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The Regime-Dependent Evolution of Credibility: A Fresh Look at Hong Kong's Linked Exchange Rate System

Boris Blagov
Hamburg University
Department of Economics
boris.blagov@wiso.uni-hamburg.de

Michael Funke
Hamburg University
Department of Economics
and
CESifo Munich
michael.funke@uni-hamburg.de

Abstract

An estimated Markov-switching DSGE modelling framework that allows for parameter shifts across regimes is employed to test the hypothesis of regime-dependent credibility of Hong Kong's linked exchange rate system. The model distinguishes two regimes with respect to the time-series properties of the risk premium. Regime-dependent impulse responses to macroeconomic shocks reveal substantial differences in spreads. These findings contribute to efforts at modelling exchange rate regime credibility as a non-linear process with two distinct regimes.

Keywords: Markov-switching DSGE models, exchange rate regime credibility, Hong Kong.
JEL-Classification: E32, F41, C51, C52

1 Introduction

Recent years have seen a resurgence of interest in exchange rate regimes. In the aftermath of the 1997–1998 Asian crisis and the global recession of 2008–2009, “crisis prevention” came to be viewed as a key criterion in choosing an exchange rate regime. With the partial collapse of the European exchange rate mechanism in September 1992, the notion that corner solutions such as free floats and super-strict pegs were preferable to intermediate regimes became widespread. The thinking was that they were less crisis-prone in the context of today’s huge and volatile financial markets on the assumption that investors will otherwise overwhelm intermediate regimes like band systems sooner or later. Put more bluntly, the options for exchange rate regimes were assumed to have hollowed out to the point where the only choices left to policymakers were whether to let exchange rates float or fix them permanently via a currency board or monetary union.

Consistent with this bipolar view, Hong Kong’s currency board system appears to be textbook corner solution. To pre-empt the weakening of confidence during the Sino-Anglo dispute on the return of Hong Kong’s sovereignty to China after 1997, the Hong Kong government adopted a currency board or linked exchange rate system on 17 October 1983, a.k.a the “Black Saturday Crisis.” Under this system, the money supply in Hong Kong was fully backed up by US dollar (USD) and the HK dollar (HKD) effectively fixed at rate of USD/HKD 7.8. Any one of the three note-issuing banks in this system wishing to print HKD notes would have to surrender an equivalent amount of USD (at the official rate) to the Hong Kong Monetary Authority (HKMA) in exchange for “Certificates of Indebtedness” that entitled the note-issuing bank to print a corresponding amount of HKD. Conversely, note-issuing banks could use their Certificates of Indebtedness in HKD to redeem an equivalent amount of USD from the HKMA. A distinctive feature of the system up to May 2005 was that no strong-side boundary existed, meaning the currency board system was asymmetric. In May 2005, however, the HKMA bit the bullet on appreciation and introduced a symmetric target zone with a HKD/USD band of 7.75 to 7.85.

A common argument for placing restraints on a currency board system is that it confers credibility in the spheres of exchange rate and monetary policy by relinquishing the devaluation option.[1] This is not always true, of course. One can point to numerous historical episodes where currency boards fail to enhance the credibility of the monetary authority. This is because the government retains its right to abandon the scheme and renege on its institutional commitments. In other words, political uncertainty about the preferences of current and future governments can erode credibility.[2] Thus, we ask how much credibility do policymakers gain by implementing

[1] It is worth adding that that currency boards have been found to perform better than soft pegs in terms of economic growth. A growing body of macroeconomic evidence suggests volatility is detrimental to economic growth, especially when financial opportunities are limited. See, for example, Aghion and Howitt (2009), pp. 329–339.

[2] Establishing a reputation is a potential solution to the credibility problem. In repeated game models upon which this argument is typically based, however, a reputational equilibrium where governments refrain from opportunism because of the long-term costs involved is just one of many possible outcomes.

a currency board? Our chosen case here is motivated by the fact that, among the countries that have adopted currency boards, Hong Kong is probably the one with the largest and most developed financial sector, and the highest capital mobility. Hence, Hong Kong's "iron" linked exchange rate system and money market seems an ideal starting point in the search for insights into our understanding of modern currency board systems.[3]

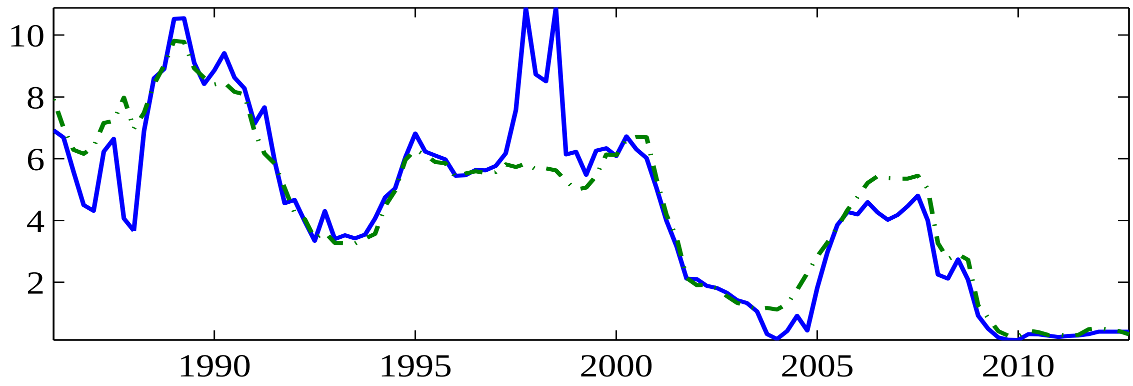

Figure 1: HKD HIBOR (—) and USD LIBOR (-.-), annualized 3-month interest rates.

Figure 1 shows the HK currency board tends to align domestic rates with US interest rates and the forward premia shed light into the time-varying credibility of Hong Kong's currency board system. With the exception of some small in 1991 and a short-lived increase of HK interest rates during the Mexican crisis in 1995, interest rate spreads relative to the US were close to zero for most of the 1989–1997 period. The Asian crisis in 1997 and its associated turbulence, however, altered this pattern dramatically. The HKD faced speculative pressure and capital outflows as HKD forward rates depreciated. The strategy of market participants was to bid up Hong Kong's interbank rate to benefit from short positions in the futures market. During this acute episode of loss of credibility, interest rate differentials surged. In 1998, they began a slow return to near-zero levels in 2000.

In contrast, financial markets in Hong Kong stayed remarkably calm during the SARS (severe acute respiratory syndrome) outbreak. If anything, confidence in the linked exchange rate system strengthened. Mirroring this, the interest rate differential between the HKD and the USD remained negligible. Moreover, the global financial crisis of 2008–2009 raised no doubts as to the credibility of Hong Kong's linked exchange rate system – the validity of the arrangement was never called into question! These sharp differences in spread movements between the Asian crisis and the global recession are quite striking given the extreme limits on Hong Kong's policy instruments.

The profile in Figure 1 suggests what it might take to call the credibility of Hong Kong's exchange rate system into question. We illustrate this by first drawing on financial market information captured by the behaviour of interest rates in the US and Hong Kong. We then develop a full-fledged DSGE model with Markov-switching to identify and interpret time-varying

[3] Oliva et al. (2001) present a signalling model to consider the choice between a currency board and a traditional peg. The model shows that the currency board's effectiveness and welfare effects hinge on its credibility.

credibility more precisely.[4] The main appeal of the structural approach is that it allows for a direct economic interpretation of observed movements in the data and fully exploits economic priors.[5]

The remainder of the article is as follows. Section 2 lays out the model. Section 3 deals with solving and estimating the Markov-switching DSGE specification. Section 4 discusses the data used in the estimation stage. Section 5 considers priors and posterior parameters. Section 6 presents the implications of the model. Section 7 discusses the robustness of the results. Section 8 concludes.

2 The Model

We use a Markov-switching DSGE Model to study the credibility of Honk Kong's linked exchange rate system. The model is based on the seminal works of Justiniano and Preston (2010) and Monacelli (2005) in combination with a fixed exchange rate and a Markov-switching component in the volatility of the interest rate risk premium.[6] In this section, we present briefly the key equations of the log-linearized system.

Consumers choose the optimal amount of consumption following the usual Euler equation with habit formation. h denotes the habit parameter, σ is the risk-aversion/inverse of the elasticity of substitution and ϑ_t is a preference shock that follows an AR(1) process.

$$c_t - hc_{t-1} = (E_t\{c_{t+1}\} - hc_t) + \frac{1-h}{\sigma}(E_t\{\pi_{t+1}\} - i_t) + \frac{1-h}{\sigma}(1-\rho_\vartheta)\vartheta_t \tag{1}$$

$$\vartheta_t = \rho_\vartheta \vartheta_{t-1} + \varepsilon_t^\vartheta \quad \text{with} \quad \varepsilon_t^\vartheta \sim N(0, \sigma_\vartheta^2). \tag{2}$$

As is standard in the small open economy (SOE) literature, consumption c_t is a bundle of domestic and foreign items. π_t stands for inflation, a weighted average of domestic and foreign goods where openness of the economy α is used as a weight. i_t is the nominal interest rate. Prices are set a lá Calvo in a hybrid manner, i.e firms are forward looking but have a degree of indexation. Inflation of domestic goods $\pi_{H,t}$ and of imports $\pi_{F,t}$ evolve according to

[4] The profile in Figure 1 suggests what it might take to call the credibility of Hong Kong's exchange rate system into question. We illustrate this by first drawing on financial market information captured by the behaviour of interest rates in the US and Hong Kong. We then develop a full-fledged DSGE model with Markov-switching to identify and interpret time-varying credibility more precisely.

[5] In recent years, the popularity of DSGE models with tight theoretical restrictions has gained ground. The trick is to make a model that closely approximates reality. The dominant pre-recession 2008–2009 DSGE paradigm viewed financial factors and/or credibility issues largely as a sideshow. The rapidly growing DSGE literature now seeks to remedy these known weaknesses, so the value of this line of remains contested. See, for example, Caballero (2010) and http://blogs.ft.com/maverecon/2009/03/the-unfortunate-uselessness-of-most-state-of-the-art-academic-monetary-economics/#axzz2V9NYYBKo.

[6] Liu and Mumtaz (2011) provide an extension of Justiniano and Preston (2010) to a Markov-switching DSGE framework from a floating exchange rate perspective.

$$(1 + \beta\delta_H)\pi_{H,t} = \beta E_t\{\pi_{H,t+1}\} + \delta_H \pi_{H,t-1} + \lambda_H mc_t \tag{3}$$

$$(1 + \beta\delta_F)\pi_{F,t} = \beta E_t\{\pi_{F,t+1}\} + \delta_F \pi_{F,t-1} + \lambda_F \psi_t + \mu_{F,t}. \tag{4}$$

where δ is the indexation parameter, β the discount factor and λ a function of the Calvo parameter. Domestic inflation is governed by the standard marginal costs driven by technology, while inflation of imported goods is subject to a cost-push shock $\mu_{F,t}$ and a function of ψ_t. This variable represents the deviations from the law of one price. It allows for a discrepancy between the prices of imported goods at the country of origin and the economy at hand. In particular, it relaxes the potentially tight link in the model between the real exchange rate (q_t) and the terms of trade (v_t) noted in Galí and Monacelli (2005)

$$\psi_t = q_t - (1 - \alpha)v_t. \tag{5}$$

Technology a_t and the import cost push shock $\mu_{F,t}$ are modelled as latent variables with normally distributed innovations to obtain

$$a_t = \rho_a a_{t-1} + \varepsilon_t^a \quad \text{with} \quad \varepsilon_t^a \sim N(0, \sigma_a^2) \tag{6}$$

$$\mu_{F,t} = \rho_\mu \mu_{F,t-1} + \varepsilon_t^{\mu_F} \quad \text{with} \quad \varepsilon_t^{\mu_F} \sim N(0, \sigma_{\mu_F}^2). \tag{7}$$

Exchange rate dynamics in SOE models are typically derived through the uncovered interest rate (UIP) parity and closed most often by a Taylor rule. Since Hong Kong has a currency board, we close the model by introducing a pegged exchange rate in accordance with Schmitt-Grohé and Uribe (2003) and Galí and Monacelli (2005) such that

$$\triangle e_t = 0. \tag{8}$$

Substituting through the UIP, we derive an important relationship between domestic and foreign interest rates. Namely, domestic rates i_t, are a function of foreign rates, i_t^*, and two additional components – an exogenous and an endogenous component such that

$$i_t = i_t^* - \chi d_t - \phi_t. \tag{9}$$

The d_t term represents the net foreign asset position. In an open economy, the agents may either borrow from domestic markets or tap into international markets. Following Benigno (2001), the borrowing rate from abroad depends on both the world interest rate i_t^* and the net foreign asset position. If the country is indebted, it has to pay a premium over the market interest. The term can be interpreted as change in the current account, i.e.

$$d_t = y_t - c_t - \alpha(q_t + \alpha v_t) + \frac{1}{\beta} d_{t-1}. \tag{10}$$

The last term in equation (9) is an exogenous AR(1) process that can be interpreted as a risk premium component

$$\phi_t = \rho_\phi \phi_{t-1} + \varepsilon_t^\phi \quad \text{with} \quad \varepsilon_t^\phi \sim N(0, \sigma_\phi^2(s_t)). \tag{11}$$

In order to capture the credibility of the Hong Kong's linked system, we allow for time variation of the variance of the shock by introducing regimes into the risk premium component.[7] $\sigma_\phi^2(s_t)$ is modelled as a regime dependent variable through a Markov-switching process with a probability matrix P.

$$P = \begin{bmatrix} p_{11} & p_{12} \\ p_{21} & p_{22} \end{bmatrix},$$

where p_{ij}, the $\{i, j\}$ element of the matrix, is the transition probability from state i to state j, $Prob(s_{t+1} = j | s_t = i)$. Thus, past states can recur over time. The proposed argument is that lower credibility of the system should lead to a risk premium and higher volatility of the interest rates.[8] By construction, we impose the rather strong assumption that the economy may only fall into specific (and a finite number of) regimes. This requires further motivation.

From the technical point of view (even if only two states might seem restrictive), this seems an obvious starting point. If the data do not support two distinctive cases, i.e. if the estimated coefficients overlap, additional regimes are inappropriate. Moreover, the introduction of more cases greatly increases the computational burden at the estimation stage.[9]

Two theories have been offered to explain regime switches in the risk premium. The first relates to the concept of sunspot shocks to agents' expectations. Here, sunspot shocks cause multiple equilibria (a low-risk premium equilibrium if rational agents are not worried about sunspot shocks, and a high-risk premium equilibrium if agents believe such shocks to be a bad). Thus, if the markets believe for some reason a currency crisis is underway, it happens. Jeanne and Masson (2000) propose an empirical test of sunspot-driven multiple equilibria in the currency crisis context. They prove that the effects of sunspot shocks are absorbed by discrete jumps in the intercept of a regression of the currency devaluation probability on fundamental variables. Therefore, a Markov regime-switching test can be used as a test for sunspot equilibria

[7] Engel (1994) and Engel and Hamilton (1990)have modelled exchange rates alternating between appreciation and depreciation regimes in a Markovian fashion. Their approach has a modicum of success in capturing the nonlinearity and regime shifts of the underlying time series and in forecasting. Marsh (2000), in contrast, shows that the Markov-switching modelling approach offers sound in-sample fit but usually fails to deliver a superior out-of-sample forecast due to parameter instability over time.

[8] This argument has been proposed and tested in Genberg and Hui (2011), who use an econometric analysis instead of building a structural model.

[9] The main pitfall is "the curse of dimensionality." A three-state Markov-switching model requires the identification of six coefficients in the probability matrix. The fact that the last column is a linear combination of the other two poses a significant problem at the estimation stage when the posterior mode is maximized. Furthermore, the Hessian at the posterior mode grows by almost three hundred elements (from 729 to 1024). These have to be estimated numerically, which greatly increases the margin for error. Moreover, the posterior distribution of a multi-state model may be highly non-Gaussian, which complicates the exploration for the Markov-Chain Monte Carlo procedure [see Sims et al. (2008)].

Another theory for regime-switching uses the "animal spirits" concept of De Grauwe (2010) and De Grauwe and Kaltwasser (2012). Here, boundedly rational and imperfectly informed agents use heuristics to make decisions in the foreign exchange market. Again, agents' psychological movements are self-fulfilling as waves of optimism and pessimism lead to fluctuations of the exchange rate even if the underlying fundamentals are unaltered by an exogenous shock. The theory of animal spirits shaping exchange rates is also consistent with a two-state regime-switching model. In theory reduced exchange rate volatility might translate into higher interest-rate volatility.[10] Hence, with a closed exchange rate channel modelling the dynamics of the interest rate in more detail is of particular interest. With the final equation (11) describing these dynamics, we are ready to solve the model.

First, we collect all of the equations into a state-space representation with the vector X consisting of all endogenous variables and Z containing the exogenous processes. The model is driven by seven exogenous variables: the four shocks of the home economy and three processes that describe the foreign economy (output, inflation and the interest rate, all modelled as latent variables).

$$B_1(s_t)X_t = E_t\{A_1(s_t, s_{t+1})X_{t+1}\} + B_2(s_t)X_{t-1} + C_1(s_t)Z_t \tag{12}$$

$$Z_t = R(s_t)Z_{t-1} + \epsilon_t \quad \text{with} \quad \epsilon_t \sim N(0, \Sigma^2(s_t)). \tag{13}$$

Here the matrices B_1, A_1, B_2, C_1 and R are time invariant. The only state-dependant matrix is $\Sigma^2(s_t)$. In the next section, we discusses how to solve and estimate (12) with actual data.

3 Solution and Estimation

The introduction of Markov Switching to DSGE models is a relatively new research area. There is yet no established way to solve and approximate these models. Several solution methods have been proposed by Davig and Leeper (2007), Farmer et al. (2011), Cho (2011) and Foerster et al. (2013). Notably, all revolve around the idea of a Minimal-State Variable Solution introduced by McCallum (1983), but explore different avenues to finding a solution, as well as how to check for uniqueness and determinacy of the system. Davig and Leeper (2007) use the notion of bounded shocks, while the last three follow the concept of Mean Square Stability (MSS). As MS-DSGE models may have more than one stable solution, each method must offer a way to choose among several solutions. Under unbounded shocks (as in this paper), the only references that provide a rationale for this problem are Cho (2011) and Farmer et al. (2011), who introduce the concept of a "no-bubble condition." We use this concept here as it is intuitively based around forward-solving the state-space system and assuming that only the non-explosive solution is economically relevant. This assumption is strengthened by a proof

[10] See Genberg and Hui (2011), p.186.

that if multiple solutions exist, only one fulfils this criterion.[11]

Following this method, the solution of (12) may be represented by two systems of equations, one for each state, i.e.

$$X_t = \Omega^*(s_t)X_{t-1} + \Gamma^*(s_t)Z_t. \tag{14}$$

This may be combined with a measurement equation for likelihood based estimation. Since the standard Kalman filter is not operable here, we may use Kim's filter as laid out in Kim and Nelson (1999) to approximate the value. The intuition of the filter is as follows. Given a point in time, using Kalman's filter we evaluate the likelihood function and by Hamilton's filter we evaluate the probability of being in each state. Since we may switch between k states, we have k^2 possible likelihood functions. Here, we have two states, so we are carrying four paths for the likelihood. Because our paths multiply at every observation by a factor of two, it is not tractable to use the whole history of the states. Instead, we collapse the k^2 different paths to one at the end of each time point by weighting them with the endogenously determined switching probabilities. Thus, we carry four instances of the Kalman filter at each iteration but extract a single likelihood value.

We proceed with Bayesian methods to estimate the model. We evaluate the posterior by imposing a prior on the parameters, including the coefficients of the probability matrix P. Therefore our posterior is also a function of P and the different states S, whereby

$$p(\theta, P, S|Y) = \frac{p(Y\,|\theta, P, S)\; p(S\,|P)\; p(\theta, P)}{\int p(Y\,|\theta, P, S)\; p(S\,|P)\; p(P, \theta)\; d(\theta, P, S)}. \tag{15}$$

We maximize the posterior using the Covariance Matrix Adaptation Evolution Strategy (CMA-ES).[12] This strategy uses a variance-covariance matrix to search for the maximum. Thus, it avoids the need to calculate numerical derivatives and has an advantage when the function has discontinuities, ridges, or local optima (which is more likely in the Markov-switching case compared to a standard DSGE model).[13] We employ a Markov-Chain Monte Carlo (MCMC) procedure to approximate the posterior distribution. For each model we estimate, we initiate four runs of 250,000 draws from which the first 50,000 are discarded and the rest are thinned by saving every 20^{th} draw to reach a sample of 10,000 per batch. In all cases, the parameters converge to the same means. Further diagnostics, graphs and convergence tables may be found in the appendix.[14]

[11] For a more detailed explanation the reader is referred to Cho (2011).

[12] See Andreasen (2008).

[13] Hansen (2006) or Van Binsbergen et al. (2012), p. 640.

[14] An alternative approach to the MCMC method is to use a GIBBS sampler, or more precisely "Metropolis within GIBBS" as in Bianchi (2012).

4 Data

We have seven exogenous variables: technology shock a_t, preference shock ϑ_t, import cost-push shock $\mu_{F,t}$, risk premium shock ϕ_t, and y_t^*, π_t^* and i_t^* (foreign output, inflation and interest rate). Accordingly, we choose seven series for the estimation. The domestic economy is represented by output, inflation, consumption, terms of trade and the HIBOR series. For the foreign variables, we take US GDP and the USD LIBOR. The data spans a period that extends from the first quarter of 1986 to the last quarter of 2012 (i.e. 108 observations). GDP is measured as quarterly real GDP per capita growth rate (HP filtered with $\lambda = 1600$). The inflation rate is the log difference of quarterly CPI. Consumption is measured as quarterly real consumption growth (also HP filtered with $\lambda = 1600$). Terms of trade are in logs and a standard measurement error is added. HIBOR and LIBOR are taken as is. All variables have been seasonally adjusted, scaled by 100 and, with the exception of real GDP, demeaned. The data has been collected primarily from the Hong Kong Monetary Authority, the statistical office and Datastream. This gives rise to the following measurement equation where $\pi^{(q)}$, $i^{(q)}$ and $i^{(q)*}$ denote the means of the variables

$$\begin{bmatrix} \triangle GDP_t \\ \triangle CONS_t \\ INFL_t \\ HIBOR_t \\ TOT_t \\ \triangle GDP_t^{US} \\ LIBOR_t \end{bmatrix} = \begin{matrix} y_t \\ c_t \\ \pi^{(q)} + \pi_t \\ i^{(q)} + i_t \\ v_t + R_v \\ y_t^* \\ i^{(q)*} + i_t^* \end{matrix} .$$

We estimate two core models: a standard DSGE model with no regime switching $\mathcal{M}_1$ and a Markov-switching version $\mathcal{M}_2$. In the next section we look at the priors and posterior estimates of both models.

	Distribution	Prior Mean	$\mathcal{M}_1$	$\mathcal{M}_2 : S_t = 1$	$\mathcal{M}_2 : S_t = 2$
p_{11}	*Beta*	0.950	—	0.961 [0.904, 0.993]	—
p_{22}	*Beta*	0.950	—	0.964 [0.925, 0.991]	—
β	*PM*	0.983	0.983	0.983	—
φ	*Gamma*	2.000	2.010 [1.625, 2.431]	2.029 [1.639, 2.458]	—
θ_H	*Beta*	0.375	0.861 [0.834, 0.887]	0.854 [0.825, 0.881]	—
θ_F	*Beta*	0.375	0.843 [0.812, 0.874]	0.846 [0.814, 0.878]	—
α	*PM*	0.500	0.500	0.500	—
σ	*Gamma*	1.000	2.684 [1.752, 3.809]	2.524 [1.564, 3.752]	—
η	*Gamma*	2.000	2.282 [1.895, 2.701]	2.412 [2.026, 2.815]	—
h	*Beta*	0.200	0.565 [0.459, 0.666]	0.575 [0.461, 0.682]	—
δ_H	*Beta*	0.200	0.422 [0.281, 0.564]	0.426 [0.291, 0.567]	—
δ_F	*Beta*	0.200	0.712 [0.602, 0.811]	0.706 [0.603, 0.802]	—
χ	*Gamma*	0.010	0.014 [0.009, 0.019]	0.017 [0.013, 0.021]	—
ρ_a	*Beta*	0.700	0.908 [0.777, 0.975]	0.905 [0.777, 0.973]	—
ρ_{μ_F}	*Beta*	0.700	0.918 [0.830, 0.972]	0.894 [0.790, 0.962]	—
ρ_ν	*Beta*	0.700	0.546 [0.381, 0.713]	0.541 [0.374, 0.703]	—
ρ_ϕ	*Beta*	0.700	0.705 [0.531, 0.857]	0.697 [0.524, 0.844]	—
c_{y^*}	*Beta*	0.850	0.900 [0.825, 0.968]	0.891 [0.820, 0.957]	—
c_{π^*}	*Beta*	0.850	0.649 [0.543, 0.743]	0.661 [0.562, 0.745]	—
c_{i^*}	*Beta*	0.850	0.923 [0.894, 0.951]	0.931 [0.898, 0.959]	—
σ_{μ_F}	*IGamma*	2.000	0.264 [0.202, 0.342]	0.273 [0.208, 0.354]	—
σ_a	*IGamma*	2.000	5.459 [4.216, 7.002]	5.142 [3.984, 6.628]	—
σ_ν	*IGamma*	2.000	11.001 [8.370, 14.300]	10.869 [8.290, 14.233]	—
σ_ϕ	*IGamma*	2.000	0.292 [0.260, 0.329]	0.101 [0.082, 0.126]	0.511 [0.418, 0.629]
σ_{y^*}	*IGamma*	1.000	0.550 [0.492, 0.615]	0.546 [0.488, 0.614]	—
σ_{π^*}	*IGamma*	1.000	1.540 [1.322, 1.797]	1.486 [1.282, 1.725]	—
σ_{i^*}	*IGamma*	1.000	0.134 [0.119, 0.150]	0.133 [0.119, 0.150]	—
R_v	*Normal*	0.000	-0.001 [-0.298, 0.296]	-0.000 [-0.117, 0.116]	—

Table 1: Estimated coefficients at the posterior mean. $\mathcal{M}_1$: Model with fixed parameters, $\mathcal{M}_2$: MS Model. 5% and 95% percentiles in brackets.

5 Priors and Posterior Estimates

Table 1 collects the most important information regarding the estimation stage. The second column shows the prior distribution of the parameters and the third the prior means. The last three columns consist of the estimated parameters for the different models. The 5^{th} and 95^{th} percentiles appear in brackets.

For the priors, we rely on values common in the literature and based on other studies of the Hong Kong economy. With few exceptions, many of the values are taken from Funke et al. (2011) and Funke and Paetz (2013). For parameters where the models are similar to each other and imply coherent dynamics (e.g. persistence and variance of shocks), we typically apply the Frisch elasticity of labour supply and the elasticity of substitution between domestic and foreign goods. Due to an absence of a financial sector, which would imply different price dynamics, we look to other studies to calibrate the price rigidity parameters. The estimates here seem to vary quite a bit. Genberg and Pauwels (2005) suggest a rather short price stickiness of about two to three quarters, while Cheng and Ho (2009) and Razzak (2003) seem to find values that correspond to around seven to eight quarters of constant prices. We set the prior on the price contracts fairly low, $\theta = 0.375$ based on Genberg and Pauwels (2005) and the degree of backward-looking agents δ at 0.2. For the debt sensitivity parameter χ, we follow Justiniano and Preston (2010), who use the value of 0.01 for a range of small open economies. In contrast to them, however, we do not fix this coefficient. Following the literature, we fix the discount factor β and the coefficient of openness α. The former is calibrated to match the steady state annual interest rate of 4.06% and the latter is fixed at 0.5, implying that home and foreign goods have equal shares in the consumer basket. For the foreign variables, we try to match the persistence of the series by fitting a simple AR(1) model to the data. This yields coefficients of about 0.85. The variance of the shocks is chosen so that it is smaller in the US compared to Hong Kong. Finally, we assume that probability of switching between regimes has a Beta distribution with a mean of 0.95 and a standard deviation of 0.05.

The posterior estimates match the Hong Kong economy quite well and are in line with most of the cited references. The risk aversion coefficient σ is around 2.6, which is typical in the small open economy literature. The Frisch elasticity of labour supply is not identified under this dataset and therefore centred around the prior distribution. The habit parameter $h = 0.56$ shows that consumption smoothing is an important factor in Hong Kong. The data supports rather sticky prices with $\theta = 0.86$. This is also evident in the backward-looking component $\delta_H = 0.4$. Technology and import shocks are relatively persistent with $\rho = 0.91$, while preference exhibits $\rho_\nu = 0.5$.

These findings are robust with respect to changes in the set of variables in the measurement equation, as well as taking different series for the same variable (e.g. taking the GDP deflator as proxy for inflation). Further, the addition of the real exchange rate or exclusion of the terms of trade do not affect the results. In the next section, we examine the most important parameters in this study – the time-varying coefficients.

6 Assessing the credibility of Hong Kong's limited exchange rate system

We model the credibility of the Hong Kong exchange regime as a two-state process, allowing the volatility of the interest rate to vary ($\sigma_\phi(s_t)$). Our hypothesis is that if the credibility of the system is low, agents will be willing to take positions against it that will increase the volatility of interest rates. We find two distinctive parameters suggesting heteroskedasticity of the risk premium. In the first state, the mean estimate of the coefficient is $\sigma_\phi(1) = 0.101$, while in the second, $\sigma_\phi(2) = 0.511$ (on a quarterly basis or about 2% in annual terms). The distributions do not overlap and the first state is associated with much smaller variance compared to the second as evident from Figure 2.

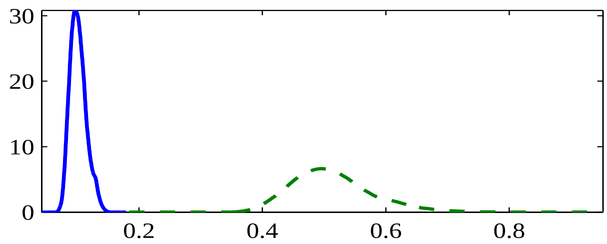

Figure 2: Posterior densities of the switching parameter under the first (—) and second regime (- -).

Conditional on the parameters, we estimate recursively the occurrence probability for each regime throughout our sample. We plot the probability for the second state ($\sigma_\phi(2) = 0.511$) in Figure 3. The bottom plot depicts the US and Hong Kong interest rates. The shaded areas represent the dates in which the probability of the second regime is above 0.5. The figure clearly indicates time variation in the risk premium on HIBOR. Three episodes are of interest. In the first, the probability peaks to one in the third quarter of 1987 and drops back after the third quarter of 1990. Next, we see a similar pattern throughout the Asian crisis, particularly between 1997Q2 and 1999Q2. Finally, there is a lone spike right before 2005 with a value of 0.8. We consider each episode in turn.

The Hong Kong stock market crashed on 19 October 1987, causing the Hong Kong market to lose almost 50% of it value.[15] The crisis spread quickly to other Asian markets, Europe and the US. Major indices such as the FTSE and Dow Jones lost over 20% of their value in a matter of days. The crash put severe pressure on Hong Kong's currency board. Our model captures this event by predicting a switch in the volatility exactly in the third quarter of 1987. An important finding is that even though the interest rates converged only two quarters later, the second regime prevailed with probability 1 for two years to come. This finding exploits the rich structure of the DSGE model. A closer look in the data shows that real GDP growth turned negative in the middle of 1989 and the economy kept shrinking for more than six quarters. With the economy recovering throughout 1992 and interest rates declining, trust in the mechanism

[15] There is a vast body of literature documenting the events and aftermath of the Black Monday. See Roll (1988), Malliaris and Urrutia (1992), and Carlson (2006).

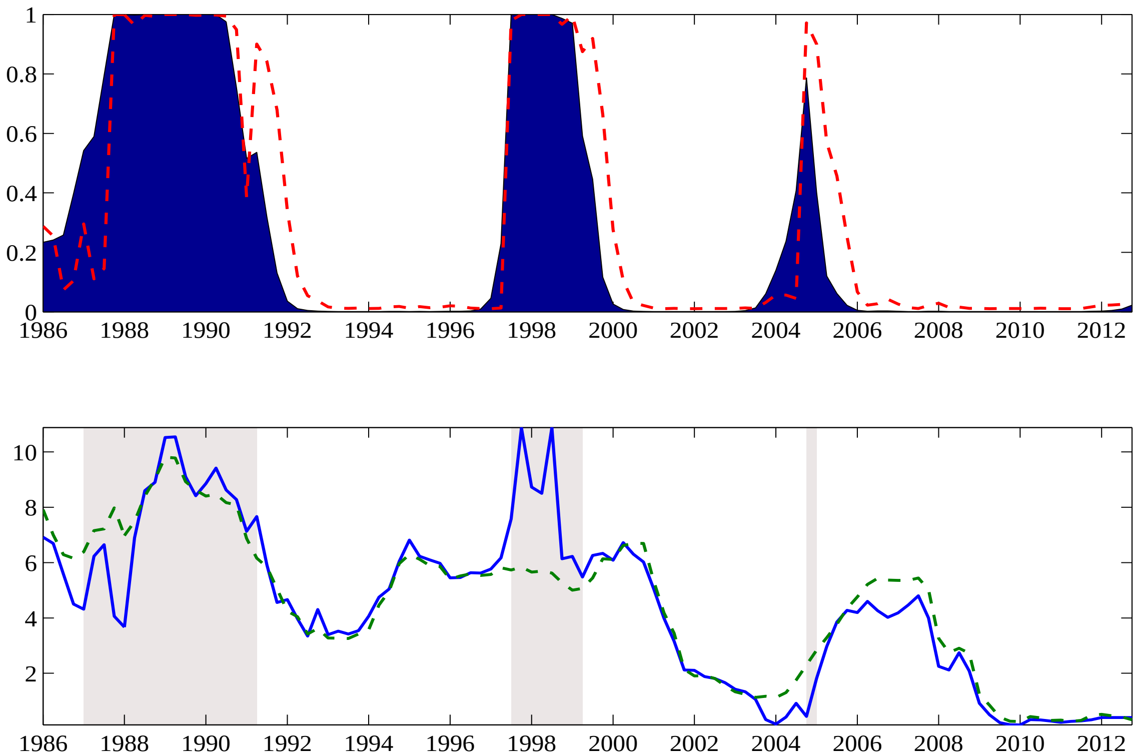

Figure 3: Top: Smoothed and not smoothed (- -) probability of the "low-credibility" state $P(\sigma = \sigma(2))$. Bottom: Annualized three-month interbank interest rates in per cent, HIBOR (—) and LIBOR (- -). Shaded areas represent the dates in which $\mathrm{P}(\sigma_\phi(s_t) = \sigma_\phi(2)) > 0.5$.

returned. Over the next seven years, the HIBOR was almost identical to the LIBOR, with the exceptions of two minor bumps in 1991 and the Mexican crisis in 1995.

The Asian crisis provoked considerable speculation against the HKD in futures markets. The 3-month HIBOR reached an all-time high, rising even more than during the Black Monday aftermath, while daily levels jumped to 16%–18% as speculators bet against the currency board on futures markets. The credibility of the linked exchange rate system was again put into question as interest rates surged. The HKMA responded in September 1998 with seven technical measures to strengthen the mechanism.[16] Those measures included a weak-side commitment against speculative attacks and depreciation and easing the borrowing conditions for the banks. The interest rate differential fell from 5% in the second quarter of 1998 to 0.8% in the third before returning to almost zero levels towards the end of 1999. Our findings suggest an almost immediate reaction to the stance taken by the HKMA with a delay of only one quarter. A similar result has been found in the study of Genberg and Hui (2011), who assess the credibility of the limited exchange rate system from an econometric perspective without introducing a specific structural model, and in Kwan et al. (2001), who look at credibility from a target-zone perspective.

The third episode appears to have been short lived. In 2004, the HKD was put under appreciation pressure. The futures market drove the interest rates down over the expectation that the HKMA would follow potential moves from the mainland for appreciation against the

[16] The official press release is available at the HKMA website: http://www.hkma.gov.hk/eng/key-information/press-releases/1998/980905.shtml.

dollar.[17] Since the technical measures of 1998 introduced only a weak-side commitment, the system was ill-prepared to cope with pressures on the strong side. In May 2005, the currency board was modified to create a symmetric band around the rate of USD/HKD 7.8. This helped calm the markets and narrow the interest-rate differential.

We find no evidence for a switch in the regimes throughout the financial crisis, even though there seem to be a negative differential similar to that of 2004. This supports our structural method; the stability of the mechanism is never questioned throughout the crisis and the monetary authority is never pushed to act.

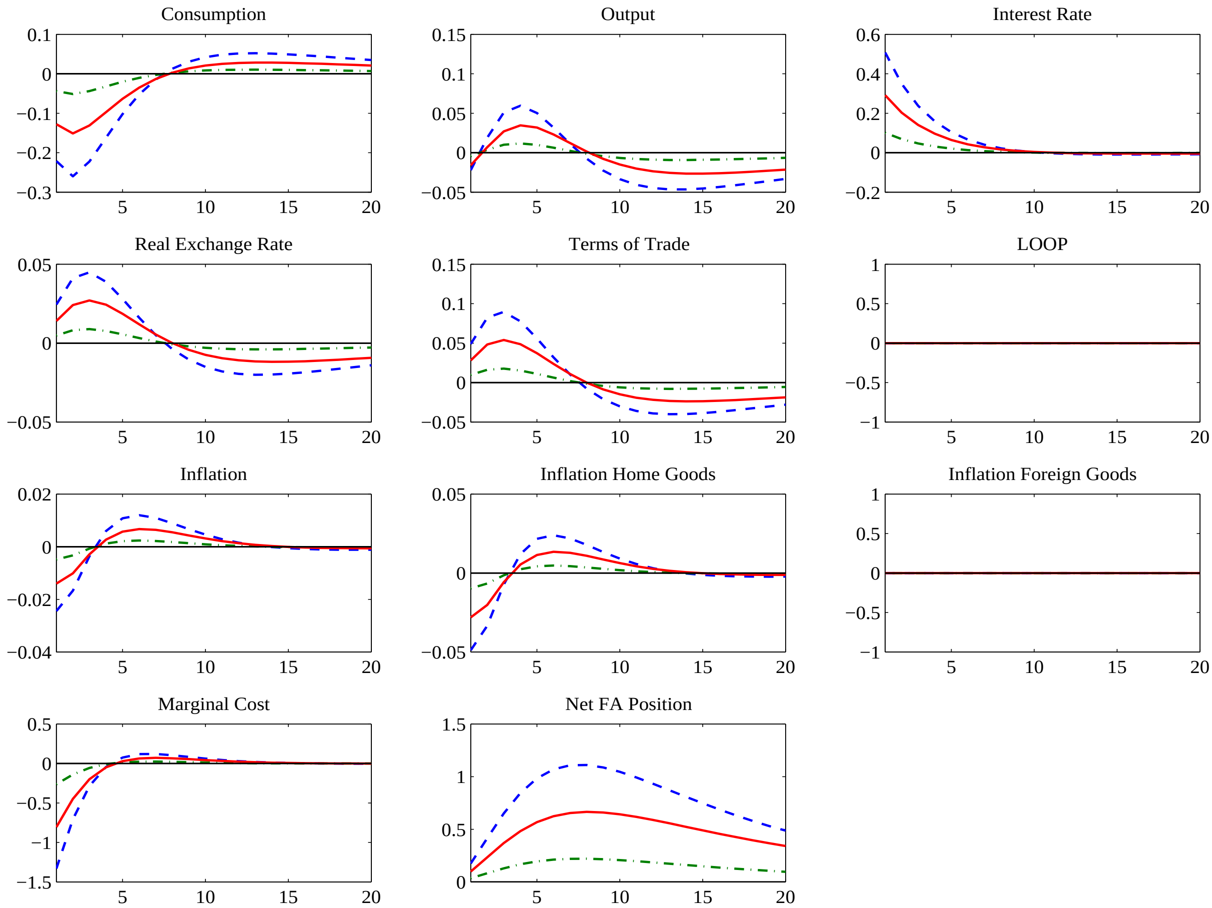

Figure 4: Impulse Responses following a risk premium shock for State 1: "high credibility" $\sigma_\phi(1)$ (-.-) and State 2: "low credibility" $\sigma_\phi(2)$ (- -) and the no-switching version $\mathcal{M}_1$ (—) .

Our framework allows us to analyze responses of the macroeconomic variables in each regime separately. Figure 4 shows the impulse responses of our variables following a shock to the interest rate such as a banking crisis or exchange rate pressures. We compare the Markov-switching model, represented by both dashed lines to the standard DSGE version, depicted by the solid red line. The green dash-dot line (-.-) is the "high-credibility" state and the blue dashed line is the "low-credibility" state. When agents trust the currency board, interest rate shocks play a negligible role on the macroeconomic variables, both real and nominal. The blue dashed line (- -) is the impulse response in a "low-credibility" state, and we see that it is much more severe. Initially, consumption and consumption growth fall sharply, around five times

[17] Genberg and Hui (2011), p. 289.

more than in the “high-credibility” state. Due to the habit formation, consumption declines even further before starting slowly to return to the steady state. Output falls initially but does not react as strongly as consumption. Combined with lower prices of domestic goods, cheaper production costs and a fixed exchange rate, GDP growth turns temporarily positive, albeit on a small scale. The economy becomes more competitive with falling domestic prices and the terms of trade improve. With the interest rate abroad lower, agents invest in foreign assets.

Crisis periods have particularly nonlinear effects on the economy because they can induce an adverse feedback loop. A low credibility regime leads to a widening of interest rate spreads, which leads to a contraction of GDP that worsens financial market conditions and widens interest rate spreads further. This, in turn, leads to a further contraction of GDP, and so on. Faced the possibility of an adverse feedback loop, the HKMA likely needs to pursue aggressively a transparent and credible commitment to a specific exchange rate target.

With a standard model, we are only able to cover a “middle-ground” scenario. The impulse responses overestimate the reactions of macroeconomic variables during times when the board is perceived as credible and underestimates the nature of interest rate shocks during the “non-credible” regime.

This class of models permits estimating the conditional variance decomposition and assessing the contribution of individual shocks to the volatility of the interest-rate series.[18] The left panel of Figure 5 shows the decomposition in the “high-credibility” state, which is dominated by variation in the foreign interest rate (up to 60%). The rest of the variance is initially attributed to its own volatility. With the advance of time, technology, the main driver behind the economy kicks in. The right panel shows a substantially different picture. Even after four years, the main volatility is attributed to the variance in the risk premium. It appears that once agents lose trust in the currency board system, it takes a long time for them to gain it back without intervention.

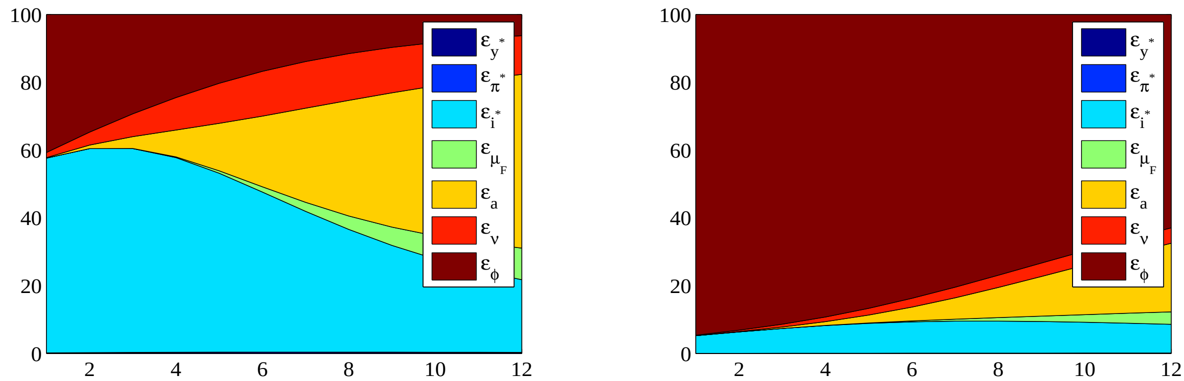

Figure 5: Variance decomposition of the interest rate for State 1: “high credibility” $\sigma_\phi(1)$ and State 2: “low credibility” $\sigma_\phi(2)$. The X-axis shows the time in quarters, the Y-axis is in per cent.

[18] The full tables with all variables may be found in the Appendix, section A.2.

7 Robustness

We perform two types of robustness checks to test the sensitivity of the results. The first is to estimate the same $\mathcal{M}_2$ specification with various data series. We explore using the GDP deflator as a proxy for inflation, excluding the terms of trade and substituting with US inflation. Furthermore, we estimate the model with series for the real exchange rate. In all cases, the identified time-varying coefficients are highly similar and the endogenously estimated regime probabilities do not change.[19]

The second type of checks is to estimate different models with the same set of data series, but specified by different switching coefficients. Due to the nature of general equilibrium models, the more flexible Markov-switching specification may allow for peculiarities of other time series to affect the estimation of the additional coefficient. Therefore, we allow simultaneous switching in all shocks and compare the results for the interest rate.

We present briefly the key results from the model with switching in all the shocks.[20] Our findings remain unchanged. Figure 6 displays the posterior densities of the switching parameters σ_ϕ (risk premium), σ_a (technology), σ_{μ_F} (inflation of imports) and σ_ν (preferences). The risk premium coefficient is around 0.1 for the "high credibility regime" and 0.5 for the "low credibility regime", which is exactly as in core model $\mathcal{M}_2$. Switching in other parameters cannot be detected as the posterior densities largely overlap. This supports our modelling in two ways. First, this is evidence that the captured heteroskedasticity is indeed a product of the interest rate and does not feed in from other variables in the structural model. Second, it shows that additional switching parameters for the volatilities are not needed as provide no further insights. The remaining parameters of the extended model are also similar to the main findings.

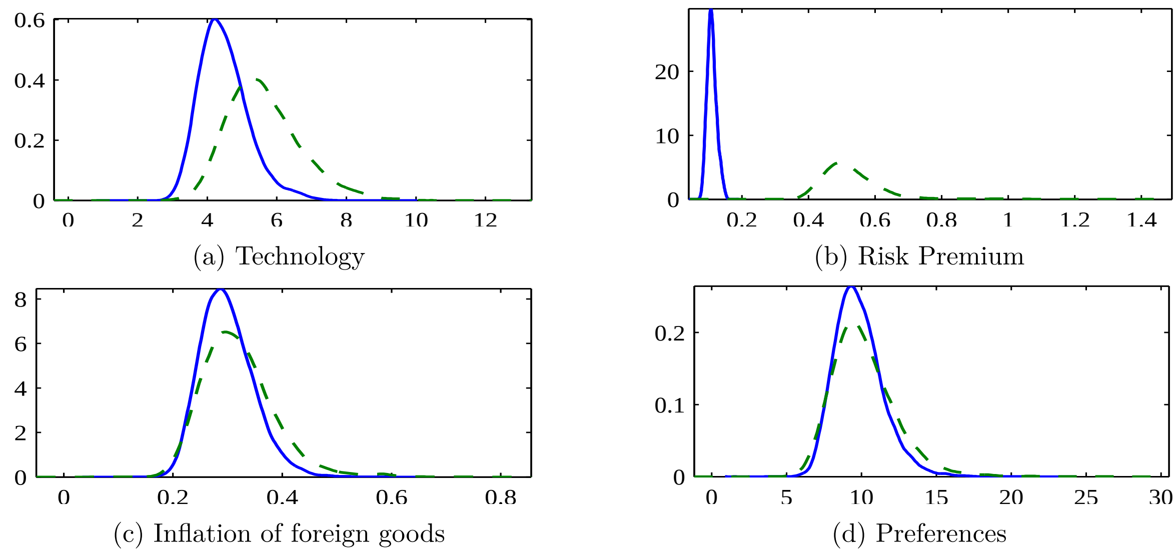

Figure 6: Posterior densities of the switching parameter under the first (—) and second regime (- -).

[19] Additional estimation results, graphs and tables are available upon request.

[20] Section A.3 in the Appendix presents a full table with the estimated parameters of this model.

8 Conclusions

This article provides a fresh look at the credibility of Hong Kong's linked exchange rate system through the lens of a structural model with stochastic volatility. Utilizing a novel Markov-switching DSGE approach, we extract evidence from financial information that the currency board has faced a loss of credibility during several prolonged periods, even during times when interest-rate differentials have otherwise been negligible.[21]

We are essentially modelling the exchange rate regime credibility as a non-linear process with two distinct regimes. Under this setup, we can see that in periods of high credibility the economy barely reacts to interest-rate shocks, yet in times of speculation against the exchange rate mechanism, the system is much more sensitive than the standard model would predict. Through conditional variance decomposition, we show that the loss of credibility may have prolonged effects before trust in the system is restored. Indeed, after the Asian crisis and during the appreciation pressure in 2005, the HKMA had to step up and strengthen the currency board before credibility could be restored.

A drawback of the estimated models is that they cannot capture the endogeneity of regime shifts. The switching parameters are exogenous, so the analysis does not allow for counterfactual policy analysis. To capture the effects of policy, one needs to know how the parameters of the Markov-switching process would have evolved for other policies. This, of course, is the Lucas critique and requires endogenization of the switching parameters in the tradition in Filardo (1994).

[21] The ability of Markov-switching frameworks to generate non-trivial connections between the dynamics of the endogenous variables and the level of uncertainty is particularly intriguing in light of the attention that uncertainty has recently received [see Bloom (2009)].

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A Appendix

A.1 Convergence diagnostics - figures and tables

In this section we present the convergence diagnostics for the Markov-switching specification, which naturally is more complicated to estimate compared to a standard case. In the interest of space, the data refers to only one of the four runs that we initiate. The other tables and figures are available upon request.

	Lag 1.	Lag 5	Lag 10	Lag 50
p_{11}	0.537	0.134	0.059	0.007
p_{22}	0.661	0.192	0.058	-0.001
φ	0.738	0.236	0.054	-0.017
θ_H	0.753	0.298	0.107	-0.023
θ_F	0.739	0.252	0.054	-0.005
σ	0.784	0.387	0.183	-0.053
η	0.724	0.236	0.055	0.008
h	0.751	0.272	0.098	0.008
δ_H	0.749	0.253	0.067	-0.016
δ_F	0.717	0.220	0.049	0.008
χ	0.723	0.260	0.086	-0.014
ρ_a	0.881	0.630	0.460	-0.023
ρ_{μ_F}	0.815	0.441	0.242	-0.009
ρ_ν	0.721	0.241	0.076	0.006
ρ_ϕ	0.727	0.223	0.044	-0.009
c_{y^*}	0.730	0.233	0.075	0.011
c_{π^*}	0.759	0.302	0.091	-0.009
c_{i^*}	0.733	0.258	0.074	0.013
σ_{μ_F}	0.788	0.362	0.169	-0.020
σ_a	0.815	0.417	0.217	-0.024
σ_ν	0.801	0.397	0.197	-0.040
σ_ϕ	0.780	0.372	0.176	0.022
σ_{y^*}	0.733	0.221	0.077	0.022
σ_{π^*}	0.751	0.261	0.064	0.057
σ_{i^*}	0.729	0.203	0.044	-0.005
R_v	0.709	0.206	0.071	0.008
σ_ϕ	0.739	0.398	0.247	0.002

	Thin	Burn	Total(N)	(Nmin)	I-stat
p_{11}	2.000	12.000	3124.000	937.000	3.334
p_{22}	2.000	12.000	3124.000	937.000	3.334
φ	2.000	12.000	3124.000	937.000	3.334
θ_H	2.000	12.000	3124.000	937.000	3.334
θ_F	2.000	12.000	3124.000	937.000	3.334
σ	2.000	12.000	3124.000	937.000	3.334
η	2.000	12.000	3124.000	937.000	3.334
h	2.000	12.000	3124.000	937.000	3.334
δ_H	2.000	12.000	3124.000	937.000	3.334
δ_F	2.000	12.000	3124.000	937.000	3.334
χ	2.000	12.000	3124.000	937.000	3.334
ρ_a	2.000	12.000	3124.000	937.000	3.334
ρ_{μ_F}	2.000	12.000	3124.000	937.000	3.334
ρ_ν	2.000	12.000	3124.000	937.000	3.334
ρ_ϕ	2.000	12.000	3124.000	937.000	3.334
c_{y^*}	2.000	12.000	3124.000	937.000	3.334
c_{π^*}	2.000	12.000	3124.000	937.000	3.334
c_{i^*}	2.000	12.000	3124.000	937.000	3.334
σ_{μ_F}	2.000	12.000	3124.000	937.000	3.334
σ_a	2.000	12.000	3124.000	937.000	3.334
σ_ν	2.000	12.000	3124.000	937.000	3.334
σ_ϕ	2.000	12.000	3124.000	937.000	3.334
σ_{y^*}	2.000	12.000	3124.000	937.000	3.334
σ_{π^*}	2.000	12.000	3124.000	937.000	3.334
σ_{i^*}	2.000	12.000	3124.000	937.000	3.334
R_v	2.000	12.000	3124.000	937.000	3.334
σ_ϕ	2.000	12.000	3124.000	937.000	3.334

Table 2: Left: Autocorrelation among the draws, based on a sample of 10,000.
Right: Raferty-Lewis convergence diagnostics with q=0.025, r=0.1, s=0.95. I-statistic larger than 5 indicates convergence problems.

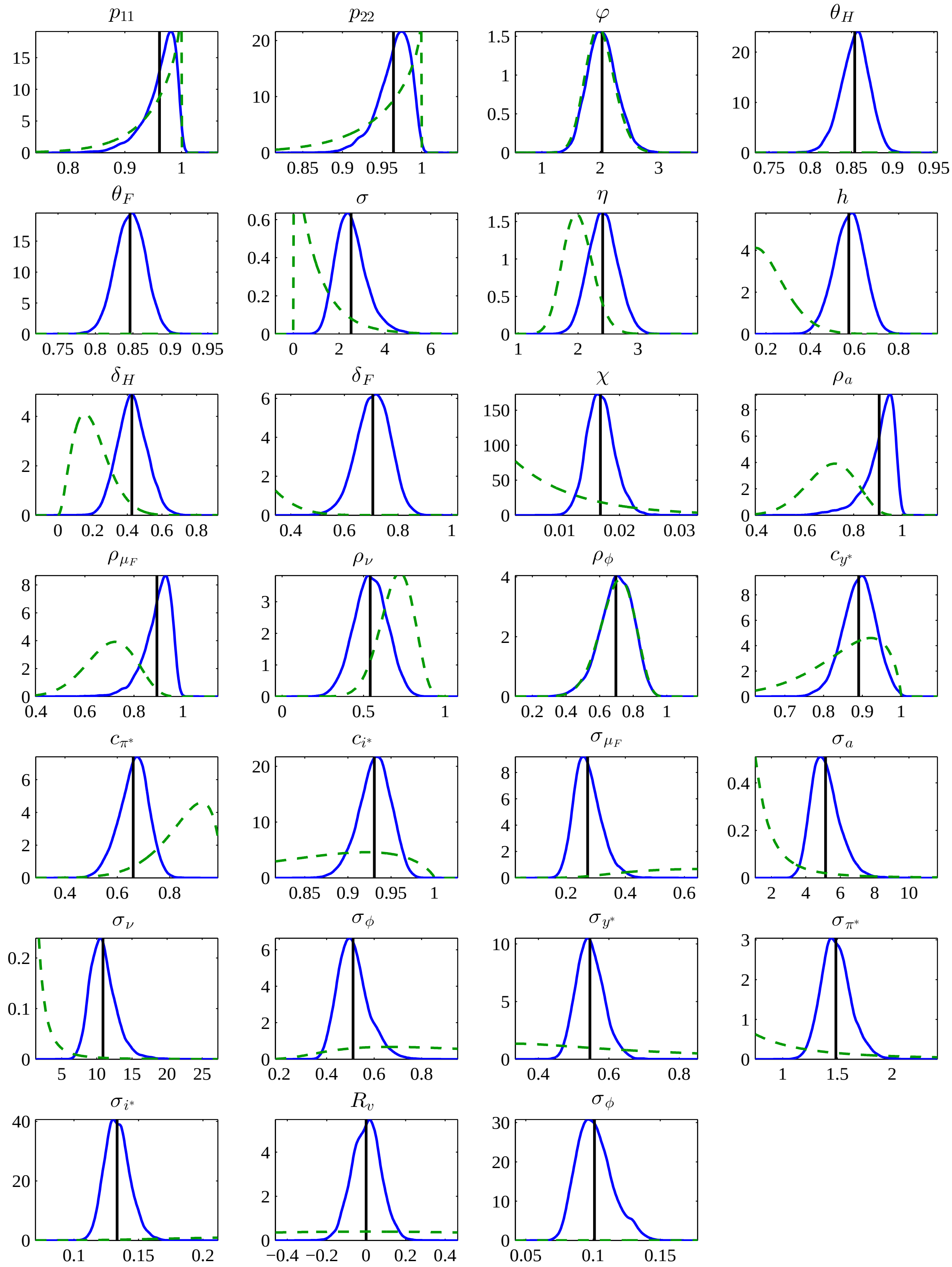

Figure 7: Prior (dashed) and posterior (solid) distributions of $\mathcal{M}_2$.

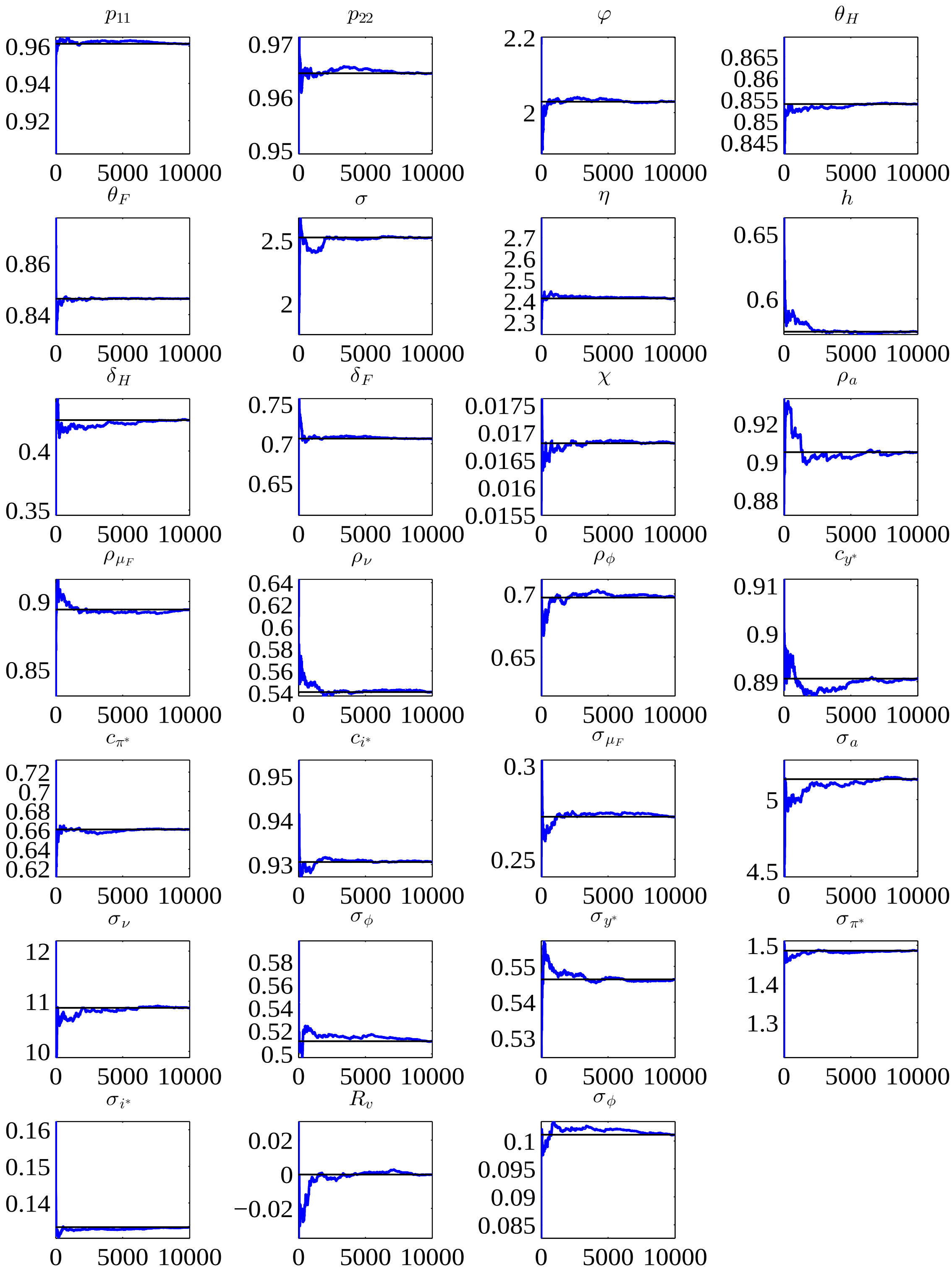

Figure 8: Recursive means of $\mathcal{M}_2$. Black line indicates the posterior mean.

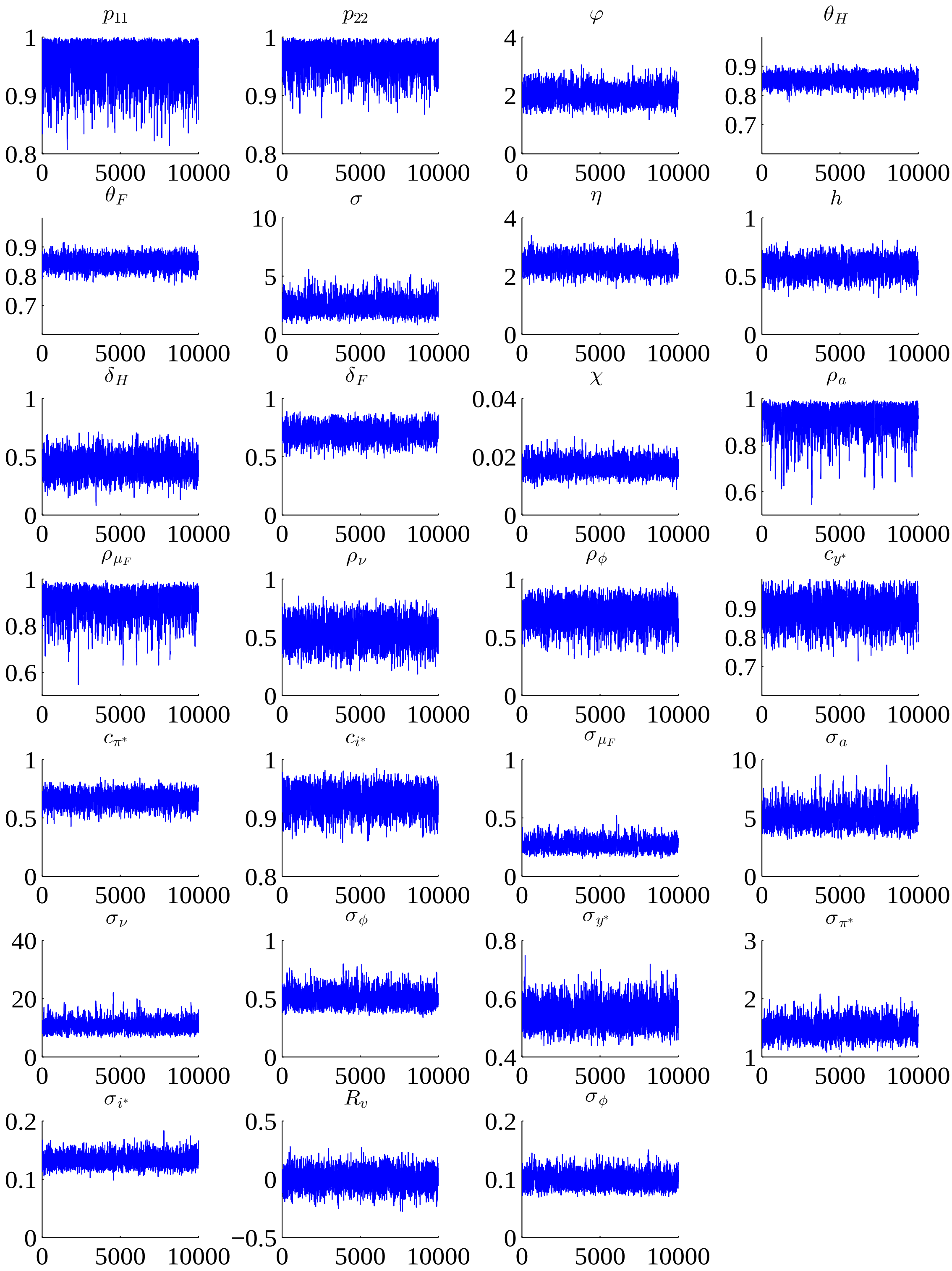

Figure 9: Trace plots of $\mathcal{M}_2$.

A.2 Variance decomposition tables

	ε_{y^*}	ε_{π^*}	ε_{i^*}	ε_{μ_F}	ε_a	ε_ν	ε_ϕ
1	0.02	7.97	0.42	3.41	10.77	77.35	0.06
4	0.04	10.31	0.60	3.35	29.74	55.91	0.05
8	0.05	8.47	0.53	2.57	46.47	41.87	0.04
12	0.05	7.05	0.45	2.19	55.08	35.15	0.03
20	0.06	5.78	0.38	2.06	62.51	29.17	0.03
40	0.06	5.05	0.41	2.86	65.94	25.65	0.03
∞	0.06	4.97	0.43	3.05	66.21	25.25	0.03

Table 3: Forecast error variance decomposition of consumption for selected periods. State 1: "high credibility" $\sigma_\phi(1)$, in per cent.

	ε_{y^*}	ε_{π^*}	ε_{i^*}	ε_{μ_F}	ε_a	ε_ν	ε_ϕ
1	0.02	7.85	0.42	3.36	10.61	76.20	1.55
4	0.04	10.17	0.59	3.30	29.34	55.17	1.38
8	0.05	8.38	0.53	2.55	46.00	41.45	1.04
12	0.05	6.99	0.45	2.17	54.61	34.85	0.88
20	0.06	5.74	0.38	2.05	62.07	28.96	0.74
40	0.06	5.02	0.41	2.85	65.52	25.49	0.66
∞	0.06	4.94	0.43	3.03	65.80	25.09	0.65

Table 4: Forecast error variance decomposition of consumption for selected periods. State 2: "low credibility" $\sigma_\phi(2)$, in per cent.

	ε_{y^*}	ε_{π^*}	ε_{i^*}	ε_{μ_F}	ε_a	ε_ν	ε_ϕ
1	0.04	57.80	0.05	31.08	8.01	3.01	0.00
4	0.02	52.01	0.02	42.50	4.37	1.08	0.00
8	0.01	54.49	0.02	40.24	4.03	1.21	0.00
12	0.01	53.28	0.02	41.20	4.31	1.17	0.00
20	0.01	50.40	0.02	44.18	4.27	1.11	0.00
40	0.02	49.79	0.02	44.85	4.23	1.10	0.00
∞	0.02	49.77	0.02	44.86	4.23	1.10	0.00

Table 5: Forecast error variance decomposition of inflation for selected periods. State 1: "high credibility" $\sigma_\phi(1)$, in per cent.

	ε_{y^*}	ε_{π^*}	ε_{i^*}	ε_{μ_F}	ε_a	ε_ν	ε_ϕ
1	0.04	57.75	0.05	31.06	8.01	3.01	0.08
4	0.02	52.00	0.02	42.48	4.37	1.08	0.03
8	0.01	54.47	0.02	40.22	4.03	1.21	0.04
12	0.01	53.26	0.02	41.19	4.30	1.17	0.04
20	0.01	50.38	0.02	44.17	4.27	1.11	0.04
40	0.02	49.77	0.02	44.83	4.23	1.10	0.04
∞	0.02	49.75	0.02	44.84	4.23	1.10	0.04

Table 6: Forecast error variance decomposition of inflation for selected periods. State 2: "low credibility" $\sigma_\phi(2)$, in per cent.

	ε_{y^*}	ε_{π^*}	ε_{i^*}	ε_{μ_F}	ε_a	ε_ν	ε_ϕ
1	0.66	45.73	0.26	0.32	50.67	2.36	0.00
4	0.10	12.54	0.33	0.05	82.94	4.03	0.01
8	0.07	7.13	0.29	0.20	89.86	2.45	0.00
12	0.07	6.19	0.27	1.20	89.83	2.43	0.00
20	0.08	5.81	0.27	4.61	86.41	2.81	0.00
40	0.09	5.45	0.37	7.87	83.29	2.92	0.01
∞	0.09	5.36	0.41	8.21	83.04	2.89	0.01

Table 7: Forecast error variance decomposition of output for selected periods. State 1: "high credibility" $\sigma_\phi(1)$, in per cent.

	ε_{y^*}	ε_{π^*}	ε_{i^*}	ε_{μ_F}	ε_a	ε_ν	ε_ϕ
1	0.66	45.72	0.26	0.32	50.66	2.36	0.02
4	0.10	12.52	0.33	0.05	82.82	4.03	0.16
8	0.07	7.12	0.29	0.20	89.76	2.45	0.11
12	0.07	6.19	0.27	1.20	89.74	2.43	0.10
20	0.08	5.81	0.27	4.61	86.30	2.81	0.13
40	0.09	5.45	0.37	7.86	83.18	2.91	0.14
∞	0.09	5.35	0.41	8.20	82.93	2.88	0.14

Table 8: Forecast error variance decomposition of output for selected periods. State 2: "low credibility" $\sigma_\phi(2)$, in per cent.

	ε_{y^*}	ε_{π^*}	ε_{i^*}	ε_{μ_F}	ε_a	ε_ν	ε_ϕ
1	0.14	0.03	57.44	0.01	0.13	1.53	40.73
4	0.33	0.03	57.40	0.23	7.94	9.58	24.50
8	0.36	0.02	36.19	3.99	34.11	13.78	11.55
12	0.30	0.02	21.41	9.35	51.26	11.43	6.24
20	0.22	0.01	10.95	14.82	63.20	7.69	3.10
40	0.17	0.01	6.96	16.91	68.47	5.52	1.96
∞	0.16	0.01	6.63	17.02	69.07	5.26	1.85

Table 9: Forecast error variance decomposition of the interest rate for selected periods. State 1: “high credibility” $\sigma_\phi(1)$, in per cent.

	ε_{y^*}	ε_{π^*}	ε_{i^*}	ε_{μ_F}	ε_a	ε_ν	ε_ϕ
1	0.01	0.00	5.22	0.00	0.01	0.14	94.61
4	0.05	0.00	8.18	0.03	1.13	1.37	89.23
8	0.09	0.01	9.43	1.04	8.89	3.59	76.94
12	0.12	0.01	8.46	3.69	20.25	4.52	62.96
20	0.12	0.00	6.22	8.41	35.87	4.37	45.01
40	0.11	0.00	4.70	11.42	46.24	3.73	33.79
∞	0.11	0.00	4.56	11.71	47.53	3.62	32.46

Table 10: Forecast error variance decomposition of the interest rate for selected periods. State 2: “low credibility” $\sigma_\phi(2)$, in per cent.

A.3 $\mathcal{M}_3$: Switching in all shocks

	Distribution	Prior Mean	$\mathcal{M}_1$	$\mathcal{M}_2 : S_t = 1$	$\mathcal{M}_2 : S_t = 2$	$\mathcal{M}_3 : S_t = 1$	$\mathcal{M}_3 : S_t = 2$
p_{11}	*Beta*	0.950	—	0.961 [0.904, 0.993]	—	0.952 [0.889, 0.990]	—
p_{22}	*Beta*	0.950	—	0.964 [0.925, 0.991]	—	0.968 [0.932, 0.992]	—
β	*PM*	0.983	0.983	0.983	—	0.983	—
φ	*Gamma*	2.000	2.010 [1.625, 2.431]	2.029 [1.639, 2.458]	—	2.036 [1.652, 2.455]	—
θ_H	*Beta*	0.375	0.861 [0.834, 0.887]	0.854 [0.825, 0.881]	—	0.844 [0.814, 0.871]	—
θ_F	*Beta*	0.375	0.843 [0.812, 0.874]	0.846 [0.814, 0.878]	—	0.838 [0.803, 0.871]	—
α	*PM*	0.500	0.500	0.500	—	0.500	—
σ	*Gamma*	1.000	2.684 [1.752, 3.809]	2.524 [1.564, 3.752]	—	2.304 [1.450, 3.397]	—
η	*Gamma*	2.000	2.282 [1.895, 2.701]	2.412 [2.026, 2.815]	—	2.426 [2.044, 2.839]	—
h	*Beta*	0.200	0.565 [0.459, 0.666]	0.575 [0.461, 0.682]	—	0.569 [0.455, 0.679]	—
δ_H	*Beta*	0.200	0.422 [0.281, 0.564]	0.426 [0.291, 0.567]	—	0.437 [0.305, 0.573]	—
δ_F	*Beta*	0.200	0.712 [0.602, 0.811]	0.706 [0.603, 0.802]	—	0.723 [0.615, 0.819]	—
χ	*Gamma*	0.010	0.014 [0.009, 0.019]	0.017 [0.013, 0.021]	—	0.017 [0.013, 0.021]	—
ρ_a	*Beta*	0.700	0.908 [0.777, 0.975]	0.905 [0.777, 0.973]	—	0.901 [0.763, 0.974]	—
ρ_{μ_F}	*Beta*	0.700	0.918 [0.830, 0.972]	0.894 [0.790, 0.962]	—	0.886 [0.769, 0.961]	—
ρ_ν	*Beta*	0.700	0.546 [0.381, 0.713]	0.541 [0.374, 0.703]	—	0.520 [0.359, 0.693]	—
ρ_ϕ	*Beta*	0.700	0.705 [0.531, 0.857]	0.697 [0.524, 0.844]	—	0.685 [0.515, 0.834]	—
c_{y^*}	*Beta*	0.850	0.900 [0.825, 0.968]	0.891 [0.820, 0.957]	—	0.889 [0.817, 0.957]	—
c_{π^*}	*Beta*	0.850	0.649 [0.543, 0.743]	0.661 [0.562, 0.745]	—	0.645 [0.547, 0.731]	—
c_{i^*}	*Beta*	0.850	0.923 [0.894, 0.951]	0.931 [0.898, 0.959]	—	0.926 [0.896, 0.954]	—
σ_{μ_F}	*IGamma*	2.000	0.264 [0.202, 0.342]	0.273 [0.208, 0.354]	—	0.298 [0.228, 0.384]	0.318 [0.227, 0.433]
σ_a	*IGamma*	2.000	5.459 [4.216, 7.002]	5.142 [3.984, 6.628]	—	4.469 [3.467, 5.779]	5.614 [4.122, 7.577]
σ_ν	*IGamma*	2.000	11.001 [8.370, 14.300]	10.869 [8.290, 14.233]	—	9.782 [7.511, 12.636]	10.113 [7.246, 13.855]
σ_ϕ	*IGamma*	2.000	0.292 [0.260, 0.329]	0.101 [0.082, 0.126]	0.511 [0.418, 0.629]	0.109 [0.088, 0.135]	0.527 [0.415, 0.682]
σ_{y^*}	*IGamma*	1.000	0.550 [0.492, 0.615]	0.546 [0.488, 0.614]	—	0.546 [0.489, 0.612]	—
σ_{π^*}	*IGamma*	1.000	1.540 [1.322, 1.797]	1.486 [1.282, 1.725]	—	1.474 [1.270, 1.696]	—
σ_{i^*}	*IGamma*	1.000	0.134 [0.119, 0.150]	0.133 [0.119, 0.150]	—	0.134 [0.119, 0.151]	—
R_v	*Normal*	0.000	-0.001 [-0.298, 0.296]	-0.000 [-0.117, 0.116]	—	-0.002 [-0.118, 0.117]	—

Table 11: Alternative specification and the benchmark model $\mathcal{M}_2$.

A.4 $\mathcal{M}_3$: Convergence diagnostics

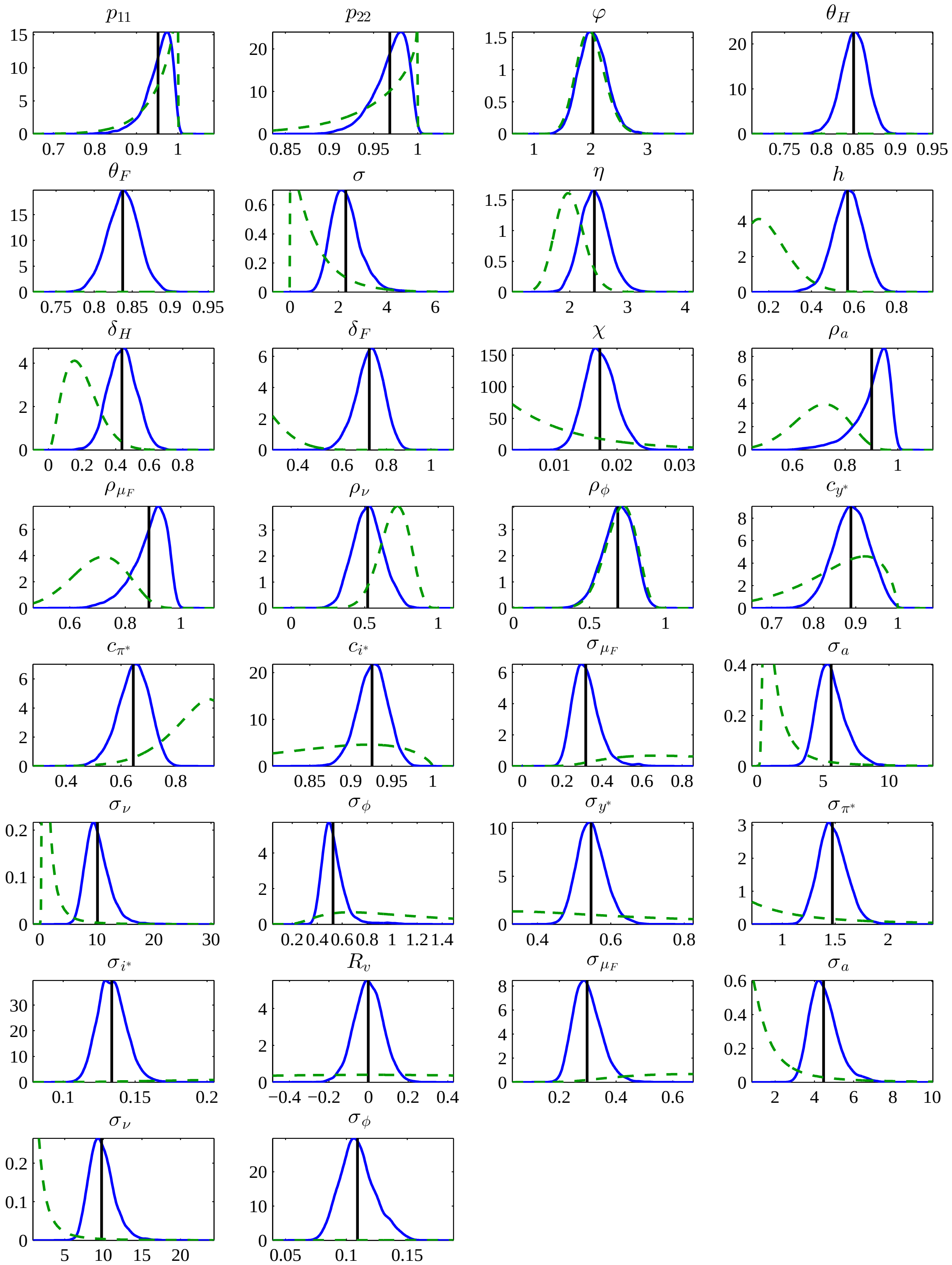

Figure 10: Prior (dashed) and posterior (solid) distributions of $\mathcal{M}_3$.

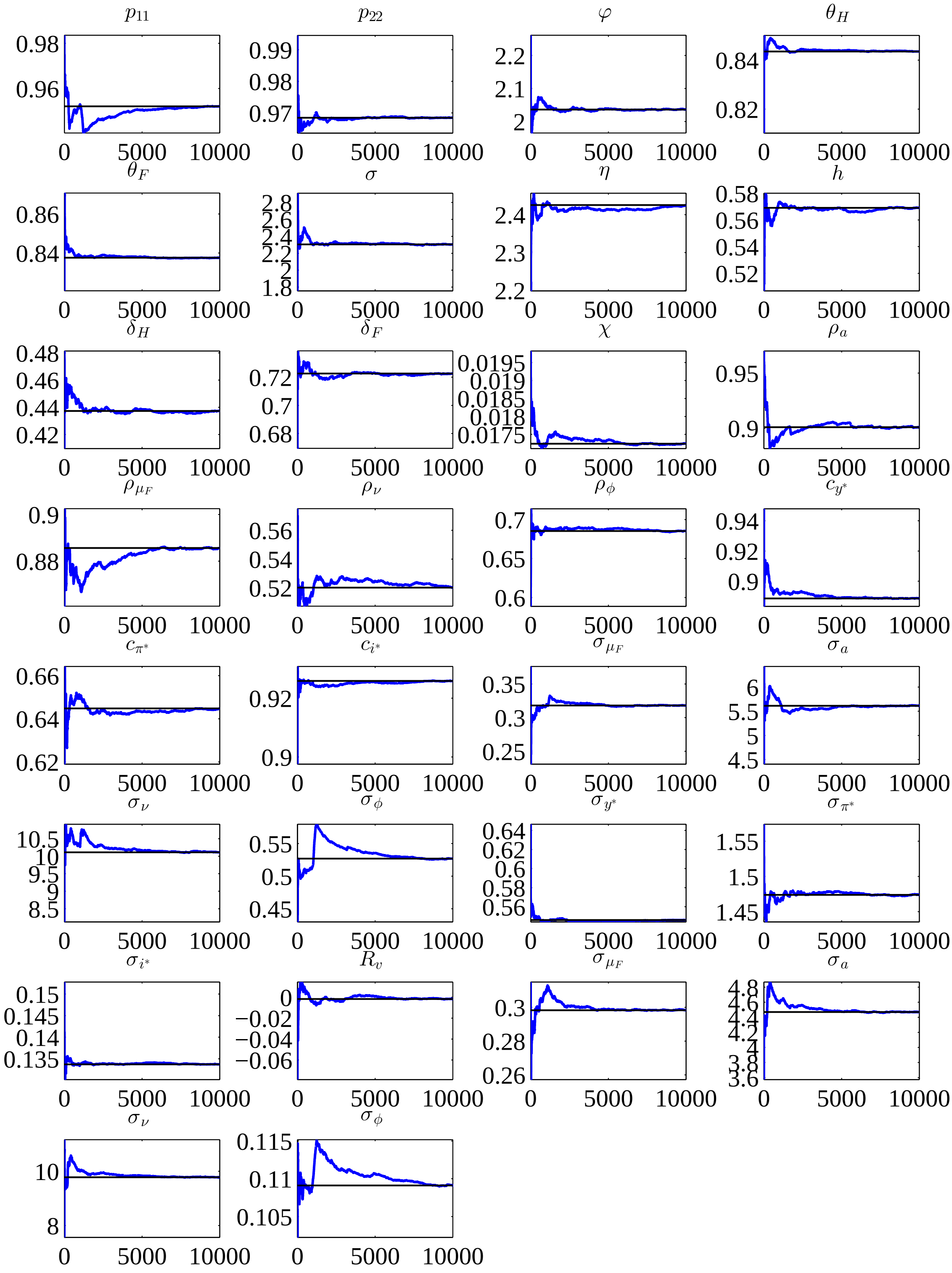

Figure 11: Recursive means of $\mathcal{M}_3$. Black line indicates the posterior mean.

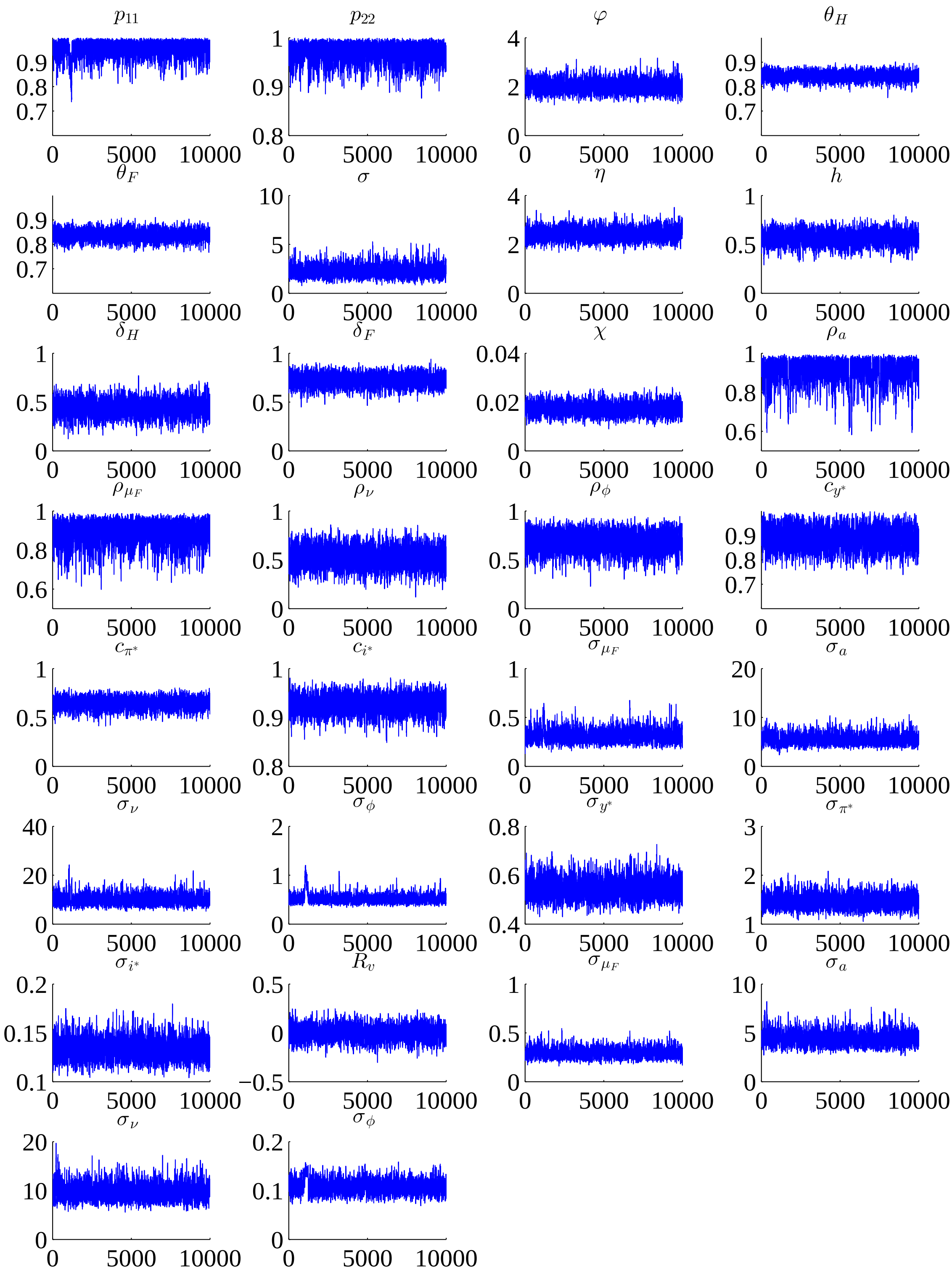

Figure 12: Trace plots of $\mathcal{M}_3$.

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