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Abstract

We employ a money-based early warning model in order to analyse the risk of a low inflation regime in the euro area, Japan and the US. The model specification allows for three different inflation regimes: Low, Medium and High inflation, while state transition probabilities vary over time as a function of monetary variables. Using Bayesian techniques, we estimate the model with data from the early 1970s up to the present. Our analysis suggests that the risks of a Low inflation regime in the euro area have been increasing in the course of the last six quarters of the sample; moreover, money growth appears to play a significant role in the assessment of such risks. Evidence for Japan and the US, on the other hand, shows that the inclusion of a monetary indicator variable does not susbstantially change the assessment of the risk of a Low inflation regime in either of the two countries.

Keywords: Money growth, deflation, inflation regimes, Markov Switching models, Bayesian inference.

JEL classification: C11, C53, E31

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1 Introduction

Since the outbreak of the financial crisis of 2008-2009, the assessment and the implications of downside risks to price stability have periodically returned under scrutiny. In particular, deep recessions like the recent global contraction would normally be expected to cause the aggregate price level to drop. Instead, inflation has barely budged across the Great Recession and its aftermath.

One accredited explanation[1] for this relatively puzzling feature hinges on the role of monetary authorities. On one hand, central banks worldwide responded timely and aggressively to the sharp contraction in economic activity and falling inflation, cutting policy rates to near zero and deploying quantitative easing to support aggregate demand. On the other hand, it has been argued that the strengthening of central banks' credibility in response to their success in delivering stable inflation over the past decade has contributed to anchor beliefs about future growth in prices and wages. Thus, more stable inflation expectations resulting from credible central banks may have contributed to weaken the link between price developments and economic fluctuations and, in turn, helped to avert deflation. Nonetheless, the 2013 outlook for the global economy has remained exposed to a great degree of uncertainty. In the G3 countries (United States, euro area, Japan), broad measures of money and credit growth have been stagnant or shrinking; unemployment has remained high while the gaps between the economies´actual output and their potential are large. Hence, the risks to price stability in the medium term are currently tilted to the downside.[2]

The question of how large these risks are has always been a widely debated one. Despite the well-recognized potential costs of deflation, relatively little is known about the economic and financial factors that contribute to deflation risk, whose definition and assessment remain subject to a considerable degree of subjectivity in the relevant economic literature.

Against this backdrop, we adopt a monetary perspective and provide a money-based early warning model in order to analyse the risk of a low-inflation regime in the euro area, as well as in Japan and in the US. The model is estimated using quarterly data spanning over the last four decades but we are particularly interested in the current inflation developments and their implications for the near future. In this paper we extend the work of Amisano and Fagan (2013), who developed an early warning indicator for shifts in inflation regimes. In particular, we model inflation dynamics using a Bayesian Markov-Switching framework in which the economy can potentially switch between three states: Low, Medium (or under control) and High inflation. Transition probabilities from one regime

[1] See, e.g., IMF (2013), Coenen and Warne (2103).
[2] See, e.g., IMF (2013).

to another are allowed to vary over time as a function of monetary variables. The idea of multiple inflation regimes finds theoretical ground in Benhabib et al. (2001), where the authors argue that economic systems that feature active interest rate feedback rules and contemplate the possibility of zero lower bound face multipla equilibria. Specifically, they demonstrate the presence of a steady state, other than the locally stable one at which monetary policy is active, in which the nominal interest rate is close to zero and inflation is possibly negative.

Our approach aims to develop a model which exploits the information content of monetary variables as 'warning signals' of the risk of the departure of inflation from the price stability regime. In other words, by using monetary variables as leading indicator of state transition probabilities we are implicitly assigning such aggregates the role of proxy of long-term inflation expectations as well as validating the fact that money growth provides a nominal anchor to the economy by reducing indeterminacy in the determination of the price level. The idea is consistent with Christiano and Rostagno (2001), who claimed that, although money is not particularly useful in forecasting inflation in the short run, a money-monitoring policy might indeed improve the economic performance of economies stuck in undesired equilibria, such as deflation, by eliminating them. Implicit here is the conjecture that inflation alternates across regimes mainly as a consequence of different monetary policy regimes.

For this reason, the analysis carried out in this paper is concerned with the prediction of inflation regimes, rather than with point forecasts of actual inflation levels. Arguably, this is a valuable piece of information for the assessment of price stability, as it provides a real time, money-based warning indicator of shifts in inflation regimes. The conclusion we draw is twofold: first, model estimations conducted with information up to the end of 2013 show that the risks of a low inflation regime have been increasing in the course of the last six quarters of the sample in the euro area, while have remained substantially unchanged in the US and even declined in Japan; second, we find that in the euro area money growth seems to play a stronger role in the assessment of such risks relatively to credit growth. In particular, we show that alternative model specifications based on different indicators such as credit aggregates, return weaker signals of the risk of euro area inflation entering the Low inflation regime. Moreover, evidence for Japan and US indicates that the inclusion of an indicator variable does not substantially change the assessment of the risk of entering a low inflation regime.

Our analysis of risks to price stability builds on a literature in which inflation is modelled as a Markov-Switching process. This strand of contributions has been extensively reviewed in Amisano and Fagan (2013), to which we refer. Nonetheless, over the last decade only, a number of diverse approaches have been used to tackle the issue of deflation risk. Decressin and Laxton (2009) a deflation vulnerability indicator originally developed in Kumar et al. (2003). In particular, Kumar et al. (2003) develop an indicator of deflation vulnerability, which they apply to a set of countries accounting for over 80 percent of world GDP. The index includes a broad range of macroeconomic variables, such as price indices, GDP growth and the output gap, the real exchange rate,equity prices, credit growth and monetary aggregates. The higher the level of indicator, the more likely it is that an economy will experience a prolonged period falling prices. Kumar et al. (2003) showed that deflationary tendencies in the global economy were somewhat higher in 2009 than during the 2002–03 deflation scare. Their finding was substantiated by means of a model-based analysis for the G3, in which the authors analyzed deflation vulnerability by employing the IMF staff's Global Projection Model, which explicitly considers the implications of the zero interest floor.

The recent contribution of Coenen and Warne (2013) builds on the strand of literature which uses structural macroeconomic models to assess the consequences of the zero lower bound on nominal interest rate. In particular, the authors employed a micro-founded open-economy model to analyse the evolution of risks to price stability in the euro area between 2008 and 2011 on the basis of model-based predictive distributions. They showed that downside risks to price stability in the euro area were considerably greater than upside risks during the first half of 2009. Following a gradual re-balancing, the risk balance turned negative again in the second half of 2011.

A loss-function-based approach to risk measurement is instead adopted in Kilian and Manganelli (2007). On the assumption that private sectors´agents can express their preference for inflation in the form of a loss function, the authors proposed formal measures of risks to price stability which feature explicit dependence from the agents´preferences. They demonstrate that risk measures are influenced by a number of factors, such as the risk aversion of the agents as well as the forecast horizon. Moreover, as empirical application of their methodology, Kilian and Manganelli (2007) estimated the risks of deflation for the United States, Germany, and Japan for different forecast horizons and, as of September 2002, found evidence of substantial deflation risks for Japan only.

Fleckenstein et al. (2013) followed a market-based approach for measuring US deflation risk and employed market prices of inflation swaps and options to extract the actual distribution of inflation. Amongst the findings of the analysis, the authors showed that the market places substantial probability weight on deflation scenarios. Their contribution validates and extends the analysis of two other recent papers: Christensen, Lopez, and Rudebusch (2011), who fit a term structure model to the Treasury real and nominal term structures and estimate the value of the implicit deflation option embedded in Treasury Inflation Protected Securities prices; and Kitsul and Wright (2012), who used inflation options to infer the risk-neutral density for inflation.

The remainder of the paper is structured as follows. Section (2) describes the econometric model and our approach to estimation. The dataset is presented in Section (3) while Section (4) summarizes the results. Section (5) details the outcome of a number of experiments aimed at assessing the robustness of our estimation results; Section (6) concludes, highlighting the avenues for ongoing and future research. In Appendix (A) we describe the data sources and in Appendix (B) we provide details on the Bayesian posterior simulation.

2 The econometric model

The model we specify, despite its relatively low level of complexity, is capable to generate inflation regimes and allows for indicator variables to act as regime predictors. In particular, we model inflation (y_t) as a stationary process that, conditional to an unobservable variable, s_t, has the following autoregressive representation:

$$\begin{aligned} y_t &= c_{s_t} + \rho y_{t-1} + \sigma_{s_t} e_t \\ e_t &\backsim NID(0,1) \\ c_{s_t} &= (1-\rho)\mu_{s_t} \\ s_t &= 1,2,3 \end{aligned} \tag{1}$$

Equation (1) shows that y_t features three different states, each of which is associated with regime-specific intercepts, c_{s_t}, and shock variances, σ_{s_t}. s_t denotes a Markov switching discrete process describing the inflation regimes: $s_t = 1$ ("Low" inflation), $s_t = 2$ ("Medium" inflation) and $s_t = 3$ ("High" inflation). Transition probabilities (henceforth TPs) across regimes are time-varying and are collected the (3×3) transition probability matrix (TPM), $\mathbf{P}_{t+1}$, defined as follows:

$$pr(s_{t+1} = j | s_t = i) = \tag{2}$$

$$\mathbf{P}_{t+1} = \begin{bmatrix} p_{11,t+1} & p_{12,t+1} & 1 - p_{11,t+1} - p_{12,t+1} \\ p_{21,t+1} & p_{22,t+1} & 1 - p_{21,t+1} - p_{22,t+1} \\ p_{31,t+1} & p_{32,t+1} & 1 - p_{31,t+1} - p_{32,t+1} \end{bmatrix} \tag{3}$$

$$p_{ij,t+1} = pr(s_{t+1} = j | s_t = i, \mathbf{z}_t), i = 1,2,3;\ j = 1,2,3 \tag{4}$$

These TPs are allowed to depend on $\mathbf{z}_t$, a set of early warning indicators, and in particular on money or credit growth aggregates. At each point in time $t+1$, TPs indicate the risks associated with regime transitions. Hence, TPs are conditional on the event that the model is in a given state at time t as well as on the lagged indicator variables. We follow Albert and Chib (1993) and use an ordered probit specification to model state transitions. In particular, given $s_{t-1} = i$,

$$\begin{aligned} s_t = j &\Longleftrightarrow \gamma_{j-1,i} < s_t^* \leq \gamma_{j,i} \\ s_t^* &= \beta_0' \mathbf{z}_t + \beta_i + \eta_t, \eta_t \sim NID(0,1) \\ cov(e_t, \eta_s) &= 0, \forall t, s \\ \gamma_{0,i} &= -\infty, \gamma_{1,i} = 0, \gamma_{m,i} = \infty \end{aligned}$$

where $\gamma_{0,i}$ and $\gamma_{m,i}$ are respectively minus and plus infinity due to the fact that s_t^* is a real valued latent variable, while $\gamma_{1,i}$ is set to 0 for identification purposes.

Hence, the free parameters are those in the β vector, i.e. the β_0 slope coefficients and the $\beta_i, i = 1,2,3$, state-specific intercepts, and the thresholds $\gamma_{h,i}, i = 1,2,3, h = 1,2,3$. In case $m = 3$, then we have $(6 + \ n_z)$ free parameters, where n_z is the number of variables in $\mathbf{z}_t$.

The resulting time-varying transition probabilities are determined as follows:

$$\begin{aligned} p_{i1,t} &= \Phi(-\beta_0' \mathbf{z}_t - \beta_i) \Leftrightarrow \beta_i = -\mathbf{z}_t' \beta_0 - \Phi^{-1}(p_{i1,t}) \\ p_{i2,t} + p_{i1,t} &= \Phi(-\beta_0' \mathbf{z}_t - \beta_i + \gamma_{2i}) \Leftrightarrow \gamma_{2i} = \beta' \mathbf{z}_t + \beta_i + \Phi^{-1}(p_{i1,t} + p_{i2,t}) \end{aligned} \tag{5}$$

A by-product of model estimation are the estimated smoothed probabilities of each state, namely the assessment of how likely each of the state is on the basis of the whole sample used for estimation

$$\begin{aligned} & p(s_t = i | \underline{\mathbf{y}}_t, \underline{\mathbf{z}}_t), i = 1,2,3; \ \tau = 1,2,...,t, \\ & \underline{\mathbf{y}}_t = \{ y_\tau : \tau = 1,2,...,t \}, \\ & \underline{\mathbf{z}}_t = \{ z_\tau : \tau = 1,2,...,t \}. \end{aligned} \tag{6}$$

Combining TPs (equation (2)) with smoothed probabilities (equation (6)) at the end of the estimation sample returns one-step-ahead regime probabilities which can be used to assess at time t the overall probability of inflation being in each one of the three states at time $(t+1)$:

$$p(s_{t+1} = j | \underline{\mathbf{y}}_t, \underline{\mathbf{z}}_t) = \sum_{i=1}^{3} p(s_t = i | \underline{\mathbf{y}}_t, \underline{\mathbf{z}}_t) \times p_{ij,t+1}, \ j = 1,2,3. \tag{7}$$

where $p(s_t = i|\underline{\mathbf{y}}_t, \underline{\mathbf{z}}_t)$ are the smoothed probabilities and $p_{ij,t+1}$ are the transition probabilites.

The model is presented here generalizes the early warning (henceforth EW) Markov Switching model used in Amisano and Fagan (2013) in order to accommodate the possible presence of three regimes. We adopt a Bayesian approach to estimation using weakly informative priors meant at conveying the prior belief that regimes tend to be persistent. Details about the choice of the priors as well as the description of the posterior simulation strategy are included in Appendix (B).

2.1 Calibration of priors on transition probabilities parameters

At this point it is important to mention that in our empirical application the values of the early warning indicators, $\mathbf{z}_t$, are standardized. This transformation is neutral and yet it pins down exactly the values of the intercepts and slopes in the ordered probit specification. Hence, it is reasonable to wonder what happens to the TPM when $\mathbf{z}_t = \mathbf{0}$, i.e.when the level of the indicator variable is aligned with its sample average. In that case we can use (5) to calibrate the prior for the parameters. As an example, let us consider a case in which we aim to center the prior around values that produce the following TPM

$$\overline{\mathbf{P}} = \mathbf{P}(\mathbf{z}_t = \mathbf{0}) = \begin{bmatrix} .80 & .19 & .01 \\ .02 & .90 & .08 \\ .005 & .14 & .85 \end{bmatrix}. \tag{8}$$

The $\overline{\mathbf{P}}$ matrix describes a situation in which the first state tends to be less persistent than the second and the second state more persistent than the third, while the third is in turn more persistent than the first, i.e. $\overline{p}_{11} < \overline{p}_{22}$, $\overline{p}_{33} < \overline{p}_{22}$, $\overline{p}_{11} < \overline{p}_{33}$. The structure of the $\overline{\mathbf{P}}$ matrix also suggests that for $s_t = 1$ and $s_t = 3$, a shift to a neighbouring state is more likely than a shift to the furthest state, while abandoning state 2 for state 3 is more likely than abandoning it for state 1 ($\overline{p}_{21} < \overline{p}_{23}$). The features displayed by (8) are indeed compatible with an interpretation of the regimes as Low, Medium and High ($s_t = 1, 2, 3$) inflation, respectively. We shall take this numerical example as reference to explain the interpretation of the resulting ordered probit parameters and as a guide to specify the central location of a prior distribution for these parameters.

Taking the point values of probabilities in (8) and using (5), the corresponding values for the ordered probit specification, i.e. intercepts, slopes and threshold parameters are:

$$\beta = \begin{bmatrix} -.84 \\ 2.05 \\ 2.57 \end{bmatrix}, \gamma = \begin{bmatrix} 1.50 \\ 3.45 \\ 1.50 \end{bmatrix}$$

These values are used as prior means. Prior uncertainty around the prior mean is modelled to be substantial and to allow for data to dominate results.

3 The dataset

We estimate our EW model for three different countries separately: the euro area (EA) Japan, and the US. This choice was motivated by the following considerations: first, data on the relevant variables are available for time spans long enough to allow the identification of regime changes; second, focusing on a set of countries instead of on a single one helps shed some light on the robustness of the results, in particular when the analyzed countries feature rather diverse inflation experiences. Finally, Japan is the one country which underwent a long deflationary phase in the course of the considered time span and as such its inflation data are expected to provide substantial insight on the characteristics of a "deflation" regime.

The country-specific data sets include quarterly data on inflation, money, output and its deflator as well as total credit. Inflation, whose sample behaviour is depicted in Figure A.1 (left column), is computed as logarithmic fourth-difference of the consumer price index. As for the variables to be included as potential indicator in our EW model, we employed both monetary and credit aggregates. In particular, for the former we opted for the broad monetary aggregate used by the respective central banks, that is M3 for the euro area, and M2 for Japan and the US. Moreover, instead of considering the raw unadjusted money growth, we developed an "adjusted" gauge which aims to correct for changes in trend in money velocity and/or in potential output. Finally, the adjusted money growth indicator was appropriately lagged and smoothed, in order to model the leading properties of money growth on inflation and to avoid for its signalling properties to be distorted by temporary shocks with no implications for future inflation (Figure A.1, right column). For technical details on the construction of the "adjusted" monetary indicator we refer to Amisano and Fagan (2013). Credit growth, the variable used in the analysis as alternative early warning indicator, is computed as logarithmic fourth-difference of loans to private sector in nominal terms, also appropriately lagged and smoothed. Table (1) summarizes the information on the series and the sample sizes being used, while the sources and the transformations applied to the data are described in Appendix (A). An additional data-related issue we had to face was the construction of historical backdata for the euro area. The procedure implemented has been thoroughly explained in Amisano and Fagan (2013), to which we refer for further details.

4 Estimation results

As already stated, we conduct our analysis on data for euro area, Japan and US. We estimate a univariate autoregressive process for inflation characterised by three regimes: a Low inflation regime, a Medium inflation regime, compatible with the notion of price stability and a High inflation regime. It is important to mention that for the euro area the Low inflation regime is calibrated, given that the occasional and short-lived negative inflation rates seen in the data throughout the estimation sample can hardly be characterised as "regimes". In particular, the upper left panel of Figure A.1 shows inflation reaching zero briefly in 1986Q4 and, much later on, dipping into negative territory in 2009Q3. In the analysis we carry out here we calibrate the euro area Low inflation state with parameter values that qualitatively correspond to the state-specific features of the parameters estimated for Japan in Amisano and Colavecchio (2013).[3] The same argument applies to the US case.

We implement MCMC posterior simulation of the parameters according to the Gibbs sampling scheme described in Appendix (B). The estimation results of the EW model with adjusted money growth as indicator variable for euro area, Japan and the US are displayed in Tables (2)-(4), respectively. For each of the parameters being estimated the tables report their mean, standard deviation and 2.5% and 97.5% quantiles of the prior and posterior distributions. In particular, in the top panel of Tables (2)-(4) are collected the parameters which describe the distribution of inflation conditional on each state (i.e. ρ, μ_i and σ_i). The figures suggest that the inflation regimes have clearly-defined features which also appear to be consistent with apriori expectations: the posterior mean of the Medium inflation regime is below 2%, in line with the current notion of price stability or inflation target, and displays a relatively subdued variance; in the case of Japan (Table (3)), the only one where the Low inflation regime has not been calibrated, state 1 is characterized by 0-mean coupled with low variance while state 3 exhibits a posterior mean above 6% and relatively larger variance. Moreover, the posterior 95% confidence sets of all coefficients in all countries (with the exception of μ_1 for Japan) never contain zero, which, according to a purely frequentist approach, is a signal of statistical significance.

The parameters determining the time-varying TPs described in the ordered probit equation (5), i.e. the slope coefficient β_0, the intercepts β_i, and the threshold parameters γ_{2i}, appear to be less well identified, mainly due to the low number of observed regime transitions in the estimation sample.

[3] Amisano and Colavecchio (2013) estimate a Markov Switching vector autoregressive model for Japan and found evidence of three different regimes of inflation, one of which displaying features and the time allocation consistent with the empirical evidence of Japanese deflationary episodes.

While the bottom panel of Tables (2) - (4) shows that for all countries the 95% confidence set for β_0 does include zero, suggesting the lack of significance of the monetary indicator in the specification on the transition probability equation, the posterior probability of the slope parameter being greater than zero is 93%, 90% and 86% for Japan, euro area and for the US, respectively. This informal check, in turn, speaks in favour of the relevance of the monetary indicator to explain the developments of the time-varying TPs. Nonetheless, more conclusive evidence might be collected by taking a fully bayesian approach and comparing the marginal likelihood of the EW model with the marginal likelihood of a model which does not include the monetary indicator variable. We perform such comparison and present some preliminary results in Section (5), in the context of model robustness analysis.

TPs can be interpreted as indicator of one-quarter-ahead regime shifts. and they are plotted in Figures (1)-(3). The respective end-of-sample values are reported in Tables (5)-(8), Panel A. Tables (5)-(8), Panel B displays the transition probabilities computed by setting the monetary indicator variable equal to its sample mean. This second set of probabilities is meant to provide a benchmark measure of regime change risks in average monetary conditions. Hence, the difference between corresponding values in Panel A and B of Tables (5)-(8) can be used to quantify the role of money growth in the assessment of risks at the end of the sample.

For the euro area, Figure (1) shows that, between the early 1980s and the burst of the financial crisis in 2007-2008, the time-varying probabilities of remaining in regime Low, $p_{11,t}$, as well as the probabilities of moving from Medium to Low inflation state, $p_{21,t}$, fluctuate around their respective average values, i.e. those obtained in average monetary conditions. From 2010 on both $p_{11,t}$ and $p_{21,t}$ increase dramatically highlighting both an increase in deflationary risks and the relevance of money growth in the assessment of such risks. Table (5) provides a closer picture of the end-of-sample figures. In particular, focussing on transition probabilities starting from regime Low (i.e. $s_t = 1$), the probability of remaining in the same regime in Low inflation regime is up to 92.5%, compared with a mean value of 80.2%, while the probability of moving to the Medium inflation state has decreased to 7.3%, as opposed to its mean value of 18.9%. Transition probabilities starting from the Medium inflation state show similar developments: the probability of going from Medium to Low inflation is up to 8.8%, from a mean of 2.6%, and the probability of remaining in Medium regime is down to 89.6% from a mean of 91.4%. These developments in transition probabilities are the direct results of the deterioration in money growth occurring since mid-2011, as shown in the upper right panel of Figure A.1.

Evidence for Japanese data is documented in Figure (2). The sharp drop observed in money growth in the mid Nineties resulted in a considerable and persistent increase in the probability of remaining in the Low inflation regime, compared with a mean value of 87%; at the same time, the probability of moving to the Medium state decreased below 8% as opposed to its mean value of 13.4%. Unsurprisingly, the evolution of the probabilities of leaving the Medium inflation regime to enter the Low inflation one mirrors those described above: in the second half of the nineties $p_{21,t}$ has reached values near to 20% and has remained around 14% ever since, against a mean value of 6.5%.

The huge spikes in US inflation in the mid and late 1970s as well as the substantial drop experienced by US money growth in the 1990s are depicted in the departure of the time-varying transition probabilities from their respective averages in Figure (3). Additionally, the EW model signals a surge in deflationary risk after 2010, which appears to have retrenched towards the end of the estimation period. The end-of-sample values summarized in Table (8) show that transition probabilities calculated at current money conditions do not differ substantially from those obtained at average monetary conditions. This in turn might suggest that in the current phase the EW variable has had a weaker effect on the probabilities of switching regimes.

All in all, the developments displayed by the time-varying TPs seem to validate the fact that money growth might indeed have a role in the assessment of the risk of inflation regime changes and that this role/impact might be different depending on the state the system finds itself into. To shed further light on this aspect, in Figures (4)-(6) we depict the relationship between TPs, computed at the posterior mean of the parameters, and indicator variable. A couple of interesting pieces of evidence, common to all countries featured in our analysis, can be drawn: first, the relationship between the monetary indicator and the probability of remaining in the Low inflation regime, $p_{11,t}$, appears to be negative; conversely, an increase in money growth would result in an increase in the probability of leave the Low inflation state. Moreover, the concave shape of the scatter plot displaying the development of $p_{22,t}$ with respect to money growth indicates that the probability of remaining in the Medium inflation state decreases when monetary conditions exceed a certain threshold, which in turn triggers an increase in the the risk of moving to the High inflation state, $p_{23,t}$.

The posterior simulation of the model provides, as a by-product, the posterior smoothed probabilities of each regime, which are plotted in Figure (7) and in Figure (8). These are the estimated probabilities of inflation being in one of the three regimes at any given point, conditioning on the whole sample evidence, as described in equation (6). The state allocation portraited is very much in line with anecdotal evidence. Interesting enough, for the euro area, since 2011Q4, the Low inflation regime smoothed probability has markedly increased and reached 45.1% in 2013Q2. It is also worth noting that the Low regime smoothed probability does not return to the peaks reached during 2009 (87.4 %), when inflation fell sharply and touched negative territory. The values at the end of sample are rather comparable with the dynamics observed between 1983Q4 and 1985Q2. Symmetrically, the smoothed probability for the Medium inflation state has fallen from 91.8% to 54.1%. These developments can be interpreted as reflecting an increased risk of the Low inflation regime.

Looking at the Japanese evidence depicted in the top panel of Figure (8), we observe that most of the estimation sample is allocated to the Low inflation regime, with a few exceptions. First, the decade 1970-1980, which featured the 70s oil price and which the model correctly assigns to the High inflation regime; second, the early 90s as well as the years of the Asian financial crisis (early 1997-1999) and, finally, the 2008-2009, when the model most likely capture a phase of increased inflation volatility. Overall, it would appear that, in a backward-looking perspective, the model does not fail to timely signal an increase in the probability of entering the low inflation state in correspondence of periods featuring decreasing inflation. As for the current phase, the EW model suggests a significant drop in the probability of the Low inflation regime, which does not come as a surprise, given the stabilization of Japanese inflation and M2 growth. At the same time, the probability of the Medium inflation regime has risen sharply, reaching 62% in 2013Q3.

As for the US case (Figure 8, bottom panel), we see that the run up to inflation in the early 1970s and inflationary spell of the late seventies are clearly picked up by the adjusted monetary indicator: the EW model assigns both episodes to the "High" inflation state. Also the acceleration of inflation taking place from 1987 until 1991 is led by money growth acceleration. Only in 2010 the model signals a substantial increase in deflation risks, in correspondence of a phase of high US inflation volatility coupled with a sharp drop in money growth. Such risks appear to have normalized towards the end of the estimation period.

Combining smoothed probabilities and transition probabilities together, it is possible to obtain one-step-ahead regime probabilities conditioned only on current observable variables, as described in equation (7). Intuitively, this corresponds to taking into consideration the uncertainty regarding the current regime. One step ahead regime probabilities are reported on the last lines of Panels A and B in Table (5). The last line of Panel A report one step ahead probabilities based on current monetary conditions, while the last line of Panel B reports one-step-ahead probabilities corresponding to average monetary conditions. It is hence possible to see that using actual monetary developments, the one-step-ahead Low regime probability is 46.5% , substantially higher than in average monetary conditions (37.5%), while the corresponding values for the one step ahead Medium regime probabilities are respectively 52.1% and 58.1%. In synthesis, the results presented can be interpreted as signalling a relevant increase in the risk of euro area inflation entering the Low inflation regime.

As for the Japanese case (Table 7), the EW model estimated with the actual monetary indicator delivers a 40.1% one-step-ahead probability of Low inflation regime as opposed to the 34.9% estimated considering average monetary conditions. The same cannot be said for the US (Table 8), whose one-step-ahead probabilities calculated at current monetary conditions do not substantially differ from those corresponding to average conditions.

5 Robustness analysis

In the two subsections that follow we assess the sensitivity of the estimation results of the model specification presented in Section (4) with respect to two different dimensions: prior settings and model specification. In particular, Subsection (5.1) addresses the former, while Subsection (5.2) focuses on the latter. Altogether, for each country, we run the following four experiments:

i. estimation of a model in which the prior for the Low inflation mean has been calibrated to 0% rather then to 0.5%;

ii. estimation of a model with looser priors on the parameters determining the time-varying transition probabilities;

iii. estimation of a model with fixed, i.e. exogenous, transition probabilities;

iv. estimation of an EW model where money growth has been replaced in the state equation by the growth rate of total loans to the private sector in nominal terms.

Due to data availability, the model specification including loans for the euro area is estimated using the 1980Q1-2013Q2 sample period, shorter than the one used to estimate the model with money growth. This circumstance somehow limits the comparability across the results of the two different model specifications.

Additional details on the outcome of these exercises are provided in Subsections (5.1) and (5.2). All in all, the results of the sensitivity analysis can be interpreted as substantially corroborating the findings of the benchmark model outlined in Section (4).[4]

[4] For the sake of brevity, we do not include the complete set of results of the robustness checks performed in this section. Additional details are available upon request.

5.1 Prior specification

In the context of prior sensitivity analysis, we focus on two crucial aspects of the benchmark model specification: first, the robustness of the estimation results with respect to the choice of the value for the mean of the Low inflation regime; second, the role of prior specification for the parameters determining the time-varying transition probabilities.

The first experiment is particularly relevant in our framework where the Low state level of inflation for the euro area and the US had to be calibrated, given that in the estimation sample low values of inflation have been experienced very seldom.[5] By estimating the same model specification with two different values for the mean of regime 1 we assess the sensitivity of the estimations with respect to the calibrated parameter. The results of this first experiment do not qualitatively differ from those presented in Section (4). In particular, both specifications provide indicators of risk of the Low inflation regimes that are largely unchanged almost throughout the entire estimation sample. Only towards the end of the sample, the model with calibrated mean of the Low inflation regime equal to 0 provides a weaker signal of entering the Low inflation regime.

The second experiment is meant to analyze the role of the priors for the parameters of the state equation in affecting the estimation results. In particular, we increase the prior standard deviations in order to allow for looser priors on the relevant parameters. Again, the results of this experiment, qualitatively very similar to those reported in Section (4), validate the robustness of the estimations also with respect to the tightness of the priors.

5.2 Model specification

One distinctive feature of the framework we implement in this paper is the exploitation of the information content of monetary variables for regime allocation. In other words, these indicators act as 'warning signals' of the risk of the departure of inflation from the price stability regime. In order to check whether the results presented in Section (4) are robust with respect to the specification of the transition probability matrix, we estimate the model with fixed transition probabilities. This experiment is conducted to evaluate how a model without indicator variables would allocate inflation to the different regimes. Resulting smoothed probabilities are coherent with those provided by the model with money growth, but at the end of the sample they tend to be smaller than those produced in the model with time-varying transition probabilities reported in Figure (1)-(3). This, in turn, seems to validate that the monetary aggregate might indeed provide valuable information for detecting the

[5] In this context, a comparison with the results obtained using Japanese data is of great interest.

risk of entering a Low inflation regime.

The last robustness check we conduct aims at assessing the regime predictor properties of money growth against those of credit aggregates. In particular, we replace money growth with the rate of growth of loans to the private sector as EW indicator, z_t, in the time-varying transition probability function. As for the previous experiments, we obtain results largely consistent with those reported in Section (4). For the sake of brevity, we report only euro area results: evidence for the regime allocation is shown in the bottom panel of Figure (7) while Table (6) summarizes the end-of-sample transition probabilities. Interestingly, it turns out that using credit growth as indicator variable produces warning signals which are weaker than those provided by the model exploiting the information content of money growth. A thorough and formalized model comparison exercise requires the calculation of the models´marginal likelihood and is currently still in progress. For this reason, we decide not to disclose any evidence in this version of the paper. Nonetheless, preliminary results seem to favour the model specification including the monetary aggregates vis-à-vis the considered alternatives, i.e. the fixed transition probability model and the EW model with credit aggregates as indicator variable, in all the analysed countries. One additional step in our agenda is the sequential evaluation of the out-of-sample predictive performance, which entails the computation of the predictive density of each observation in the sample conditioned only on the past value of the observable variables. This type of analysis is particularly useful to assess the forecasting performance of the model over different subsamples.

6 Conclusion

In this paper we employ an early warning model in order to analyse the risk of a low inflation regime in the euro area, Japan and the US. In particular, we adopt a monetary perspective and extend the work of Amisano and Fagan (2013), who developed a money-based early warning indicator for shifts in inflation regimes. In our specification inflation is characterized by a regime-switching model in which the economy can potentially switch between three states: Low, Medium (or under control) and High inflation. Moreover, transition probabilities from one regime to another are allowed to vary over time as a function of monetary variables. An important feature of this approach is that it allows to exploit the information content of these variables as 'warning signals' of the risk of the departure of inflation from a regime of price stability.

We apply the model to quarterly data spanning over the last four decades, according to country data availability, and we employ Bayesian estimation techniques. The results obtained partially support the view that money growth has valuable information content in signalling changes in inflation regimes, but the strength of such signal varies across countries and it is difficult to assess, mainly due to the limited number of regime shifts observed throughout the data sample This, in turn, translates into high uncertainty in the estimates of the parameters of the transition probability equation. Nonetheless, after performing a few robustness exercises, we are able to draw the following conclusions. First, model estimations conducted with information up to the end of 2013 show that the risks of a low inflation regime have been increasing in the course of the last six quarters of the sample in the euro area, while have remained substantially unchanged in the US and even declined in Japan; second, preliminary results of a model comparison exercise have shown that in the euro area money growth seems to play a stronger role in the assessment of deflation risks relatively to credit growth: in particular, model specifications employing credit aggregates as warning indicators return weaker signals of the risk of euro area inflation entering the Low inflation regime. Finally, evidence for Japan and US indicates that the inclusion of an indicator variable does not substantially change the assessment of the risk of entering a low inflation regime in comparison with a simple Markov switching specification with fixed transition probabilities.

A Data

A.1 Euro area

From 1992Q1 to 2012Q3 we employed official euro area data from the European Central Bank (ECB) website. For the period before 1992 we used German data. In particular, the Bundesbank real time data base was the source for data spanning from 1962Q1 to 1998Q4 while earlier (yearly) data were obtained from Buba 1988 and interpolated using the BFL interpolation procedure to obtain quarterly data. Data on loans to private sector are only available from 1980Q1 and provided by the ECB.

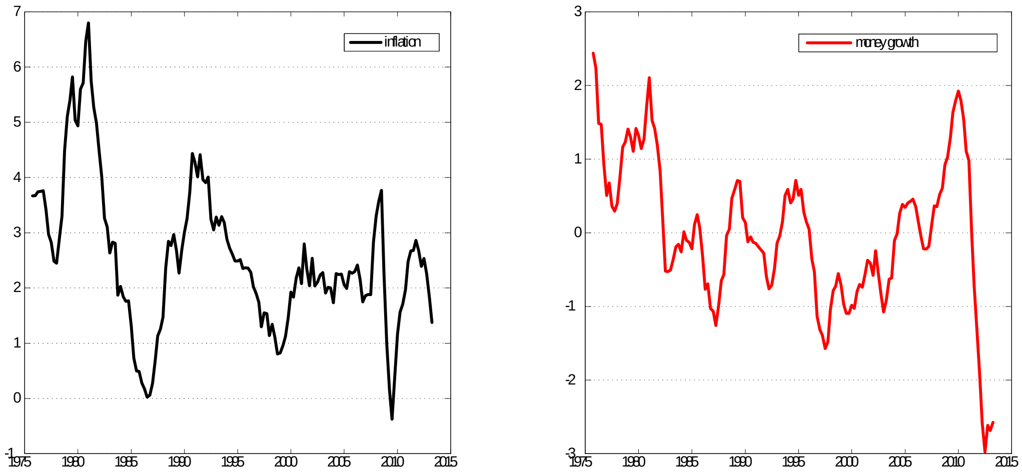
A.2 Japan

The sources for the CPI index and for the money supply (M2) series, both starting in 1957Q1, are the Japanese Statistics Bureau and the Bank of Japan, respectively. Data on credit to private non-financial sector (from all sectors) start in 1964Q4 and are provided by the Bank for International Settlements.

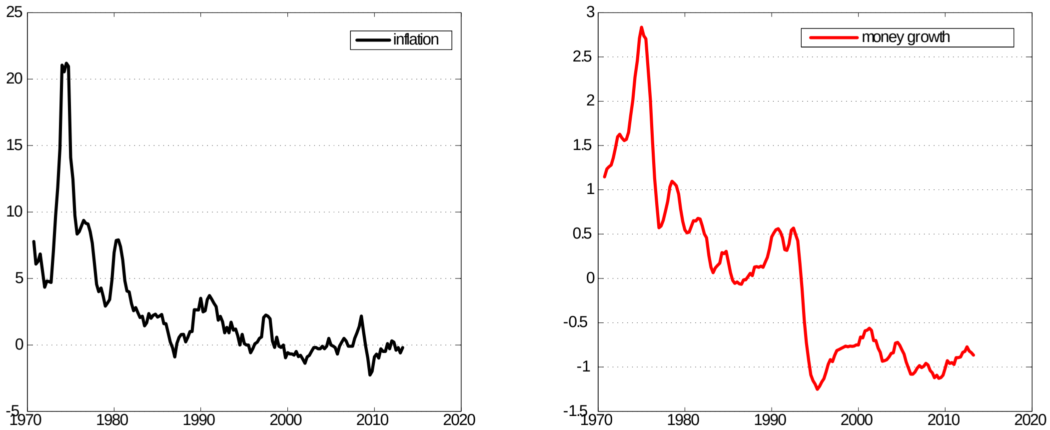
A.3 US

The source of the data is the Federal Reserve Economic Data provided by the Federal Reserve Bank of St. Louis. Quarterly data on the monetary aggregate (M2) are available since 1950Q1; therefore we used yearly data on M2 from Mitchell and Ame (1998), subject to quarterly BFL interpolation. Data on credit to private non-financial sector (from all sectors) start in 1952Q1 and are provided by the Bank for International Settlements.

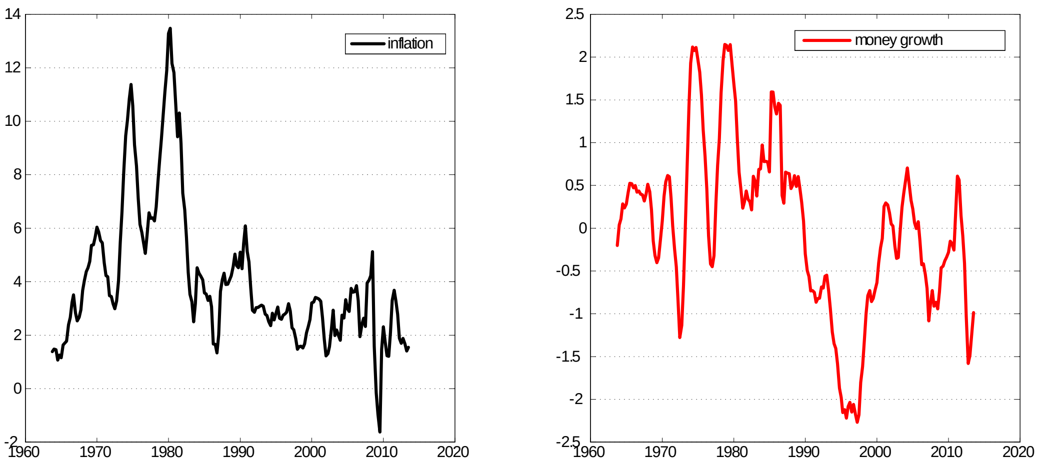
Figure A.1: Inflation and money growth indicator



Euro area data, last observation: 2013Q2



Japan data, last observation: 2013Q3



US data, last observation: 2013Q4

B Model estimation

B.1 Prior specification

Priors on the ordered probit specification for regimes are Gaussian centered on the values described in Section (2) and prior standard deviations set to 0.5. The slope coefficient on money growth is centered on 0.2 and with prior standard deviation equal to 0.5. The autoregressive coefficient is Gaussian with mean equal to 0.90 and standard deviation equal to 0.5. The state-specific means are endowed with a weak prior for the medium and high inflation states, respectively centered on 1.95% and 6% and with prior standard deviations of 0.3 and 1.0 respectively. The low inflation mean is calibrated at 0.50%.

The regime-specific standard deviations of shocks are given an inverse Gamma distribution centered on the standard deviation of residuals obtained by estimating a linear AR model for inflation data prior to the sample used for estimation, with a degrees of freedom parameter set equal to 4. This is a weakly informative prior.

Several robustness exercises have been conducted to assess the role of the priors and the results have been qualitatively confirmed for a range of alternative prior specifications.

B.2 Simulation strategy

Let us call $\mathbf{s}$ the $(T \times 1)$ vector with discrete states s_t, t =1,2,...T; $\mathbf{s}^*$ the $(T \times 1)$ vector with continuous latent variable s_t^*, t =1,2,...T; $\mathbf{y}$ the $(T \times 1)$ vector with observations on endogenous variable y_t, t =1,2,...T; $\mathbf{z}$ the $(T \times 1)$ vector with observations on indicator variables $\mathbf{z}_t$, t =1,2,...T; θ the vector with all parameters in the model defined as

$$\theta = \begin{bmatrix} \theta_{me} \\ \theta_{se} \end{bmatrix},$$

where θ_{me} contains all free parameters in the measurent equation (the 3-regime AR(1) specification for inflation), and θ_{se} contains all parameters appearing in the state equation (the ordered probit specification for transition probabilities). Then the Gibbs-data augmentation strategy is as follows:

- draw from $p(\mathbf{s}, \mathbf{s}^*|\mathbf{y}, \mathbf{z}, \theta)$ using the sequential partition:

$$p(\mathbf{s}, \mathbf{s}^*|\mathbf{y}, \mathbf{z}, \theta) = p(\mathbf{s}|\mathbf{y}, \mathbf{z}, \theta) \times p(\mathbf{s}^*|\mathbf{s}, \mathbf{y}, \mathbf{z}, \theta)$$

  (the first factor drawn using filtering and simulation smoother on discrete states and the second factor by drawing from truncated distributions the shocks in the state equation)

- draw from $p(\theta|\mathbf{s}, \mathbf{s}^*, \mathbf{y}, \mathbf{z})$ as follows (rem that conditioning on relevant latent variables the parameters in the measurement and state equations are independent, since shocks in these are independent)

$$\begin{aligned} p(\theta|\mathbf{s}, \mathbf{s}^*, \mathbf{y}, \mathbf{z}) &= p(\theta_{me}|\mathbf{s}, \mathbf{s}^*, \mathbf{y}, \mathbf{z}) \times p(\theta_{se}|\mathbf{s}, \mathbf{s}^*, \mathbf{y}, \mathbf{z}) \\ p(\theta_{me}|\mathbf{s}, \mathbf{s}^*, \mathbf{y}, \mathbf{z}) &= p(\theta_{me}|\mathbf{s}, \mathbf{y}) \\ p(\theta_{se}|\mathbf{s}, \mathbf{s}^*, \mathbf{y}, \mathbf{z}) &= p(\theta_{se}|\mathbf{s}^*, \mathbf{z}) \end{aligned}$$

These two steps can be performed by using the same algorithm as in Amisano and Fagan (2013).

B.2.1 Drawing parameters

Drawing from the posterior distribution of the parameters in the autoregressive equation is straightforward. For additional details we refer to Amisano and Fagan (2013). Drawing the β parameters in the ordered probit specification for transition probabilities can be done just by trivial extension of the algorithm used in Amisano and Fagan (2013). As for the threshold parameters γ, they have a conditional posterior distribution which is uniform over a domain defined by draws on the continuous latent variables, as described in Albert and Chib (1993), and therefore they can be easily simulated.

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Table 1: Data used in the application

country	*source*	*end date*	*money*	*inflation*	*start date*	*credit*	*start date*	*lag order*	*MA order*
Euro area	ECB, Bundesbank	2013Q2	M3	HICP	1962Q1	loans to priv. non-fin. sector	1980Q1	9	5
Japan	BoJ, BIS	2013Q3	M2	CPI	1957Q1	loans to priv. non-fin. sector	1964Q1	9	5
US	FRED, REF, BIS	2013Q4	M2	CPI	1950Q1	loans to priv. non-fin. sector	1952Q1	9	5

The effective sample start indicates the first observation of the sample after performing the necessary adjustments of the monetary aggregate.

Table 2: Posterior simulation, Euro Area

	prior mean	prior sd	prior low q	prior up q	post mean	post sd	post low q	post up q
ρ	0.90	0.50	-0.10	1.90	0.88	0.03	0.81	0.94
μ_1	0.50	0.00	0.50	0.50	0.50	0.00	0.50	0.50
μ_2	1.95	0.30	1.36	2.54	1.89	0.27	1.39	2.43
μ_3	6.00	1.00	4.07	7.94	6.16	0.83	4.50	7.77
σ_1	0.56	0.31	0.27	1.30	0.58	0.19	0.28	1.02
σ_2	0.56	0.29	0.27	1.29	0.27	0.03	0.23	0.33
σ_3	0.56	0.29	0.27	1.28	0.44	0.10	0.30	0.69
β_0	0.20	0.50	-0.78	1.19	0.23	0.19	-0.13	0.62
β_1	-0.85	0.50	-1.84	0.14	-0.85	0.40	-1.62	-0.03
β_2	2.04	0.50	1.07	3.02	1.95	0.26	1.43	2.43
β_3	2.57	0.50	1.60	3.55	2.25	0.36	1.57	2.97
γ_{21}	1.50	0.12	1.31	1.69	1.50	0.12	1.31	1.69
γ_{22}	3.45	0.14	3.21	3.69	3.50	0.14	3.23	3.69
γ_{23}	1.50	0.12	1.31	1.69	1.50	0.12	1.31	1.69

z_t=standardized MA(5) of lagged (9) M3 growth; estimation sample: 1962Q1-2013Q2

Table 3: Posterior simulation, Japan

	prior mean	prior sd	prior low q	prior up q	post mean	post sd	post low q	post up q
ρ	0.90	0.51	-0.08	1.91	0.82	0.08	0.67	0.97
μ_1	0.01	1.01	-1.94	1.99	-0.30	0.54	-1.32	0.93
μ_2	1.95	0.30	1.36	2.54	1.90	0.30	1.26	2.43
μ_3	5.98	0.99	4.04	7.90	6.53	1.01	4.50	8.46
σ_1	0.56	0.30	0.27	1.29	0.47	0.10	0.31	0.75
σ_2	0.56	0.27	0.27	1.27	0.53	0.11	0.38	0.81
σ_3	0.56	0.32	0.27	1.31	0.80	0.26	0.36	1.38
β_0	0.20	0.50	-0.79	1.18	0.45	0.30	-0.18	1.01
β_1	-0.84	0.49	-1.81	0.10	-1.11	0.39	-1.86	-0.30
β_2	2.04	0.50	1.05	3.03	1.51	0.43	0.74	2.46
β_3	2.57	0.50	1.59	3.54	2.24	0.47	1.34	3.20
γ_{21}	2.00	0.58	1.05	2.95	2.04	0.57	1.06	2.96
γ_{22}	3.53	1.44	1.13	5.88	4.27	0.86	2.46	5.82
γ_{23}	1.00	0.58	0.05	1.94	1.42	0.47	0.24	1.98

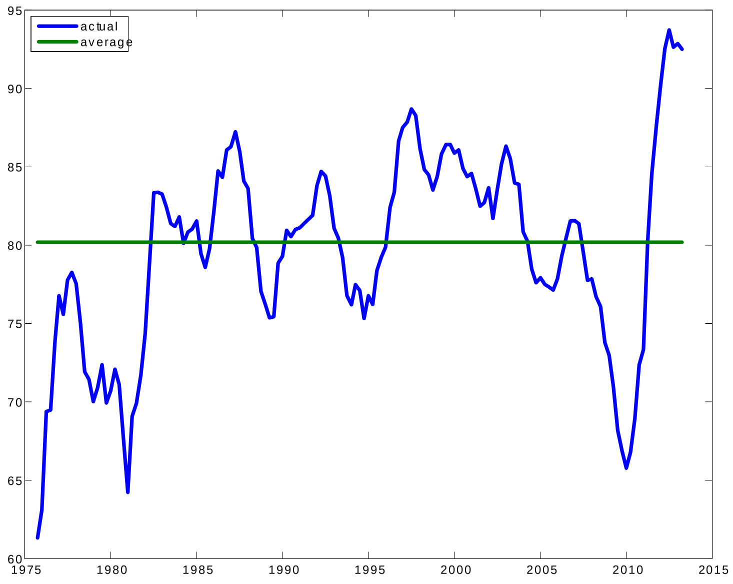
z_t=standardized MA(5) of lagged (9) M2 growth; estimation sample: 1964Q4-2013Q3

Table 4: Posterior simulation, US

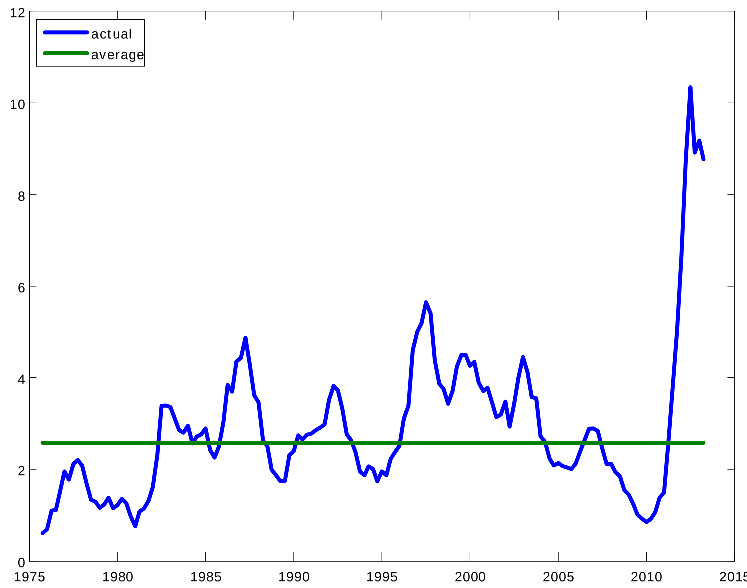
	prior mean	prior sd	prior low q	prior up q	post mean	post sd	post low q	post up q
ρ	0.89	0.50	-0.08	1.85	0.96	0.04	0.86	1.01
μ_1	0.50	0.00	0.50	0.50	0.50	0.00	0.50	0.50
μ_2	1.95	0.30	1.37	2.52	1.98	0.30	1.39	2.56
μ_3	6.00	0.98	4.06	7.96	6.29	1.13	4.21	8.64
σ_1	0.56	0.30	0.27	1.28	0.74	0.54	0.24	1.93
σ_2	0.56	0.28	0.27	1.27	0.51	0.20	0.29	1.06
σ_3	0.56	0.30	0.27	1.30	0.83	0.29	0.28	1.32
β_0	0.20	0.50	-0.77	1.18	0.29	0.31	-0.49	0.80
β_1	-0.85	0.50	-1.82	0.13	-0.98	0.41	-1.75	-0.10
β_2	2.05	0.50	1.07	3.02	2.00	0.41	1.22	2.82
β_3	2.56	0.50	1.59	3.52	2.59	0.38	1.83	3.32
γ_{21}	1.50	0.12	1.31	1.69	1.50	0.12	1.31	1.69
γ_{22}	3.45	0.14	3.21	3.69	3.46	0.14	3.22	3.69
γ_{23}	1.50	0.12	1.31	1.69	1.50	0.11	1.32	1.69

z_t=standardized MA(5) of lagged (9) M2 growth; estimation sample: 1952Q1-2013Q4

Figure 1: Euro Area: Time-varying transition probabilities



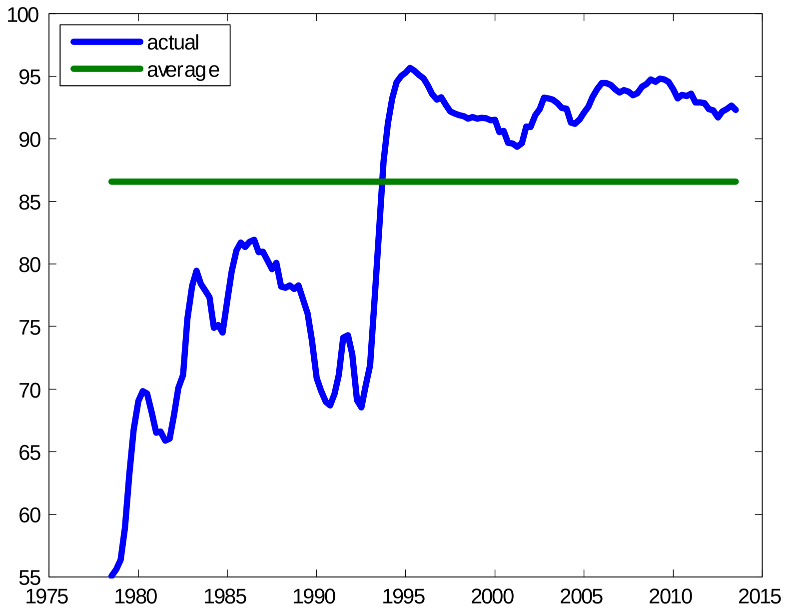
probability to stay in regime 1 when starting from regime 1



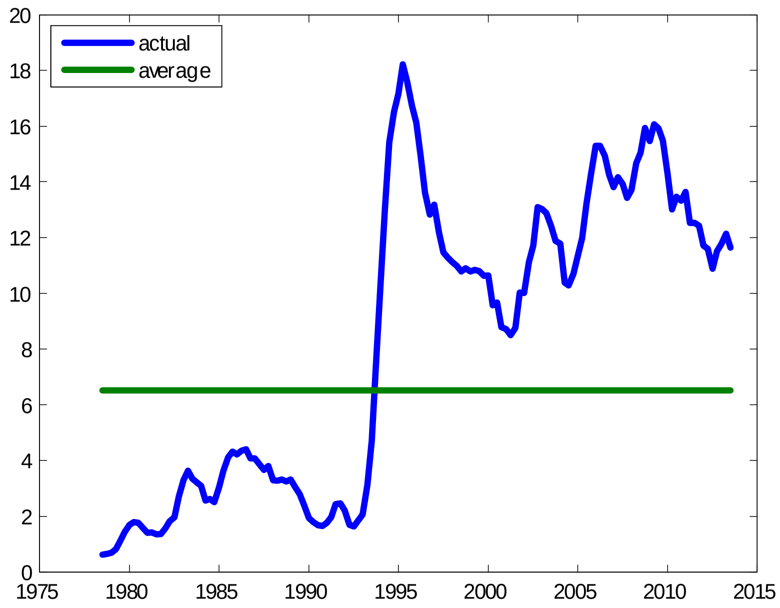
probability to move to regime 1 when starting from regime 2

Note: transition probabilities are computed at the posterior mean of parameter estimates and using actual money growth.

Figure 2: Japan: Time-varying transition probabilities



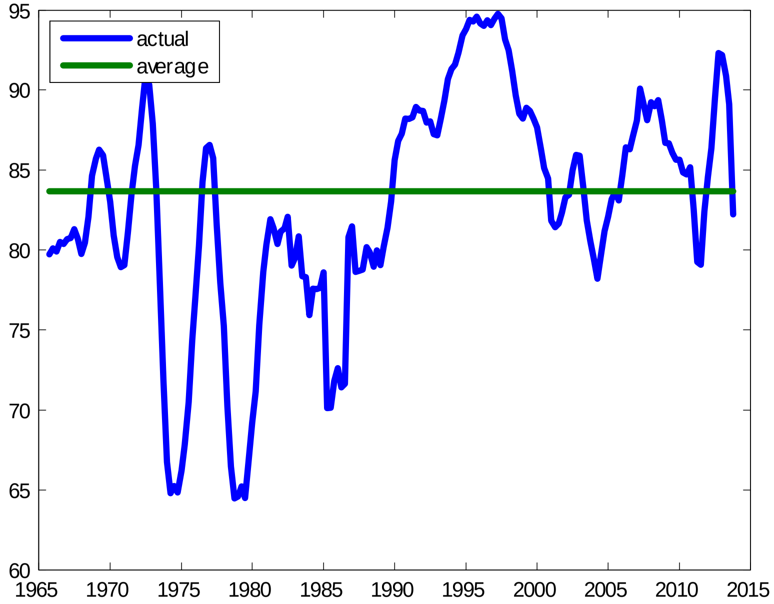
probability to stay in regime 1 when starting from regime 1



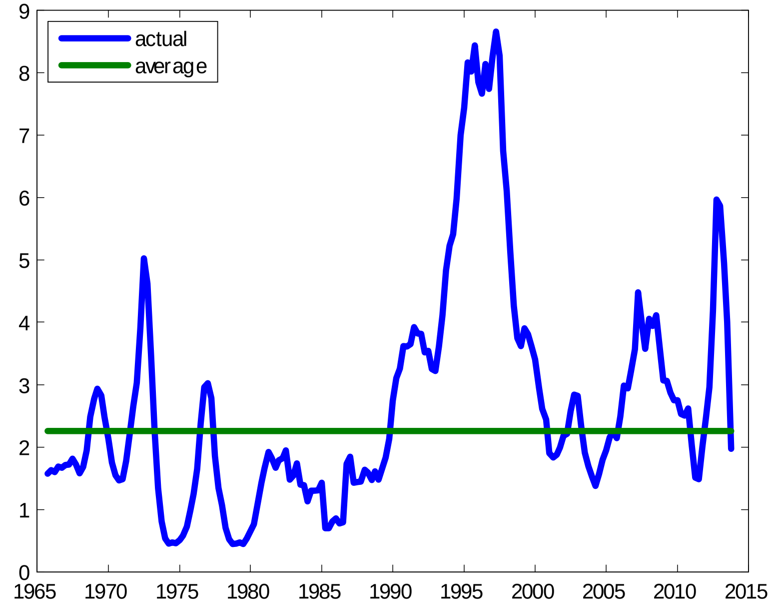
probability to move to regime 1 when starting from regime 2

Note: transition probabilities are computed at the posterior mean of parameter estimates and using actual money growth.

Figure 3: US: Time-varying transition probabilities



probability to stay in regime 1 when starting from regime 1



probability to move to regime 1 when starting from regime 2

Note: transition probabilities are computed at the posterior mean of parameter estimates and using actual money growth.

Table 5: Euro Area: Transition probabilities, smoothed probabilities and one step ahead probabilities computed at 2013Q2

Panel A: transition probabilities as determined by current monetary conditions				
	st+1 = Low	st+1 = Medium	st+1 = High	smoothed
st = Low	92.5	7.3	0.2	45.1
st = Medium	8.8	89.6	1.6	54.1
st = High	4.9	38.9	56.2	0.8
one step ahead	46.5	52.1	1.4	

Panel B: transition probabilities computed at "average" monetary conditions			
	st+1 = Low	st+1 = Medium	st+1 = High
st = Low	80.2	18.9	1.0
st = Medium	2.6	91.4	6.0
st = High	1.2	21.5	77.3
one step ahead	37.6	58.1	4.3

Note: Panel A reports transition probabilities, smoothed probabilities and one step ahead probabilities computed at the end of the sample used for estimation (1975Q1-2013Q2) and using actual monetary conditions. Panel B reports the corresponding probabilities computed using neutral values for monetary conditions.

Table 6: Euro Area: Transition probabilities, smoothed probabilities and one step ahead probabilities computed at 2013Q2

Panel A: transition probabilities as determined by current credit conditions				
	st+1 = Low	st+1 = Medium	st+1 = High	smoothed
st = Low	84.5	14.9	0.6	28.8
st = Medium	4.4	91.9	3.7	62.8
st = High	1.7	25.5	72.8	8.3
one step ahead	27.3	64.2	8.6	

Panel B: transition probabilities computed at "average" credit conditions			
	st+1 = Low	st+1 = Medium	st+1 = High
st = Low	79.3	19.6	1.0
st = Medium	2.9	91.6	5.6
st = High	1.0	21.1	78.9
one step ahead	24.8	64.9	10.4

Note: Panel A reports transition probabilities, smoothed probabilities and one step ahead probabilities computed at the end of the sample used for estimation (1983Q1-2013Q2) and using actual credit conditions. Panel B reports the corresponding probabilities computed using neutral values for credit conditions.

Table 7: Japan: Transition probabilities, smoothed probabilities and one step ahead probabilities computed at 2013Q3

Panel A: transition probabilities as determined by current monetary conditions				
	st+1 = Low	st+1 = Medium	st+1 = High	smoothed
st = Low	92.3	7.7	0.0	35.6
st = Medium	11.6	88.3	0.1	61.6
st = High	2.7	28.0	69.3	2.7
one step ahead	40.1	57.9	2.0	

Panel B: transition probabilities computed at "average" monetary conditions			
	st+1 = Low	st+1 = Medium	st+1 = High
st = Low	86.6	13.4	0.1
st = Medium	6.5	93.2	0.3
st = High	1.2	19.3	79.5
one step ahead	34.9	62.7	2.3

Note: Panel A reports transition probabilities, smoothed probabilities and one step ahead probabilities computed at the end of the sample used for estimation (1964Q4-2013Q3) and using actual monetary conditions. Panel B reports the corresponding probabilities computed using neutral values for monetary conditions.

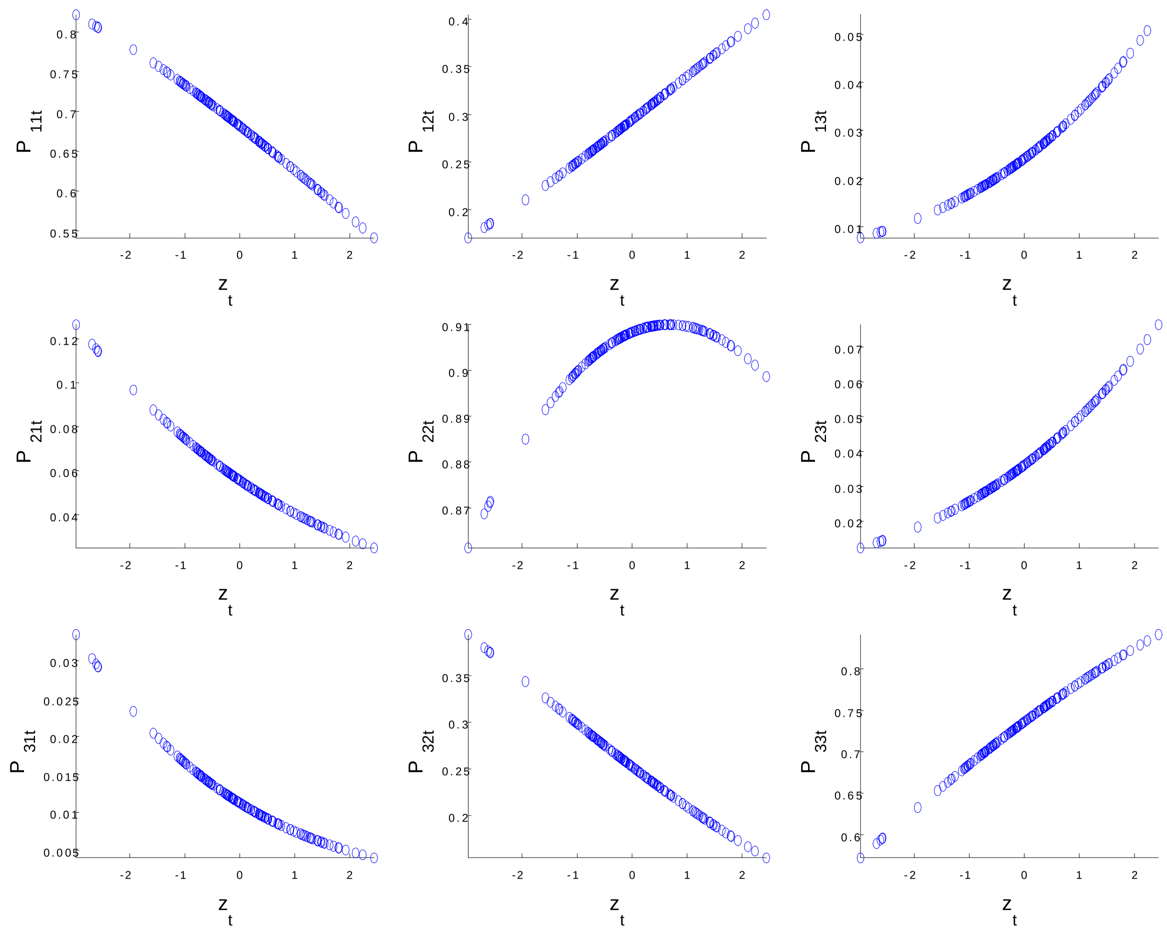
Table 8: US: Transition probabilities, smoothed probabilities and one step ahead probabilities computed at 2013Q4

Panel A: transition probabilities as determined by current monetary conditions				
	st+1 = Low	st+1 = Medium	st+1 = High	smoothed
st = Low	82.2	17.0	0.8	31.7
st = Medium	2.0	90.0	8.0	56.0
st = High	0.4	12.0	87.6	12.3
one step ahead	27.2	57.3	15.5	

Panel B: transition probabilities computed at "average" monetary conditions			
	st+1 = Low	st+1 = Medium	st+1 = High
st = Low	83.7	15.7	0.7
st = Medium	2.3	90.5	7.2
st = High	0.5	13.2	86.3
one step ahead	27.9	57.3	14.9

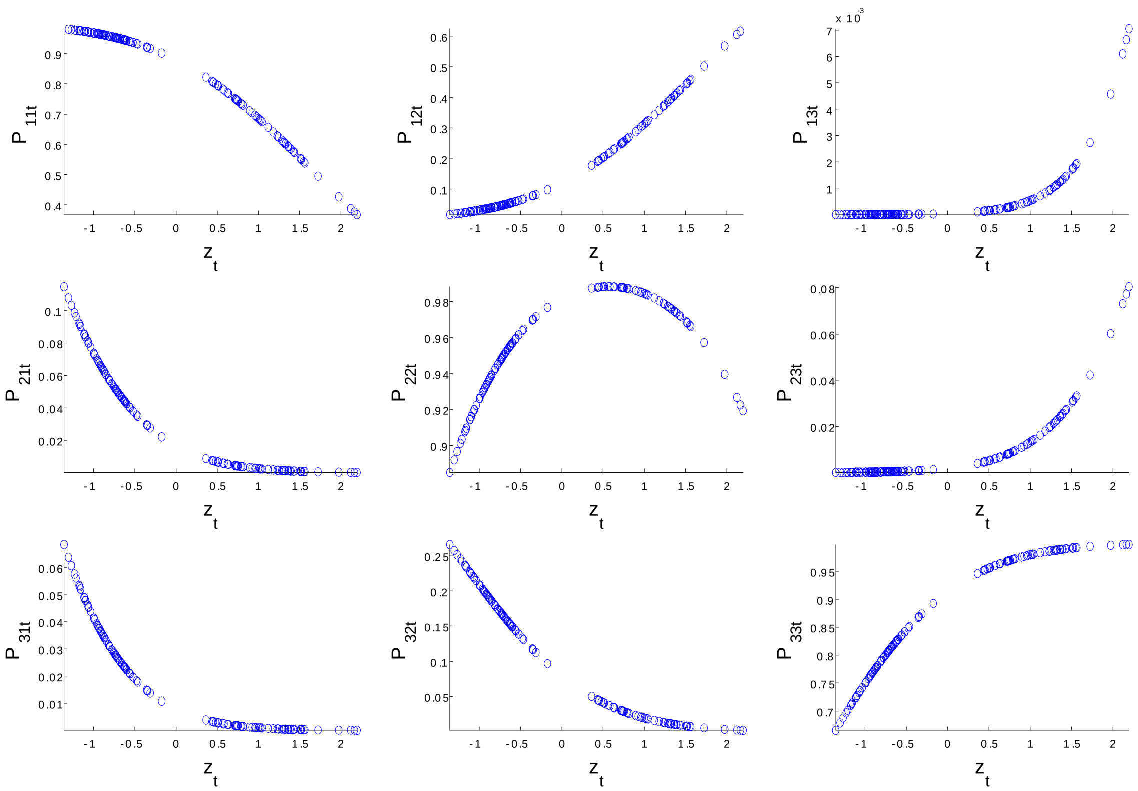
Note: Panel A reports transition probabilities, smoothed probabilities and one step ahead probabilities computed at the end of the sample used for estimation (1952Q1-2013Q4) and using actual monetary conditions. Panel B reports the corresponding probabilities computed using neutral values for monetary conditions.

Figure 4: Time-varying transition probabilities as function of the indicator variable, euro area



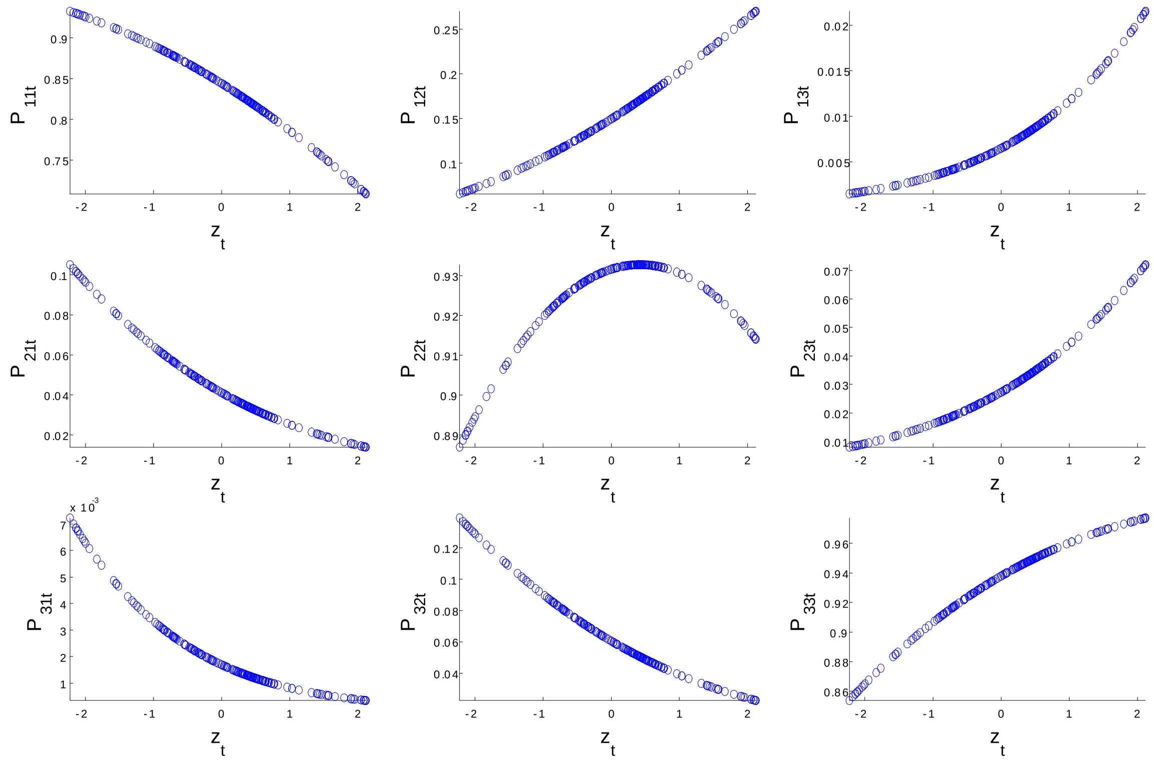
Time-varying transition probabilities are computed using the posterior mean values of the parameters.

Figure 5: Time-varying transition probabilities as function of the indicator variable, Japan



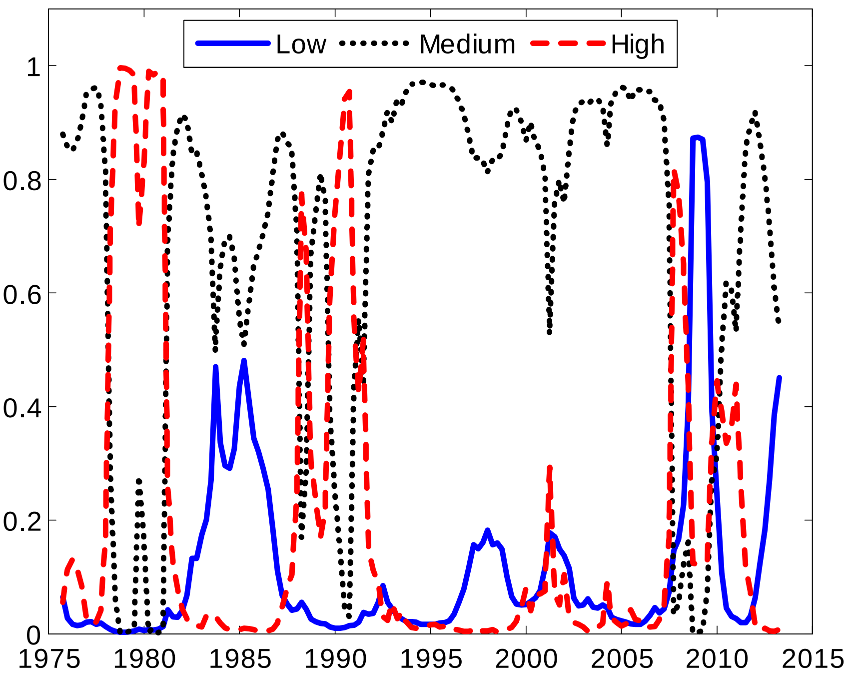
Time-varying transition probabilities are computed using the posterior mean values of the parameters.

Figure 6: Time-varying transition probabilities as function of the indicator variable, US

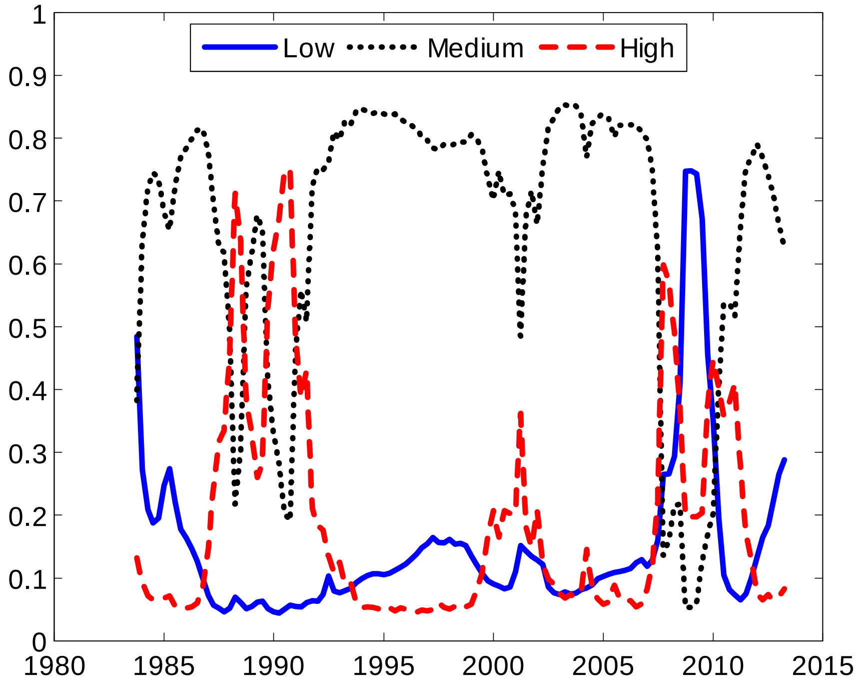


Time-varying transition probabilities are computed using the posterior mean values of the parameters.

Figure 7: Smoothed probabilities for inflation regimes, Euro Area

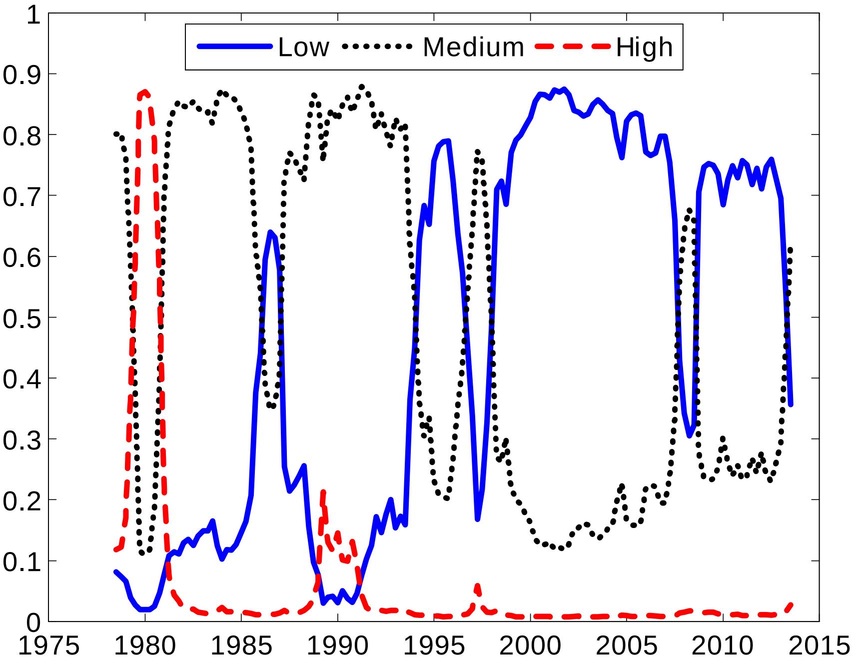


Estimated smoothed prob obtained with z_t=M3 growth (1962Q1-2013Q2)

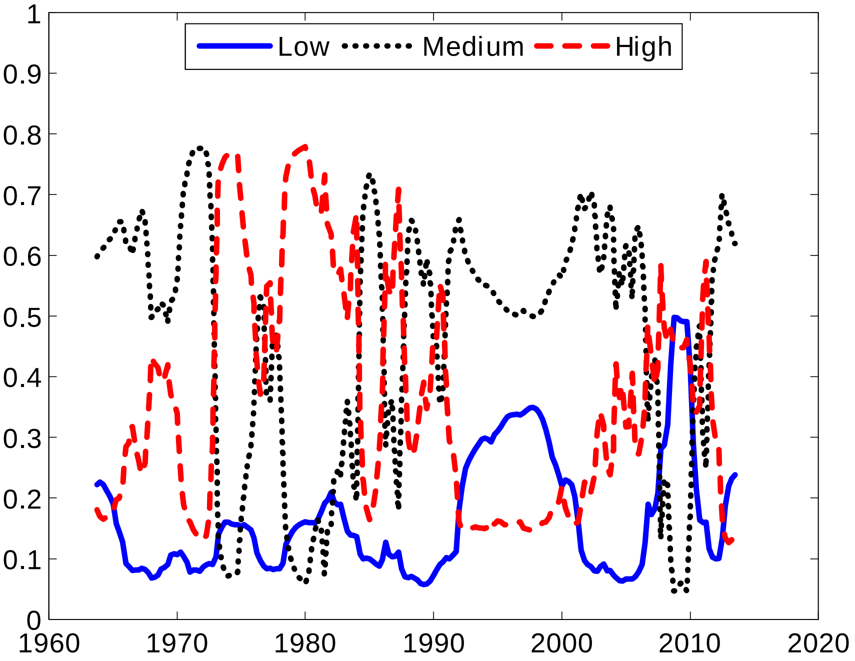


Estimated smoothed prob obtained with z_t=credit growth (1980Q1-2103Q2)

Figure 8: Smoothed probabilities for inflation regimes, Japan and US



Japan: Estimated smoothed prob obtained with z_t=M2 growth (1964Q4-2013Q3)



US: Estimated smoothed prob obtained with z_t=M2 growth (1952Q1-2013Q4)