bill: Secondary Market Rate. The source is Federal Reserve. The rate is divided by four to express it at a quarterly basis.

The Table below reports summary statistics for the different state variables:

<b>Panel A: Summary Statistics</b>							
	<i>Mean</i>	<i>Median</i>	<i>Maximum</i>	<i>Minimum</i>	<i>Std. Dev.</i>	<i>Skewness</i>	<i>Kurtosis</i>
$y_t$	0.0083	0.0417	1.3360	-1.5094	0.5496	-0.0302	2.9290
$y_t^f$	-0.0074	0.0748	1.3521	-1.5046	0.5654	-0.0368	2.9736
$\pi_t$	1.0022	0.7879	3.1137	0.1494	0.6164	1.1872	3.9637
$\pi_t^f$	0.9609	0.8091	2.3212	0.3234	0.5029	1.0852	3.2649
$i_t$	1.4657	1.3550	3.7550	0.2200	0.7096	0.8295	4.1004
<b>Panel B: Correlations</b>							
	$y_t$	$y_t^f$	$\pi_t$	$\pi_t^f$	$i_t$		
$y_t$	1.00	0.98	0.18	0.06	0.09		
$y_t^f$	0.98	1.00	0.14	-0.01	0.01		
$\pi_t$	0.18	0.14	1.00	0.84	0.57		
$\pi_t^f$	0.06	-0.01	0.84	1.00	0.77		
$i_t$	0.09	0.01	0.57	0.77	1.00		

## D Specification Tests

The reduced form of the model (see Section 2.4) is given by:

$$X_t = \Omega_1(S_t)X_{t-1}^f + \Omega_2(S_t)X_t^f + \Omega_3(S_t)X_{t-1} + \Gamma(S_t)\varepsilon_t \quad \varepsilon_t \sim N(0, \Sigma(S_t)), \quad (\text{D-1})$$

with  $\Sigma$  being the regime-dependent diagonal variance-covariance matrix for the shocks contained in  $\varepsilon_t$ . Recall that there are 16 different regimes. Let the smoothed regime probability for regime  $i$ , with a slight abuse of notation, be denoted by  $P_{i,t}$ . That is,  $P_{i,t} = P[S_t = i | I_t]$ . Because regimes are unobserved, the residuals are essentially unobservable to the econometrician. We therefore use the econometrician's best estimate for the residuals given full sample information and the smoothed probabilities, that is, using equation (19):

$$\bar{\varepsilon}_t = \sum_{i=1}^S (\Gamma(S_t = i))^{-1} \left[ X_t - \Omega_1(S_t)X_{t-1}^f - \Omega_2(S_t)X_t^f - \Omega_3(S_t)X_{t-1} \right] P_{i,t}$$

We denote the regime-dependent variance covariance matrix for these residual, again with a slight abuse of notation by  $V_t$ , that is:

$$V_t = \sum_{i=1}^S \Sigma(S_t = i) P_{i,t}$$

We perform our different tests on the standardized residuals  $z_t = V_t^{-\frac{1}{2}} \bar{\varepsilon}_t$ . We test for a zero mean

and no second-order correlation by testing whether or not  $b_1, b_2$ , and  $b_3$  are zero in:

$$E [z_t] - b_1 = 0 \quad (\text{D-2})$$

$$E [(z_t - b_1)(z_{t-1} - b_1)] - b_2 = 0 \quad (\text{D-3})$$

$$E [(z_t - b_1)(z_{t-2} - b_1)] - b_3 = 0 \quad (\text{D-4})$$

Define  $\hat{z}_t = (z_t - b_1)^2 - 1$ . We test for a well-specified variance by testing whether or not  $b_4, b_5$ , and  $b_6$  are equal to zero in:

$$E [\hat{z}_t] - b_4 = 0$$

$$E [\hat{z}_t \hat{z}_{t-1}] - b_5 = 0$$

$$E [\hat{z}_t \hat{z}_{t-2}] - b_6 = 0$$

We test for excess skewness and kurtosis by testing whether or not  $b_7$  and  $b_8$  are equal to zero in:

$$E [(z_t - b_1)^3] - b_7 = 0 \quad (\text{D-5})$$

$$E [(z_t - b_1)^4 - 3] - b_8 = 0 \quad (\text{D-6})$$

We estimate  $b_1$  to  $b_8$  using GMM with a Newey and West (1987) weighting matrix with number of lags equal to 5. The tests for zero mean, unit variance, zero skewness, and zero excess kurtosis follow a  $\chi^2(1)$  distribution, the tests for second order autocorrelation a  $\chi^2(2)$  distribution. The joint mean and variance tests follow a  $\chi^2(3)$  distribution. We also perform a small sample analysis of the test statistics. For each series, we use the estimated parameters from the model to simulate a time-series of similar length as our sample. For 500 of such simulated time-series, we calculate the test statistics, and use the resulting distribution to derive empirical probability values.

To investigate whether our model adequately captures the covariance between the factor shocks, we test whether the following conditions hold:

$$E [z_{l,t} z_{j,t}] = 0, \text{ for } l, j \in \{y_t, \pi_t, i_t\}; l \neq j.$$

We test for each of the 3 variables whether its shocks have a zero covariance with the two other shocks. This test follows a  $\chi^2$  distribution with 2 degrees of freedom. In addition, we report the joint test for the covariances between all factor shocks which follows a  $\chi^2$  distribution with 3 degrees of freedom. Our  $p$ -values results from a small sample analysis of the test statistics, analogous to that performed for the univariate tests.

## E Impulse Response Analysis

Recall from equation (8) and the discussion in Appendix A that the forward solution, a fundamental RE equilibrium to our model, can be characterized as follows:

$$X_t = \Omega^*(S_t)X_{t-1} + \Gamma^*(S_t)\varepsilon_t \quad \varepsilon_t \sim N(0, \Sigma(S_t)).$$

Conditional on time  $t$  information including  $S_t$ , the one-step ahead prediction of  $X_{t+1}$  is given by  $E_t X_{t+1} = F(S_t, 1)X_t$  where  $F(S_t, 1) = E[\Omega^*(S_{t+1})|S_t]$ . The  $k$ -step ahead prediction of  $X_t$  is then, computed recursively as  $E_t X_{t+k} = F(S_t, k)X_t$  where  $F(S_t, 0) = I_3$  and

$$F(S_t, k) = E[F(S_{t+1}, k-1)\Omega^*(S_{t+1})|S_t],$$

for  $k \geq 1$ . The initial value of  $X_t$  is given by  $\Gamma^*(S_t)\varepsilon_t$ . Therefore, the impulse responses of  $X_{t+k}$  to the initial innovation at time  $t$  conditional on  $S_t$  are expressed as:

$$IR(S_t, k) = F(S_t, k)\Gamma^*(S_t)\varepsilon_t, \tag{E-1}$$

which is just a function of the current state  $S_t$ . Note that the volatility regimes simply determine the initial size of a given shock and do not affect the impulse response dynamics. Therefore, the relevant regime variable in  $F(S_t, k)\Gamma^*(S_t)$  is  $S_t = s_t^{mp}$ . For instance, in the case of a supply shock and the initial volatility regime being 1, we can set  $\varepsilon_t = (\sigma_{AS}(s_t^\pi = 1) \ 0 \ 0)'$ . Therefore, the impulse-response analysis only needs to consider regime-switching in the monetary policy stance. For each shock, there are two impulse responses starting from the initial monetary policy regime, depending on the volatility regime of the shock.

## F Computing the Unconditional Variance and its Decomposition conditional on Regimes

In this appendix, we show how to compute the unconditional and regime-dependent variances of  $X_t$ , implied by the rational expectations solution of the model. For expositional purposes, we rewrite equation (8) as follows:

$$X_t = \Omega(S_t)X_{t-1} + V(S_t)u_t \quad u_t \sim N(0, I_3), \tag{F-1}$$

where  $V(S_t) = \Gamma(S_t)\Sigma^{1/2}(S_t)$ . To compute the variance of  $X_t$ , we first define the regime variable  $S_t$ , the corresponding transition probability matrix and its ergodic probabilities. Recall that the regime



variable  $S_t$  comprises 4 different regime variables  $S_t = (s_t^{mp}, s_t^\pi, s_t^y, s_t^i)$ , each potentially taking on 2 states. Thus the variable  $S_t$  has 16 different states, indexed in the following way:

$S_t$	$s_t^{mp}$	$s_t^\pi$	$s_t^y$	$s_t^i$	$S_t$	$s_t^{mp}$	$s_t^\pi$	$s_t^y$	$s_t^i$
1	(A	H	H	H)	9	(P	H	H	H)
2	(A	H	H	L)	10	(P	H	H	L)
3	(A	H	L	H)	11	(P	H	L	H)
4	(A	H	L	L)	12	(P	H	L	L)
5	(A	L	H	H)	13	(P	L	H	H)
6	(A	L	H	L)	14	(P	L	H	L)
7	(A	L	L	H)	15	(P	L	L	H)
8	(A	L	L	L)	16	(P	L	L	L)

where A and P stand for ‘active’ and ‘passive’ monetary policy regimes.; H and L stand for ‘high’ and ‘low’ volatility regimes for supply, demand and monetary policy shocks, respectively.

Because the 4 regime variables are assumed independent, the transition probabilities for  $S_t$  are easily computed from the underlying transition probabilities for the 4 separate regime probabilities. Let  $S_t = i$  correspond to  $(s_t^{mp}, s_t^\pi, s_t^y, s_t^i) = (i_1, i_2, i_3, i_4)$  where  $i_1 = A$  or  $P$ ,  $i_2, i_3, i_4 = H$  or  $L$ . The state  $S_t = j$  is defined analogously. Then, the transition probability of switching from  $i$  to  $j$ ,  $P_{ij} \equiv \Pr(S_t = j | S_{t-1} = i)$  for  $i, j = 1, \dots, 16$  is given by  $P_{ij} = P_{i_1, j_1}^{mp} \times P_{i_2, j_2}^\pi \times P_{i_3, j_3}^y \times P_{i_4, j_4}^i$ . If the regime-switching model is ergodic, the unconditional probabilities, denoted by  $P_i = \Pr(S_t = i)$ , satisfy

$$\sum_{i=1}^S P_{ij} P_i = P_j, \quad \sum_{i=1}^S P_i = 1$$

where  $S = 16$  and  $i, j = 1, 2, \dots, S$ .

Now multiply with  $X_t'$  on both sides of (F-1) and take expectations conditional on the state  $S_t = i$ :

$$E[X_t X_t' | S_t = i] = \Omega(S_t = i) E[X_{t-1} X_{t-1}' | S_t = i] \Omega'(S_t = i) + V(S_t = i) V'(S_t = i) \quad (\text{F-2})$$

as  $X_{t-1}$  and  $u_t$  are independent. Since there is no drift term in our model,  $E[X_t X_t' | S_t = i]$  is the variance of  $X_t$  conditional on  $S_t = i$ ,  $\text{Var}(X_t | S_t = i)$ . Following Ang, Bekaert, and Wei (2008),

$E[X_{t-1}X'_{t-1}|S_t = i]$  can be written as:

$$\begin{aligned}
E[X_{t-1}X'_{t-1}|S_t = i] &= \sum_{j=1}^S E[X_{t-1}X'_{t-1}|S_{t-1} = j] \Pr(S_{t-1} = j|S_t = i) \\
&= \sum_{j=1}^S E[X_{t-1}X'_{t-1}|S_{t-1} = j] \frac{\Pr(S_{t-1} = j)}{\Pr(S_t = i)} \Pr(S_t = i|S_{t-1} = j) \\
&= \sum_{j=1}^S E[X_{t-1}X'_{t-1}|S_{t-1} = j] \frac{P_j}{P_i} P_{ji}
\end{aligned}$$

Plugging this expression into (F-2), we have

$$E[X_t X'_t | S_t = i] = \Omega(i) \left( \sum_{j=1}^S E[X_{t-1} X'_{t-1} | S_{t-1} = j] \frac{P_j}{P_i} P_{ji} \right) \Omega'(i) + V(i) V'(i) \quad (\text{F-3})$$

where  $\Omega(i) = \Omega(S_t = i)$  and  $V(i) = V(S_t = i)$ .

In order to obtain a closed form expression for  $E[X_t X'_t | S_t = i]$ , we define  $v^x$  and  $v$  as follows:

$$v^x = \begin{bmatrix} \text{vec}(E[X_t X'_t | S_t = 1]) \\ \text{vec}(E[X_t X'_t | S_t = 2]) \\ \dots \\ \text{vec}(E[X_t X'_t | S_t = 16]) \end{bmatrix}, \quad v = \begin{bmatrix} \text{vec}(V(1)V'(1)) \\ \text{vec}(V(2)V'(2)) \\ \dots \\ \text{vec}(V(16)V'(16)) \end{bmatrix}$$

Then, equation (F-2) for all  $i = 1, \dots, S$  can be expressed as:

$$v^x = \Sigma^\Omega v^x + v, \quad (\text{F-4})$$

where  $(i, j)$ -th element of the matrix  $\Sigma^\Omega$  is given by:

$$\Sigma_{ij}^\Omega = \left[ \frac{P_j}{P_i} P_{ji} \Omega(i) \otimes \Omega(i) \right].$$

Therefore,  $v^x = (I_{n^2S} - \Sigma^\Omega)^{-1} v$  where  $n = 3$ . By reshaping  $v^x$  back into a matrix form, we have the formula for  $E[X_t X'_t | S_t = i]$  for all  $i = 1, 2, \dots, S$ . Finally,  $\text{Var}(X_t) = E(X_t X'_t)$  can be obtained as:

$$\begin{aligned}
\text{Var}(X_t) &= E(X_t X'_t) = E(E[X_t X'_t | S_t]) \\
&= \sum_{i=1}^S E[X_t X'_t | S_t = i] \cdot P_i \\
&= \sum_{i=1}^S \text{Var}[X_t | S_t = i] \cdot P_i.
\end{aligned}$$