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The yield spread and GDP growth – time varying leading properties and the role of monetary policy

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Abstract

The yield spread is a well documented leading indicator of GDP growth. Estrella (2005) proposes a model to explain this relationship. Within the model, the leading properties of the yield spread are determined by the monetary policy. Accordingly, changes of the leading properties that have been reported in many studies should correspond to changes of the monetary policy. This paper analyzes whether and what form of time variation of the leading properties can be found in four major industrialized countries (France, Germany, the UK and the US). The results are connected with time varying behavior of the monetary policy by modeling a joint state dependency of the leading properties and the reaction parameters of the monetary policy. Time variation of the leading properties seem to exist in all countries under consideration. For the US and Germany they are best modeled as a structural break while France and the UK exhibit recurring phases. Evidence for a link between the time variations of the monetary policy and the leading properties can be found. However, a clear determination of the leading properties by the monetary policy cannot be confirmed.

JEL classification: C32, E37, E43, E52

Keywords: leading indicator; yield spread; GDP growth; monetary policy; Markov-Switching

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1 Introduction

The yield spread, the difference between long-term and short-term interest rates, has been reported as a leading indicator of real output and recessions in different industrialized countries by several authors: e.g. Estrella and Hadrouvelis (1991), Plosser and Rouwenhorst (1994), Estrella and Mishkin (1997). The predictive power of the yield spread differs from country to country, but can be identified even in relatively small economies such as Belgium or the Netherlands, see Bernard and Gerlach (1998).

The leading properties between yield spread and real activity are not time-stable. Haubrich and Dombrosky (1996) report a worse fit of a simple prediction model for US real GDP growth using the yield spread in years between 1985 and 1995 than in the decades before. The case of structural stability has been addressed in the studies of Estrella et al. (2003), Chauvet and Potter (2002) and Venetis et al. (2003). Using monthly data up to 1998 for Germany and the US, Estrella et al. (2003) find structural breaks in the relation between yield spread and the US industrial production, but no breaks for Germany and within a prediction model for US recessions. The latter result has been contradicted by Chauvet and Potter (2002) using Bayesian methods. Applying time varying smooth transition models, Venetis et al. (2003) find time dependence for the US, Canada and the UK, considering GDP growth up to the year 2000. In addition to time dependency Venetis et al. find threshold effects for the mentioned countries. Threshold behavior has also been reported earlier in Galbraith and Tkacz (2000) for Canada and the US and in Duarte et al. (2005) for the Euro zone.

A theoretical model recently proposed by Estrella (2005) gives a formal explanation why the yield spread is a leading indicator of GDP growth. The model is based on the expectation hypothesis that long term interest rates reflect expectations on future short term interest rates. The model implies that the leading properties of the yield spread are determined by the behavior of the central bank regarding inflation and especially the output gap. In line with the assumption that expectations on long term interest rates drive the leading properties are the results of the analysis of Hamilton and Kim (2002) who decompose the yield spread in interest rate expectations and a term premium. They find only the interest expectations as a significant leading indicator. However, Favero et al. (2002) find both parts significant using a different decomposition. Following the hypothesis, that the monetary policy determines the leading properties of the yield spread, the observed time variation of the leading properties should correspond to time variations of the central bank behavior.

This paper aims at examining the structural stability of the relationship between the yield spread and future real activity not only by allowing for one single transition, but also by modeling state dependent leading properties that allow for recurring phases. Discrete states in form of a Markov-Switching model as well as continuous states in form of an autoregressive time-varying coefficient model are considered. Four industrialized countries are dealt with in this paper: the US, the UK, France and Germany. Moreover, the paper links the results on state dependency to the theoretical model of Estrella (2005) by also considering different states and implicit changes of the monetary policy.

To capture changes in monetary policy two approaches are used in this paper. Clarida et al. (2000) propose that institutional changes of the central banks, like the assumption of office of a new chairman or the introduction of a currency union, correspond to structural instability of the reaction parameters of monetary policy rules. Accordingly, in the first approach sub-samples are formed according to major institutional changes. The leading properties are compared between these sub-samples. This approach has also been chosen by Estrella (2005) to prove the theoretical model. The second approach used in this analysis replaces the formation of sub-samples by a data driven method. The different monetary policy regimes shall be captured by the different states of a Markov-Switching model, compare Sims (1999), Wesche (2003) or Bueno (2004) who apply Markov-Switching processes to model the time variation of the behavior of central banks. If the time-variation of the monetary policy governs the time-variation of the relation between yield spread and real activity, the states identified within the Markov-Switching model for the monetary policy rule shall be similar to the states of a Markov-Switching model for the relation between yield spread and real activity. This hypothesis is assessed in this analysis via estimation of a joint Markov-Switching model for the leading relationship and the monetary reaction function. Within this model both, reaction parameters of the monetary policy as well as the leading properties are assumed to follow the same latent Markov-Switching process.

This paper contributes to the literature in two aspects. Firstly, time varying leading behavior of the yield spread with respect to GDP growth is assessed by explicitly modeling state dependency in form of a Markov-Switching process and an autoregressive process and forms an extension to the existing literature that employs models with a single transition in time only. Secondly, the state dependency is connected to states of the monetary policy using Markov-Switching monetary policy rules. The identification of Markov-Switching states is used to assess the theoretical results of the model of Estrella (2005) that different states of the monetary policy imply different states for the leading properties of the yield spread and that the relation between expectations of real activity and the monetary policy reaction on the output gap is inverse. The joint Markov-Switching model for the monetary policy rule and the leading relationship provides an alternative approach to evaluate theoretical considerations concerning the relation between monetary policy and the predictive power of the yield spread than the comparison of sub-samples. The analysis can be paraphrased by three questions:

I. Do time varying leading properties of the yield spread exist?

II. Do the leading properties and the monetary policy share the same state dependency?

III. Can the results be interpreted in the framework of the theoretical model of Estrella (2005)?

The outline of the paper is as follows. The data sets are described in Section 2. Section 3 discusses different models to capture the time variation of the leading properties of the yield spread concerning GDP growth. In Section 4 the first approach to connect the time variation with monetary policy changes, namely the comparison of sub-samples, is presented. Section 5 introduces the joint Markov-Switching model and compares its fit to alternative models without joint states replying to the Question II and III. Section 6 concludes.

2 Data description

The analysis deals with four industrialized countries: France, Germany, the United Kingdom and the Unites States. The data is taken from the OECD database. For all countries the yield spread is represented by the difference between three months interbank rates and interest rates of governmental papers with 10 years maturity. For France and Germany the national three month rates (Pibor and Fibor) are continued from the first quarter in 1999 onwards by the Euribor rates due to the European currency union. For the same reasons the central bank rates of Germany and France are continued by the interest rates of the ECB from the first quarter in 1999 onwards.[1] The growth rates of GDP are calculated on the basis of seasonal adjusted values. In the case of Germany up to 1991 only West German GDP data is available. This series is directly followed by GDP data for the reunified country without overlapping. In order to obtain a single series in a first step quarterly growth rates are calculated; the growth rate at the matching point is replaced by the mean of the 10 surrounding growth rates. This series is used to count back a GDP series starting with the last observation of GDP. The resulting series can be used to obtain growth rates at horizons longer than one quarter. The resulting GDP series is used to calculate the output gap, too. This is done for all countries via

[1] Concerning the consideration of the European currency union for the estimation of the monetary policy rules see the discussion in Sub-Section 5.2.

the Hodrick-Prescott filter. As a measure of inflation the growth rates of the consumer price indices are used. Summary statistics for all variables are given in Table 1.

For all countries the time horizon of the analysis is between 1973:2 and 2002:4. The first sample point in time is chosen according to the end of the era of fixed exchange rates according to the Bretton Woods system and its direct successors. The end of fixed exchange rates is the starting point of independent monetary policies in countries aside from the US.

3 Time varying leading properties

First, it is analyzed whether and what type of time variation of the leading properties of the yield spread exists. In addition to time varying coefficient approaches based on latent state variables threshold effects are considered.

3.1 Modeling Time Variation

A simple linear prediction model for GDP growth which is widely used in the literature is given by:[2]

$$\Delta y_t^k = \alpha + \beta S_t + \varepsilon_t, \quad \varepsilon_t \overset{i.i.d.}{\sim} (0, \sigma^2), \tag{1}$$

where $\Delta y_t^k = (400/k) \cdot (\log y_{t+k} - \log y_t)$ is the k-step ahead growth rate of GDP. S_t denotes the spread between a long-term interest rate and short-term interest rate. This model reflects the relation between yield spread and real activity. The yield spread has predictive power for different horizons k ranging from two quarters to four or five years.[3] This analysis is restricted to a horizon of one year or $k = 4$.

Galbraith and Tkacz (2000) use threshold models to introduce non-linearity in the relationship between the yield spread and real activity.

$$\Delta y_t^k = \alpha_0 + \beta_0 S_t + (\alpha_1 + \beta_1 S_t) I_{(z_{t-d} > \lambda)} + \varepsilon_t, \quad \varepsilon_t \overset{i.i.d.}{\sim} (0, \sigma^2) \tag{2}$$

Given the threshold parameter λ, the model in Equation (2) is a switching regression model. If the threshold variable z_{t-d} is below λ, the regression is based on the parameters α_0 and β_0, otherwise on α_1 and β_1. Galbraith and Tkacz (2000) use the spread itself as the threshold variable and find

[2]See e.g. *Estrella and Hadrouvelis* (1991), *Plosser and Rouwenhorst* (1994), *Estrella and Mishkin* (1997), *Bernard and Gerlach* (1998).

[3]Compare the authors listed in Footnote 3.1.

it to be significant for the US and Canada. Duarte et al. (2005) apply several candidates and find lagged GDP growth as a threshold variable for the Euro zone.[4]

If the threshold variable z_{t-d} is equal to the time indicator t, this model is equivalent to the assumption of an unknown structural break in the parameters α and β. Structural breaks have been found for the US in 1984 by Estrella et al. (2003) regressing the industrial production on the yield spread and for the Euro zone in 1992 by Duarte et al. (2005). One property of structural break models is that changes in time are only possible once in one direction.

This assumption is rather restrictive. It implies that only one event changes the leading properties for the whole sample and that the former state is impossible to recur. Therefore time-varying coefficients are applied. Two different processes that govern the predictive relation between the yield spread and GDP growth are considered, on the one hand an autoregressive process, on the other a Markov-Switching process. The Markov-Switching model is given by

$$\Delta y_t^k = \alpha + \beta_{v_t} S_t + \varepsilon_t, \quad \varepsilon_t \overset{i.i.d.}{\sim} N(0, \sigma^2) \tag{3}$$

where v_t denotes the state variable that can take discrete values and evolves according to a Markov-chain. Assuming the number of states is N, the Markov-chain can be described by the following transition matrix:

$$P = \begin{bmatrix} p^{11} & p^{21} & \cdots & p^{N1} \\ p^{12} & p^{22} & \cdots & p^{N2} \\ \vdots & \vdots & \ddots & \vdots \\ p^{1N} & p^{2N} & \cdots & p^{NN} \end{bmatrix}.$$

The probability of transition to the state n conditional on being in state m is:

$$p^{nm} = P(v_t = n | v_{t-1} = m).$$

In the case of $N = 2$ a (single) structural break can be modeled by a restricted Markov-Switching model with non-recurring states where the process starts in state 1 and when state 2 is reached stays there. This corresponds to the restriction of $p^{22} = 1$.

In addition to the Markov-Switching model a linear time varying coefficient model is considered

[4] A generalization of the threshold model is the smooth transition model which should reflect not an abrupt but steady change of the parameters according to the change variable z_{t-d}. *Venetis et al.* (2003) apply a time-varying smooth transition model on the leading relationship between yield spread and GDP growth.

in which the β coefficient does not follow a discrete process but evolves in an autoregressive process

$$\begin{aligned} \Delta y_t^k &= \alpha + \beta_t S_t + \varepsilon_t, \quad \varepsilon_t \overset{i.i.d.}{\sim} N(0,\sigma^2) \\ \beta_t &= \mu_\beta + \phi\beta_{t-1} + e_t, \quad e_t \overset{i.i.d.}{\sim} N(0,\sigma_\beta^2). \end{aligned} \tag{4}$$

The autoregressive time varying coefficient model gives a the quite well suited method to picture the time variation of the leading properties by plotting the smoothed values of the latent β coefficient.

Both models can be enriched with a threshold process like described in Equation (2), such that the relationship between yield spread and GDP growth is governed by a latent state variable and an observable threshold variable. The consideration of an additional threshold variable shall control for the possibility that a threshold process can capture the time variation of the leading properties. Such a result suggests that rather than different states of the monetary policy, non-linear behavior of the economic agents causes the time variation. The following Equation gives the Markov-Switching model with a threshold:

$$\Delta y_t^k = \alpha_0 + \beta_{0,v_t} S_t + (\alpha_1 + \beta_{1,v_t} S_t) I_{(z_{t-d}>\lambda)} + \varepsilon_t, \quad \varepsilon_t \overset{i.i.d.}{\sim} N(0,\sigma^2). \tag{5}$$

Threshold behavior can also be considered within the autoregressive time varying coefficient model. Concerning the β coefficient, the threshold is modeled to evoke a level shift ($+\beta_1$) resulting in the following model:

$$\begin{aligned} \Delta y_t^k &= \alpha_0 + \beta_t S_t + (\alpha_1 + \beta_1 S_t) I_{(z_{t-d}>\lambda)} + \varepsilon_t, \quad \varepsilon_t \overset{i.i.d.}{\sim} N(0,\sigma^2) \\ \beta_t &= \mu_\beta + \phi\beta_{t-1} + e_t, \quad e_t \overset{i.i.d.}{\sim} N(0,\sigma_\beta^2). \end{aligned} \tag{6}$$

All models used in this analysis are estimated via Maximum Likelihood assuming Gaussian errors. For the autoregressive time varying coefficient model the Kalman filter is applied to obtain the likelihood. The model selection is done via the Akaike criterium (AIC) and the Schwarz criterium (BIC). The BIC is more conservative than the AIC and favors less parameterized models.

3.2 State Dependent Models: Results

The linear model in Equation (1) is chosen as benchmark model and is compared with the Markov-Switching and the autoregressive time varying coefficient models described in the foregoing subsection. Table 2 gives the results of the estimation of Equations (1), (3) and (4). The linear model is compared with the Markov-Switching model, a restricted Markov-Switching model (allowing for a single break) and an autoregressive time varying coefficient model. The number of states for the

Markov-Switching model are determined by a preliminary analysis regarding the AIC and BIC. For the UK, the model with three states yields the best AIC however the BIC is best for a two state model. For the other three countries, both criteria favor a two state model.

For the US, the best model of the set presented in Table 2 is the restricted Markov-Switching model according to the AIC and BIC. The ML estimate of the restricted model is also the ML estimate of the unrestricted indicating a structural break taking place in 1984:2, which is in line with Estrella et al. (2003) and Chauvet and Potter (2002).[5] After the break in 1984 the β coefficient turns out to be insignificant. The R^2 is slightly better for the autoregressive time varying coefficient model. This model itself is preferred to the linear one according to both criteria. In contrast to the structural break hypothesis the β coefficient regains significance in the second half of the nineties according to the autoregressive time varying coefficient model (see Figure 2, upper panel). This movement goes along with the observation that the yield curve is able to indicate the 2001 recession, see e.g. Stock and Watson (2003).

Also for the UK all time varying coefficient models are favored against the linear model according to both criteria. The unrestricted Markov-Switching model is preferred to its restricted counterpart. According the AIC, it is also preferred to the autoregressive time varying coefficient model. For the BIC the autoregressive time varying coefficient model is preferred. Both models indicate that two phases, one at the beginning of the eighties and the other at the beginning of the nineties, show up with high β coefficients (see Figure 1 and 2, second panel each). The corresponding ML estimates for the parameters of the Markov-Switching model are $\widehat{\beta}_1 = 2.0713$ and $\widehat{\beta}_2 = 0.2080$. According to the autoregressive time varying coefficient model the yield spread is an insignificant leading indicator for long phases.

For Germany the AIC prefers the time varying coefficient models over the linear model vice versa for the BIC. However the difference between the BIC of the linear model and the restricted Markov-Switching model is rather small. If one follows the hypothesis of a time varying β, the results favor a structural break in 1993:2. The AIC and BIC prefer the restricted Markov-Switching to the autoregressive time varying coefficient model; like in the US, the ML estimate of the restricted model is equivalent to the ML estimate of the unrestricted model. This break seems to be accompanied by a change in growth dynamics like in the US, see Aßmann et al. (2006).

In the case of France, a large likelihood gain is obtained by admitting for time varying coeffi-

[5] Interestingly, this date corresponds to the well documented reduction of business cycle volatility, see e.g. *Kim and Nelson* (1999).

cients. According to the AIC and the BIC, the autoregressive time varying coefficient model fits the data best. Phases with rather high value of the β coefficient are followed by phases in which the yield spread has no predictive power with an insignificant β coefficient, compare Figure 2. Three times the β coefficient becomes insignificant: 1979, 1993 and 2001. 1979 and 1993 can be linked specific developments of the European currency union: the introduction in the ECU in 1979 and the ratification of the Maastricht treaty in 1993, which was accompanied by a strict tightening of the French fiscal policy.

3.3 State Dependent Threshold Models: Results

Several authors find threshold effects in the leading relationship between yield spread and GDP growth. Preliminary analysis leads to the choice of lagged GDP growth as a threshold variable as found in Duarte et al. (2005). The inclusion of the threshold is meant to improve the fit of the model and to monitor whether time varying behavior can in fact be traced back to threshold effects. Table 3 gives the results of the models described in Equations (2), (5) and (6). Figure 3 shows the expected value of β_t according to the estimates of the autoregressive time varying coefficient model with threshold.

In the case of the US the time varying coefficient models are preferred to the linear model in the class of threshold models according to both AIC and BIC. The restricted Markov-Switching model is the best model with an implicit break point in 1984:2. The estimate of threshold value is -0.1014. It is the best model to fit the data according to the AIC; according to the BIC, it is its counterpart without a threshold (see Table 2).

For the UK, the linear model again has the highest values of both AIC and BIC. There is evidence for the hypothesis of one structural break against the linear model but not against the unrestricted Markov-Switching. The Markov-Switching model with the threshold is the best model according to the AIC. The estimated threshold value is 2.101 for all models. According to the BIC the autoregressive time varying coefficient model is better. The threshold specification is preferred for each model according to both criteria. Thus this relation is valid for all models considered in this section.

For the German data, the restricted Markov-Switching with a threshold is the best model according to the AIC. The estimated threshold value is 3.041. The implied break point is identical with the one of the model without a threshold in 1993:2. Again, the estimate of the unrestricted Markov-Switching model is equal to the estimate of the restricted. According to the BIC, the linear threshold model is still preferred; the linear model without a threshold is the best model (according to the AIC the restricted Markov-Switching without a threshold is the best model, see Table 2).

In the case of France, time varying coefficient models also dominate when a threshold is considered. According to both criteria the autoregressive time varying coefficient model with a threshold is the best model of all considered in Section 3. The plot of the β coefficient in time given in Figure 3 shows the same pattern as that of the time varying coefficient model without a threshold: phases with high values for β_t are followed by phases where β_t is near zero.

To sum up, strong evidence for time varying leading properties exists for the US, the UK and France. For the US, the time variation can be characterized as a structural break that causes the loss of the leading property of the yield spread. For the UK and France, recurring phases with high and low values for β can be found. In the case of Germany, slight evidence of a structural break exists.

For all countries and all models the threshold leads to an improvement according to the AIC. By the inclusion of a threshold the BIC is improved for all specifications in case of the UK. For the US and France, this is only true for the autoregressive time varying coefficient model. However the relations between the types of modeling time variation are not changed by the inclusion of a threshold. A threshold process is not the only source of the time variation of β.

4 Monetary policy and the leading properties of the yield spread

The previous section dealt with regime switching behavior of the predictive relations between the yield spread and real GDP growth. Evidence of at least two different states has been found for all countries. According to the theoretical model of Estrella (2005) different monetary regimes should correspond to different predictive relations between the yield spread and real GDP growth.[6] This model implies that the expectation of future real activity is a linear function in the yield spread. Furthermore, the precursive relation is inverse to the reaction of the central bank on deviations of real activity from its desired level. Within this model central bank behavior is characterized by a reaction function or monetary policy rule.

Two approaches are applied to analyze the link between monetary policy and the leading properties of the yield spread. They are also used to assess the implications of the model of Estrella

[6] The main equations of the model of *Estrella* (2005) as well as the deduced link between expectations on output and the yield spread are reported in Appendix A.

(2005).

1. Sub-samples are chosen according to points in time that possibly mark institutional changes of the central banks. Estimates of the sub-samples are then compared to find out whether they differ significantly. This approach has been applied before by Feroli (2004) and Estrella (2005) and is presented Sub-Section 4.2.

2. The second approach is to model a joint process for both the behavior of the central bank and the leading properties. This is done by assuming a joint Markov-Switching process explained in Section 5.

In the following sub-section the econometric model of a monetary policy rule according to Clarida et al. (2000) is introduced. The corresponding monetary policy rule can also be found in the theoretical model of Estrella (2005). The monetary policy rule is used to characterize the behavior of the central bank; it also is the basis of the joint Markov-Switching model in Sub-Section 5.

4.1 Monetary policy rules

Based on the assumption that the nominal interest rate i_t is the main instrument of central banks, Taylor (1993) proposes the following characterization of monetary policy:

$$i_t^* = \overline{r} + \overline{\pi} + \eta(\pi_t - \overline{\pi}) + \xi(y_t - \overline{y_t}). \tag{7}$$

The desired interest rate i_t^* is the sum of the equilibrium real rate $\overline{r}$, the inflation target $\overline{\pi}$ and the reaction to the deviations of inflation and output from their target values ($\overline{\pi}$ and $\overline{y}_t$) according to the coefficients η and ξ. The importance of the inflation and the business cycle goals is reflected by these coefficients. The inflation target is assumed to be invariant while the target output varies according to the variation of potential output.

This monetary policy rule heeds contemporaneous regressors and no further expectations. A forward looking policy rule can be established by replacing π_t and y_t by expected future values. Furthermore, it can be observed that interest rates are highly persistent. Central bank rates are changed in small steps only. Central banks seem to smooth interest rates. Such a behavior shall be captured by the parameter ρ that Clarida et al. (2000) include in the following policy rule

$$\begin{aligned} i_t^* &= \zeta + \eta(E[\pi_{t+j}] - \overline{\pi}) + \xi(E[y_{t+k}] - \overline{y_{t+k}}) & (8) \\ i_t &= (1-\rho)i_t^* + \rho i_{t-1}. & (9) \end{aligned}$$

Clarida et al. (2000) deduce the following linear model from Equations (8) and (9):

$$i_t = (1-\rho)[\zeta - \eta\overline{\pi} + \eta\pi_{t+j} + \xi x_{t+k}] + \rho i_{t-1} + \epsilon_t, \tag{10}$$

where x_{t+k} denotes the output gap $(y_{t+k} - \overline{y_{t+k}})$ and ϵ_t consists of the expectation errors concerning future inflation and output gap. Clarida et al. (2000) solve the endogeneity problem concerning the regressors π_{t+j} and x_{t+k} via instrumenting them and perform a GMM estimation.

4.2 Institutional change points

The hypothesis of the correspondence between states of monetary policy and states of the leading properties has been analyzed by Estrella (2005) and Feroli (2004) for the US. Both divide the sample into sub-samples, which are in line with the terms of office of Federal Reserve Board chairmen. This approach reflects the idea that different chairmen pursue different aims and impose different monetary policy regimes (rules). As a first step, such institutional change points are regarded within the linear model regressing the yield spread on GDP growth as changes of the β coefficient. With each institutional phase a "new" phase of the precursive relationship is assumed. For the US, two changes are considered, the change to Paul Volcker and the one to Alan Greenspan. Changes of the predecessors of Paul Volcker are neglected.

The great influence of the chairmen like in the US cannot be found in the other three countries. In the UK and France, central banking's political dependency did not end until the 1990s. Therefore, the beginning of the legislation of Margaret Thatcher is considered as a possible change point for the UK. Possible dates for France and Germany are the introduction of the European currency system in 1979 (ECU), which the UK entered in 1990 and left in 1992, and the introduction of the Euro in 1999. Furthermore, the ideas of monetarism influenced monetary politics in the mid seventies, e.g. Germany introduced the money supply management in 1975 and the Bank of England committed itself to monetarism in 1976. Finally, for Germany the monetary union between the Federal Republic of Germany and the German Democratic Republic is considered. An overview of all considered dates that can be assumed as institutional change points of monetary policy is given in Table 4.

The possible change points, $\{cp_n\}_{n=1}^{N}$, are taken into account via dummy variables

$$\Delta y_t^k = \alpha_0 + \beta_0 S_t + \sum_{n=1}^{N} \beta_n S_t I_{(t>cp_n)} + \varepsilon_t \ , \quad \varepsilon_t \sim N(0, \sigma^2). \tag{11}$$

The dummy variables establish sub-samples in which the β coefficient can take different values. Not all institutional change points considered lead to significantly different leading properties. Insignificant dummies or change points are eliminated stepwise. The resulting change points in monetary policy that deliver significantly different phases of the leading properties, are reported in Table 5. For all countries, evidence for at least one institutional change point compared with the linear model exists.

For the final sub-samples monetary policy rule according to Equation (10) are estimated to monitor the impact of the institutional change on the reaction parameters. Two versions of this policy rule are estimated for each sub-sample, one with contemporaneous regressors ($k = j = 0$) and one forward looking version ($k = j = 1$). The differences between the sub-sample estimates of the reaction parameters are given in Table 6, which also includes the differences of the sub-sample estimates of the β coefficients. Theory suggests that the differences of the β coefficients are accompanied with differences of the ξ coefficients with contrary sign.

Moreover, the fit of the final change point models is compared to optimal break points, see Table 5. Optimal break points are obtained via sequential estimation of the linear model for the sub-samples for any possible break point. The break point (or the set of break points) with the highest likelihood is chosen as optimal break point. The number of break points that are chosen corresponds to the number of significant institutional change points. The optimal break points, their likelihoods and values for the model selection criteria are given in Table 5.

In case of the US, the beginning of the era of Alan Greenspan in 1987 is the best institutional change point according to the model selection criteria. Two change points exhibit worse AIC and BIC. On the other hand, a model with no change point is rejected by both criteria, too. Compared to the 1979 change point connected with Paul Volcker the change point in 1987 exhibits a far better likelihood value. Before the break $\widehat{\beta}_1 = 0.9139$ and the corresponding estimator $\widehat{\beta}_2 = 0.1506$ and is insignificant. However, the change of the chairman does not come along with a break of the reaction parameters. The differences $\xi_1 - \xi_2$ and $\eta_1 - \eta_2$ are both insignificant and of the wrong sign compared to the suggestions of the theory. This is true for both, the forward looking and the contemporaneous policy rule.

Given the assumption of one change point, the institutional change point is compared with an optimal break point. The optimal break date is 1984:1 and is better according to the AIC and the BIC. The optimal break date corresponds to the findings of the Markov-Switching model where the smoothed probability decreases under 0.5 in 1984.[7]

[7] The difference of the likelihood value between the optimal break point model and the restricted Markov-Switching

For the UK, the two institutional change points 1976:1 and 1979:3, namely, the commitment to monetarism by the Bank of England and the beginning of the legislation of Margaret Thatcher remain after elimination of those not preferred by the selection criteria. Parameter estimates for the sub-samples only in one case yield a result as suggested by the theory. The difference of ξ between the second and the third sub-sample of the forward looking reaction function has the opposite sign compared to the difference of β, but is just significant at the 10 % level. A model with two optimal break points yields 1978:4 and 1980:3 and is preferred according to both criteria compared to the two institutional change points.

In the case of Germany, the single best institutional change point according to both criteria is the monetary union with the GDR in 1990:2. Differences between the estimates of the reaction parameters of the monetary policy rule of the sub-samples are insignificant for both specifications. The optimal break point is somewhat later in 1993:1. It is preferred by both model selection criteria.

Finally, for France one institutional change point in 1979:2 is preferred against specifications with more institutional change points and against no change point according to both criteria. Before 1979:2 $\widehat{\beta}_1 = 0.7659$ and the corresponding estimate $\widehat{\beta}_2 = 0.3633$. This change point also corresponds to the end of the first regime in the Markov-Switching model according to the smoothed probabilities. However, the differences between the sub-sample estimates of the monetary policy rule cannot be found significant. Compared to an optimal break point (1978:3) mixed evidence arises. The AIC favors the optimal point in 1978:3 while the BIC favors the institutional change point.

The approach to evaluate institutional change points yields no evidence for a link between the leading properties of the yield spread and the monetary policy. Indeed, estimates of the leading properties in different sub-samples according to institutional change points show significant differences. However, with exception of France other break points are preferred by both criteria compared to those proposed by institutional changes. In the case of France the results in Section 3 reject the hypothesis of a single change point. Furthermore, the institutional change points that bring up significant differences of the leading properties do not divide the estimates of the reaction parameters of the monetary policy in the same manner.

model vanishes as soon as one evaluates the likelihood of the Markov-Switching model with the smoothed and not with the filtered state probabilities.

5 Joint Markov Switching model

The given institutional change points may not reflect the behavior of the central banks. It is possible that institutional changes have some delay until they influence policy. It is also possible within the ongoing institutions, behavior changes are triggered by exogenous shocks other than changes in politics or chairmen rotation. Therefore, a second approach to identify monetary policy regimes is chosen based on a Markov-Switching model.

5.1 Markov-Switching monetary policy rules

The Markov-Switching monetary policy rule can be seen as an extension of Equation (7). Thereby the reaction parameters ξ and η as well as the target values are modeled to be state dependent, whereas the state variable w_t can take discrete values. Each value represents a monetary policy regime.

The resulting econometric model is the following Markov-Switching model:

$$i_t = (1-\rho)(\tilde{\zeta}_{w_t} + \eta_{w_t}\pi_{t+j} + \xi_{w_t}x_{t+k}) + \rho i_{t-1} + \sigma_{u_t}e_t \ , \quad e_t \sim N(0,1) \tag{12}$$

$$p_w^{ij} = P(w_t = i | w_{t-1} = j) \ , \quad p_u^{ij} = P(u_t = i | u_{t-1} = j)$$

In accordance with Wesche (2003), an own Markov-Switching process, u_t, is introduced for the error variance to provide a better fit to the data.

In case of $k = j = 0$, the model in Equation (12) represents a Markov-Switching Taylor rule with interest rate smoothing. In the case $k > 0$ and $j > 0$, the endogeneity problem appears again. However, a GMM estimation is not tractable for a Markov-Switching model. Therefore Wesche (2003) replaces π_{t+j} and x_{t+k} by forecasted values where the information set is restricted up to time t. She applies an extra forecasting model, namely a VAR model with inflation, output gap and interest rates as variables.

Alternatively, Kim (2004) proposes a bias correction for the endogenous explanatory variables based on instrument variables Z_t:

$$\begin{aligned} i_t &= (1-\rho)(\tilde{\zeta}_{w_t} + \eta_{w_t}\pi_{t+j} + \xi_{w_t}x_{t+k}) + \\ &\quad \gamma_1(\pi_{t+j} - \delta_1 Z_t) + \gamma_2(x_{t+j} - \delta_2 Z_t) + \rho i_{t-1} + \sigma_{u_t}e_t \end{aligned} \tag{13}$$

If δ_1 and δ_2 are replaced by their OLS estimates according to the regression of π_{t+j} on Z_t and x_{t+k} on Z_t respectively, the model can be estimated like the classical Markov-Switching model employing the Hamilton filter.[8]

The estimation of monetary policy rules encounters the problem of measuring inflation and output gap, see among others Bueno (2004). To obtain a measure of the output gap the Hodrick-Prescott filter is applied on GDP. The growth rates of the consumer price indices are used as a measure of inflation. Furthermore, in the literature, different interest rates are used to reflect the policy instruments. Wesche (2003) uses the overnight rate while e.g. Clarida et al. (2000) use central bank rates as it is done here, too. The instrumental variables Z_t used in this analysis are lagged series of short and long term interest rates, inflation, GDP growth and output gap.

Based on Clarida et al. (2000) central banks are assumed to be forward looking. Both, j and k are set to 1. Additionally, a specification of the monetary policy rule with contemporaneous regressors is considered. For the US, the UK and France, the forward looking ($j = k = 1$) results will be reported while these exhibit hardly interpretable coefficient estimates for Germany so that for this country only the case ($j = k = 0$) is analyzed.[9]

5.2 Modeling state dependent monetary policy and leading properties

The hypothesis that monetary policy regimes correspond to the regimes of the predictive relation between yield spread and GDP growth will be assessed by imposing the same Markov Switching process on Equation (3) and (12). This can be expressed as $v_t = w_t$ and results in the following model:

$$\begin{aligned} i_t &= (1-\rho)(\tilde{\zeta}_{w_t} + \eta_{w_t}\pi_{t+j} + \xi_{w_t}x_{t+k}) + \rho i_{t-1} + \sigma_{u_t}e_t \\ \Delta y_t^k &= \alpha + \beta_{w_t}S_t + \varepsilon_t \\ p^{ij} &= P(w_t = i|w_{t-1} = j). \end{aligned} \tag{14}$$

Because of the results in Section 3 for the US and Germany a joint structural break is assumed, too. The structural break is again implemented as a restricted Markov-Switching model by $p^{22} = 1$. If the restricted model is preferred according to the model selection criteria its results will be presented instead of those of the unrestricted.

Furthermore, evidence of threshold effects within the leading properties can be found. Therefore

[8] If the relation between the instruments and the regressors also depends on the Markov-Switching process (δ_{1,w_t} and δ_{2,w_t}), the Hamilton filter has to be modified, see *Kim* (2004). For the Hamilton filter see *Hamilton* (1989).

[9] By and large, results for the US, the UK and France do not alter qualitatively when changing from $j = k = 1$ to $j = k = 0$.

they are included in the joint Markov-Switching model resulting in:

$$\begin{aligned} i_t &= (1-\rho)(\tilde{\zeta}_{w_t} + \eta_{w_t}\pi_{t+j} + \xi_{w_t}x_{t+k}) + \rho i_{t-1} + \sigma_{u_t}e_t \\ \Delta y_t^k &= \alpha_0 + \beta_{0,w_t}S_t + (\alpha_1 + \beta_{1,w_t}S_t)I_{(z_{t-d}>\lambda)} + \varepsilon_t \\ p^{ij} &= P(w_t = i|w_{t-1} = j). \end{aligned} \tag{15}$$

The threshold does not effect the monetary reaction function. This implies, that the threshold effect itself is not caused by monetary policy.

Figure 4 shows the smoothed probabilities of being in state 1 (high value of the β coefficient) based on the estimation of Equation (14). For the US, the structural break in the leading relationship can also be found within the joint Markov-Switching model. In case of the UK, obvious deviations of the inference on the states exist. However, the two phases with high β values identified in the single Markov-Switching model seem to remain. For Germany, the estimate of the joint Markov-Switching model does not signal a structural break. Two phases of a high β value can be found. The end of the second phase corresponds to the structural break in the single Markov-Switching model of the leading properties. In the case of France the smoothed probabilities of being in state 1 are identical to those of the single Markov-Switching model of the leading properties apart from a short period at the end of the eighties.

Figure 5 gives the smoothed probabilities of being in state 1 for the joint Markov-Switching model including a threshold, see Equation (15). The picture for the US seems to be unchanged by the threshold. Minor deviations of the state probabilities can be found for the UK and France, while the inclusion of a threshold leads to an obvious change of the probabilities of the states in the case of Germany. Just a smaller period around 1990 remains with probabilities higher than 0.5.

For Germany and France, a special situation arises by the beginning of the European currency union replacing the monetary policy of the national central banks by the monetary policy of ECB from 1999 onwards. Therefore, after 1998 the same (European) variables are used to continue the national variables for the monetary policy equation, namely central bank rate, inflation and output gap. For Germany and France, the currency union implies that if the leading properties are determined by the monetary policy, both countries will share the same regimes of the joint Markov-Switching model from 1999 onwards. This is not the case. For France the yield spread has leading properties until 2001, while Germany experiences no regime change after 1993.

5.3 Joint Markov-Switching model: Model evaluation

The joint Markov-Switching model will be compared to a model with two independent Markov Switching processes ($v_t \neq w_t$). Testing the null hypothesis of one joint Markov Switching process against the alternative of two different ones leads to the problem that the models are not nested in the classical sense. Therefore, standard parametric tests are not tractable.

Again the model selection criteria, AIC and BIC, are used. Furthermore, the R^2 of the leading relation is monitored. The joint Markov-Switching models, Equation (14) and Equation (15), are estimated and compared to the estimates of two separated Markov-Switching models given by Equation (3), or (5), respectively, and Equation (12). Furthermore, combinations of a Markov-Switching process for one equation and a linear process for the other are considered. Namely, the likelihood of the estimate of Equation (1) together with the likelihood of Equation (12) and the resulting selection criteria are reported to represent the case of a linear leading relationship and a Markov-Switching monetary policy rule. The case of a Markov-Switching leading relationship and a monetary policy rule with a linear mean equation, is reported, too. This is done to account for the case that one of the two data generating processes can be approximated best by a linear model. Then the model selection criteria would prefer the joint Markov-Switching model to the separated Markov-Switching model although no joint process exists.

Results of the model selection criteria for the model in Equation (14) and its challengers are given in Table 7. The results are presented for the forward-looking monetary policy rule in case of the US, the UK and France so that the bias correction of Kim (2004) is applied, compare Equation (13).[10] In the case of Germany, results for the contemporaneous and homoscedastic monetary policy rule are reported.

According to the AIC, the divided Markov-Switching model (Equations (3) and (12)) is preferred. The BIC favors the joint Markov-Switching model given in Equation (14). The same situation can be observed in the UK. Interestingly, the R^2 of the regression of the yield spread is improved in both countries by the joint estimation compared to the estimation of the single Markov-Switching model. This is also true for France. Although, neither the AIC nor the BIC favor the joint model. According to the AIC the divided Markov-Switching models is preferred; according to the BIC the monetary policy rule is linear and the leading relationship is Markov-Switching. This result implies that only according to the AIC evidence of a Markov-Switching monetary policy rule in France.

For Germany, the divided Markov-Switching model is preferred according to the AIC. The linear

[10] The contemporaneous specification gives qualitatively similar results.

leading relationship and a Markov-Switching monetary policy rule prevails regarding the BIC. The R^2 for the leading relationship is only slightly improved by the joint Markov-Switching model as compared to the linear model.

Moreover, the hypothesis of a joint Markov-Switching process is assessed for the case of a threshold process present in the leading relation. Table 8 reports the results of the model selection criteria. The inclusion of a threshold does not change the relations between the models for the US, the UK and France. In the case of Germany, the consideration of the threshold comes up with some evidence of the joint Markov-Switching process.

Summarizing the results of Table 7 and 8, evidence for a joint time variation of the leading relation and the reaction parameters of the monetary policy can at least be found for the US, the UK and Germany. For France, no evidence of a link between the time variation of the monetary policy and the leading relation can be found.

In the case of the US, the BIC favors a joint break of both mean equations independent from the inclusion of a threshold. However, according to the AIC, a single Markov-Switching process for the monetary policy is preferred.

The same is true for the UK with the difference that the joint time variation is no break but a joint switching. Again, the AIC favors separated processes.

According to the AIC, the joint Markov-Switching model with a threshold is the best to fit the German data. Given the threshold this is also true for the BIC. However, the overall preferred model according the BIC is the combination of a (stable) linear leading relationship and a Markov-Switching monetary policy rule.

The rejection of the hypothesis of a joint Markov-Switching according to the AIC for the US, the UK and France corresponds to the differing inference on the states of both separated Markov-Switching models. Figure 6 shows the smoothed probabilities of being in state 1 according to the estimates of the Markov-Switching model for the monetary policy rule given in Equation (12). For all countries the inference on the states of the monetary policy leads to different results in comparison with the inference states of the leading properties (see Figure 1). However, the results of the BIC (US and UK) and of the AIC in the case of Germany stress also the existence of some similarities between the regimes of the separate models. For instance the structural breaks of the leading properties in Germany and the US are accompanied by changes of the monetary policy.

5.4 Joint Markov-Switching model: Parameter estimates

Table 9 gives the ML estimates and their standard deviations of the parameters of Equation (14) and Table 10 reports the estimates of the parameters of Equation (15). Differences of the parameters of interest as well as the corresponding t-statistics are given in Table 11. Theory suggests that a high value of β comes along with a weak reaction of the monetary policy on the output gap and vice versa. For the comparison of the states of the Markov-Switching model this implies if $\beta_1 - \beta_2$ is negative $\xi_1 - \xi_2$ should be positive. For the US this is not true. The opposite is the case. The decrease of the value of the β coefficient is accompanied with an even significant decrease of the reaction on the output gap. This result does not change by the inclusion of a threshold process.

For all other countries, the signs of the differences are in keeping with the theory. For the UK, $\xi_1 - \xi_2$ is significantly negative and within the model without a threshold $\beta_1 - \beta_2$ is significantly positive. In case of the threshold model, the sign of the differences between the βs depend on the threshold. The difference above the threshold ($\Delta y^4_{t-4} > -0.41$) is significantly positive as theory suggests and the series Δy^4_{t-4} has values above this most of the time, compare Table 1. The impact of the threshold – that the differences between the β have varying signs – can be found for Germany and France, too.

In the case of Germany, $\xi_1 - \xi_2$ is significantly negative for both models. In the model without threshold $\beta_1 - \beta_2$ is positive but not significantly different to zero. In the model with threshold, this is true for $\beta_{1,1} - \beta_{1,2}$, too. $\beta_{0,1} - \beta_{0,2}$ is significantly negative but corresponds to the rarer case of $\Delta y^4_{t-4} > 3.53$.

For France, the model without threshold exhibits an estimate for $\xi_1 - \xi_2$ that is insignificant. The inclusion of a threshold changes this; $\xi_1 - \xi_2$ and $\beta_{1,1} - \beta_{1,2}$ are in keeping with theory. In the case $\Delta y^4_{t-4} > 3.93$ the β difference is negative but the Δy^4_{t-4} takes values below this level most of the time.

The parameter estimates of the joint Markov-Switching models show some evidence for the implications of the theoretical model of Estrella (2005) at least for the UK and France. However, one has to take into account that no model selection criterium favors the joint Markov-Switching model in the case of France. The main contradiction is found in the US. There is evidence for a joint break but the change has the wrong direction.

Partly, this contradiction might stem from the quite restrictive break model. This model disregards the recurrence of the leading property according to the autoregressive time varying coefficient model. However, this recurrence seems to be accompanied by a new regime in monetary policy according to the model in Equation (12), compare Figure 2 and 4, upper panel. This can also be shown via the inclusion of the smoothed probabilities of the monetary state in the linear regression in the following way:

$$\Delta y_t^k = \alpha + \beta_1 P(w_t = 1|I_T)S_t + \beta_2(1 - P(w_t = 1|I_T))S_t + \varepsilon_t. \tag{16}$$

The R^2 of this model is 0.3484 compared to 0.2967 for the linear model. Assuming Gaussian errors both model selection criteria (AIC and BIC) prefer the higher parameterized model to the linear model. The relevant parameter estimates of the models Equation (12) and (16) are given in the following table in which the t-statistics are underset in paranthesis:

$$\widehat{\beta}_1 - \widehat{\beta}_2 = \underset{(2.914)}{0.557} \qquad \widehat{\xi}_1 - \widehat{\xi}_2 = \underset{(-0.929)}{-0.104} \qquad \widehat{\eta}_1 - \widehat{\eta}_2 = \underset{(5.477)}{0.936}$$

The signs of the differences are as theory suggests. Although the difference of the reaction to the output gap between the states is insignificant. The monetary policy states are differentiated by the reaction on inflation. According to the model of Estrella (2005), a higher inflation reaction leads to less importance of the yield spread as a leading indicator.

6 Conclusion

Evidence for time varying leading properties of the yield spread can be found for all countries considered. For the US and Germany, the type of time variation fitting the data best is a structural break in 1984 and 1993 respectively. For the UK and France, recurring phases of different values of β can be identified. Thus, the reply to the first question of this analysis is unambiguous. However, the type and the intensity of variation depends on the country under consideration. For France a large likelihood gain can be obtained by taking time varying leading properties into account while for Germany the evidence is rather weak.

No clear determination of the time variation of the leading properties by the time variation of the monetary policy can be proven. However, some coincidences can be found between the leading properties of the yield spread and monetary policy. The structural breaks of the leading properties in the US and Germany are accompanied by changes in the monetary policy in the neighborhood these dates according to the Markov-Switching monetary policy rules. The first regime change of the Markov-Switching model of the leading properties of France corresponds to the introduction of the ECU in 1979; a further regime change in 1993 corresponds to the ratification of the Maastricht treaty followed by a strict tightening of fiscal policy in France. These coincidences are contradicted by large deviations, e.g. the Markov-Switching regimes of the US monetary policy rule do not determine the structural break of the leading properties of the yield spread.

On the one hand, one can conclude that leading properties of the yield spread and monetary policy do not share the same states. On the other hand, evidence of a connection can be found with the exception of France. According to the BIC joint Markov-Switching models are preferred in the US, the UK and Germany. Furthermore, the signs of the differences between the states of the reaction parameters and the β coefficients are in keeping with the suggestion of the model of Estrella (2005) in case of the UK, Germany and France. However, for the US the parameter estimates contradict the theory although there is evidence for the joint Markov-Switching model.

Sources other than monetary policy seem to influence the leading properties of the yield spread or the way monetary policy influences the leading properties. The model of Estrella (2005) gives first insights but is based on the expectation hypothesis that may not hold. Time varying term or risk premia might be important (twice, structural breaks of the leading properties of the yield spread are accompanied by volatility reductions of the business cycle). The influence of international financial markets and exchange rate fluctuations are disregarded, too. These are only examples of aspects that can influence the decision to buy or sell long term bonds other than interest rate expectations. Explaining why and how the yield spread leads GDP growth needs further investigation.

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Appendix

A The model of Estrella [2005]

The macroeconomic model of Estrella (2005) consists of the following equations. The IS curve is given by:

$$x_t = b_1 x_{t-1} - b_2 r_{t-1} + \epsilon_t^x, \tag{17}$$

where x_t denotes the output gap and r_{t-1} the long term real interest rate. Random shocks to the output gap are denoted as ϵ_t^x. The dynamics of inflation π_t is modeled via the Phillips curve:

$$\pi_t = \pi_{t-1} + a x_{t-1} + \epsilon_t^\pi, \tag{18}$$

where again random shocks are considered via ϵ_t^π. The central bank behavior is modeled according to the following reaction function:

$$i_t = g_i i_{t-1} + g_\pi \pi_t + g_x x_t + (1 - g_r - g_\pi)\pi^*, \tag{19}$$

where π^* denotes the desired level of inflation. The long term nominal interest rate R_t is determined by the Fisher equation:

$$R_t = r_t + \frac{1}{2}(E_t[\pi_{t+1}] + E_t[\pi_{t+2}]) \tag{20}$$

where E_t represents the expectations at time t. Furthermore, the expectations hypothesis is given by:

$$R_t = \frac{1}{2}(i_t + E_t[i_{t+1}]). \tag{21}$$

Solving this model yields the relationship between yield spread $(R_t - i_t)$ and expectations of future output:

$$E_t[x_{t+1}] = \frac{2}{g_x}(R_t - i_t) + \frac{1 - g_i}{g_x}(i_t - \pi^*) + \frac{g_\pi}{g_x}(\pi^* - \pi_t - a x_t). \tag{22}$$

If alternatively the reaction function is modeled forward-looking

$$i_t = g_i i_{t-1} + g_\pi E_t[\pi_{t+1}] + g_x E_t[x_{t+1}] + (1 - g_r - g_\pi)\pi^* \tag{23}$$

the Equation (22) will change to:

$$\begin{aligned} E_t[x_{t+2}] \quad &= \quad \frac{2}{g_x}(R_t - i_t) + \frac{1 - g_i}{g_x}(i_t - \pi^*) + \frac{g_\pi}{g_x} \\ &\quad \frac{1}{2 - ab_2}\left[\pi^* - (2 + ab_2)\pi_t - a(2 + 2b_1 + b_2)y_t + 2ab_2 R_t\right]. \end{aligned} \tag{24}$$

According to both approaches the coefficient of the spread for the expectations of real activity is the inverse of the monetary policy reaction parameter g_x. Equations (17) through (23) are given in Estrella (2005). Equation (24) is deduced correspondingly.

B Tables

Table 1: Data description – summary statistics

		mean	std-dev	lower quartile	upper quartile
US	Δy^4	2.943	2.321	1.771	4.266
	S	0.652	1.972	-0.238	1.908
	i	6.624	2.578	5.000	7.845
	x	0.029	4.700	-3.035	3.096
	π	4.911	3.473	2.606	6.442
UK	Δy^4	2.120	2.052	1.536	3.424
	S	0.593	1.938	-0.730	2.070
	i	9.763	3.250	6.645	12.408
	x	-0.001	4.677	-3.231	2.512
	π	7.077	6.855	2.025	9.466
GER	Δy^4	2.010	2.012	0.697	3.302
	S	0.971	1.727	-0.228	2.219
	i	4.565	1.866	3.000	6.000
	x	0.040	3.685	-2.369	2.007
	π	3.027	2.569	1.156	4.454
FRA	Δy^4	2.182	1.452	1.187	3.369
	S	1.042	1.484	0.099	2.143
	i	9.611	2.182	7.954	10.985
	x	0.094	3.381	-2.416	2.687
	π	5.498	4.503	2.034	9.426

Summary statistics of all endogenous and exogenous variables. Δy^4: growth rates of GDP; S: yield spread; i: central bank rate; x: output gap; π: inflation.

Table 2: Time varying leading properties: Model Comparison

		linear[a]	MS 2[b]	MS 2[c]	auto[d]
US	LL	-224.85	–	-214.31	-215.93
	AIC	2.0527	–	1.9758	1.9904
	BIC	2.0893	–	2.0368	2.0514
	R^2	0.2967	–	0.4278	0.4335
	BP[e]			1984:2	
UK	LL	-221.54	-209.25	-214.62	-211.38
	AIC	2.0229	1.9392	1.9786	1.9494
	BIC	2.0595	2.0124	2.0396	2.0104
	R^2	0.2150	0.3674	0.3221	0.4077
	BP[e]			1975:2	
GER	LL	-220.95	–	216.65	-218.65
	AIC	2.0176	–	1.9068	2.0149
	BIC	2.0542	–	2.0578	2.0759
	R^2	0.1735	–	0.2420	0.2292
	BP[e]			1993:2	
FRA	LL	-181.5	-158.86	-177.14	-157.8
	AIC	1.6622	1.4852	1.6409	1.4667
	BIC	1.6988	1.5585	1.7019	1.5277
	R^2	0.2455	0.5137	0.3140	0.5404
	BP[e]			1978:3	

a: Represents results of the linear model. see Equation 1.
b: Represents results of the Markov-Switching model (Equation 3).
c: Represents results of the restricted Markov-Switching model with $p^{22} = 1$.
d: Represents results of the autoregressive time varying coefficient model (see Equation 4).
e: Represents the point in time where $P(v_t = 2|I_T) > 0.5$ firstly (implicit break point).

Table 3: Time varying leading properties of threshold models: Model Comparison

		linear[a]	MS 2[b]	MS 2[c]	auto[d]
US	LL	-218.02	–	-205.38	-207.74
	AIC	2.0182	–	1.9314	1.9436
	BIC	2.0914	–	2.0412	2.0412
	R^2	0.3813	–	0.5138	0.5378
	$\hat{\lambda}$	-1.4829	–	-0.1014	-0.4468
	BP[e]			1984:2	
UK	LL	-203.83	-191.65	-195.69	-194.57
	AIC	1.8904	1.8167	1.8441	1.8250
	BIC	1.9636	1.9387	1.9539	1.9226
	R^2	0.4307	0.4566	0.5201	0.5180
	$\hat{\lambda}$	2.1010	2.1010	2.1010	2.1010
	BP[e]			1975:2	
GER	LL	-214.68	–	-209.66	-212.98
	AIC	1.9881	–	1.9699	1.9908
	BIC	2.0613	–	2.0798	2.0885
	R^2	0.2637	–	0.3303	0.2962
	$\hat{\lambda}$	3.0410	–	3.0410	3.0410
	BP[e]			1993:2	
FRA	LL	-175.32	-149.53	-171.70	-148.80
	AIC	1.6335	1.4372	1.6279	1.4126
	BIC	1.7067	1.5593	1.7378	1.5103
	R^2	0.3250	0.5747	0.3749	0.5990
	$\hat{\lambda}$	3.7516	3.7516	3.9878	3.7516
	BP[e]			1979:1	

a: Represents results of the linear model with a threshold. see Equation 2.
b: Represents results of the Markov-Switching model with a threshold (see Equation 5).
c: Represents results of the restricted Markov-Switching model with $p^{22} = 1$.
d: Represents results of the autoregressive time varying coefficient model with a threshold (see Equation 6).
e: Represents the point in time where $P(v_t = 2|I_T) > 0.5$ firstly (implicit break point).

Table 4: Monetary Policy Events

US	1979:3	Paul Volcker becomes chairman of the FED
	1987:4	Alan Greenspan becomes chairman of the FED
UK	1976:1	Commitment to monetarism
	1979:3	Margaret Thatcher becomes PM
	1990:4	Entry into the European currency system
	1992:3	Withdrawal from the European currency system
	1997:2	Commitment to inflation targeting (independence from politics)
GER	1975:4	Commitment to money supply management
	1979:2	Introduction of the European currency system
	1990:2	Monetary Union with GDR
	1999:1	Introduction of the Euro
FRA	1979:2	Introduction of the European currency system
	1993:3	Independence from politics
	1999:1	Introduction of the Euro

Table 5: Monetary Policy Events and optimal break points

		without break	institutional change point[a]	optimal break point
US	logLik	-224.85	-215.68	-210.86
	AIC	2.053	1.979	1.945
	BIC	2.089	2.028	2.006
	date	-	1987:4	1984:1
UK	logLik	-221.54	-215.04	-208.98
	AIC	2.023	1.982	1.946
	BIC	2.059	2.043	2.031
	date	-	1976:1 1979:3	1978:4 1980:3
GER	logLik	-220.95	-216.44	-213.13
	AIC	2.018	1.986	1.965
	BIC	2.054	2.035	2.026
	date	-	1990:2	1993:1
FR	logLik	-181.50	-176.48	-174.96
	AIC	1.662	1.626	1.621
	BIC	1.699	1.675	1.682
	date	-	1979:2	1978:3

For the calculation of the model selection criteria of the institutional change point models the change point is interpreted as a exogenous regressor. In the case of the optimal break point model it is an additional parameter and is penalized correspondingly.

a: All previously considered institutional change points (see Table 4) that do not improve the model according to the AIC and BIC have been eliminated. Only the best models according the model selection criteria are reported here.

Table 6: Results for $\beta_1 - \beta_2$. $\xi_1 - \xi_2$ and $\eta_1 - \eta_2$

	USA	UK	GER[a]	FRA
$\beta_1 - \beta_2$	0.763 (4.191)	−0.742 (−2.457)	0.564 (2.878)	0.483 (3.670)
$\xi_1 - \xi_2$	0.201 (0.575)	−4.333 (−1.279)	−0.041 (−0.183)	0.193 (0.270)
$\eta_1 - \eta_2$	0.350 (0.450)	1.006 (0.859)	0.004 (0.012)	−0.003 (−0.005)
$\beta_1 - \beta_3$		−1.181 (−4.201)		
$\xi_1 - \xi_3$		−4.626 (−1.589)		
$\eta_1 - \eta_3$		−1.639 (−1.399)		
$\beta_2 - \beta_3$		−0.439 (−3.491)		
$\xi_2 - \xi_3$		−0.293 (−0.125)		
$\eta_2 - \eta_3$		−1.639 (−2.258)		
Forward-looking policy rule				
$\beta_1 - \beta_2$	0.763 (4.191)	−0.742 (−2.457)	0.564 (2.878)	0.483 (3.670)
$\xi_1 - \xi_2$	0.121 (0.203)	−0.307 (−0.293)	−0.011 (−0.029)	0.062 (0.133)
$\eta_1 - \eta_2$	1.894 (0.951)	−0.627 (−1.298)	−0.163 (−0.357)	0.556 (1.172)
$\beta_1 - \beta_3$		−1.181 (−4.201)		
$\xi_1 - \xi_3$		0.247 (0.261)		
$\eta_1 - \eta_3$		−0.668 (−1.383)		
$\beta_2 - \beta_3$		−0.439 (−3.491)		
$\xi_2 - \xi_3$		0.554 (1.373)		
$\eta_2 - \eta_3$		−0.668 (−0.085)		

Results for sub-sample estimates according institutional change points of the linear leading relationship and monetary policy reaction functions.
Values of the t-statistics in parentheses.

Table 7: Markov Switching Monetary Policy Rules: Two vs. joint MS

		two MS processes	joint MS process	MS and linear[a]	linear and MS[b]
US	LogLik	−246.90 (−32.59/−214.31)	-251.35	−257.44 (−32.59/−224.85)	−262.98 (−48.67/−214.31)
	AIC	2.414	2.436	2.481	2.513
	BIC	2.670	2.667	2.701	2.709
	R^2	0.4278	0.4368	0.2967	0.4278
UK	LogLik	−332.41 (−123.16/−209.25)	-336.08	−344.70 (−123.16/−221.54)	−344.63 (−135.38/−209.25)
	AIC	3.184	3.199	3.268	3.249
	BIC	3.440	3.431	3.487	3.444
	R^2	0.3674	0.3686	0.2150	0.3674
GER[c]	LogLik	−259.15 (−42.50/−216.65)	-262.79	−263.45 (−42.50/−220.95)	−271.14 (−54.49/−216.65)
	AIC	2.479	2.494	2.491	2.542
	BIC	2.674	2.664	2.649	2.676
	R^2	0.2420	0.1797	0.1735	0.2420
FRA	LogLik	−178.46 (−19.60/−158.86)	-184.29	−201.10 (−19.60/−181.50)	−188.82 (−29.96/−158.86)
	AIC	1.797	1.831	1.974	1.845
	BIC	2.053	2.063	2.194	2.041
	R^2	0.5137	0.5199	0.2455	0.5137

In case of models not joining a common state the likelihood can be factored in one part for the monetary policy rule and one for the leading relation. The corresponding log likelihood values are given in parentheses under their sum. The first value in the parenthesis corresponds to the monetary policy rule.

a: Markov-Switching policy rule and linear leading model.

b: Linear policy rule and Markov-Switching leading model.

c: The variance is modeled homoscedastically.

Table 8: Joint Markov-Switching in Monetary Policy Rules and leading properties with threshold

		two MS processes	joint MS process	MS and linear[a]	linear and MS[b]
US	LogLik	-237.97 $(-32.59/-205.38)$	-242.42	-250.61 $(-32.59/-218.02)$	-254.05 $(-48.67/-205.38)$
	AIC	2.360	2.382	2.447	2.460
	BIC	2.653	2.651	2.703	2.692
	R^2	0.5138	0.5136	0.3813	0.5138
	$\hat{\lambda}$	-0.1014	-0.1014	-1.4829	-0.1014
UK	LogLik	-314.81 $(-123.16/-191.65)$	-317.45	-326.99 $(-123.16/-203.86)$	-327.03 $(-135.38/-191.65)$
	AIC	3.016	3.022	3.090	3.081
	BIC	3.260	3.242	3.285	3.264
	R^2	0.4566	0.4738	0.4307	0.4566
	$\hat{\lambda}$	2.1010	-0.4101	2.1010	2.1010
GER[c]	LogLik	-252.16 $(-42.50/-209.66)$	-250.53	-257.18 $(-42.50/-214.68)$	-264.14 $(-54.49/-214.68)$
	AIC	2.497	2.464	2.506	2.560
	BIC	2.802	2.745	2.762	2.804
	R^2	0.3303	0.3620	0.2637	0.3303
	$\hat{\lambda}$	3.0410	3.5263	3.0410	3.0410
FRA	LogLik	-169.13 $(-19.60/-149.53)$	-173.09	-194.92 $(-19.60/-175.32)$	-179.49 $(-29.26/-149.53)$
	AIC	1.749	1.767	1.945	1.797
	BIC	2.054	2.047	2.202	2.041
	R^2	0.5747	0.5682	0.3250	0.5747
	$\hat{\lambda}$	3.7516	3.9263	3.7516	3.7516

In case of models not joining a common state the likelihood can be factored in one part for the monetary policy rule and one for the leading relation. The corresponding log likelihood values are given in parenthesis under their sum. The first value in the parenthesis corresponds to the monetary policy rule.

a: Markov-Switching policy rule and linear leading model.

b: Linear policy rule and Markov-Switching model for the leading relationship.

c: The variance is modeled homoscedastically.

Table 9: Parameter estimates for the joint Markov-Switching model

	USA	UK	GER	FRA
p_{11}	1.000 (0.000)	0.918 (0.037)	0.974 (0.023)	0.935 (0.035)
p_{22}	0.974 (0.027)	0.686 (0.109)	0.890 (0.082)	0.926 (0.040)
ζ_1	−2.403 (6.747)	5.564 (2.606)	5.186 (1.018)	4.254 (2.452)
ξ_1	1.832 (1.013)	0.104 (0.268)	−0.265 (0.272)	0.022 (0.434)
η_1	2.135 (1.318)	0.931 (0.315)	1.414 (0.503)	0.717 (0.434)
ζ_2	3.619 (1.319)	0.783 (1.162)	2.287 (0.665)	4.645 (2.773)
ξ_2	0.882 (0.573)	1.768 (0.439)	0.862 (0.312)	2.489 (1.657)
η_2	0.598 (0.372)	1.055 (0.203)	0.327 (0.192)	−0.033 (0.493)
ρ	0.967 (0.020)	0.910 (0.018)	0.912 (0.024)	0.977 (0.015)
γ_x	−0.021 (0.010)	−0.048 (0.029)		−0.009 (0.019)
γ_π	−0.045 (0.014)	−0.063 (0.022)		−0.021 (0.016)
σ_1^2	0.007 (0.005)	0.064 (0.017)	0.106 (0.016)	0.042 (0.008)
σ_2^2	0.342 (0.072)	1.937 (0.463)		0.917 (0.339)
p_{11}^u	0.717 (0.148)	0.984 (0.022)		0.976 (0.019)
p_{22}^u	0.782 (0.141)	0.979 (0.024)		0.900 (0.078)
α	3.227 (0.221)	2.213 (0.186)	1.579 (0.226)	1.660 (0.120)
β_1	1.101 (0.122)	1.675 (0.231)	0.915 (0.333)	1.030 (0.087)
β_2	0.088 (0.133)	0.223 (0.089)	0.467 (0.131)	0.053 (0.084)
σ^2	2.641 (0.377)	2.172 (0.329)	3.078 (0.425)	0.764 (0.113)

Maximum Likelihood estimates of the Model in Equation 14. Asymptotic standard deviations in parenthesis.

Table 10: Parameter estimates for the joint Markov-Switching model with threshold

	USA	UK	GER	FRA
p_{11}	1.000 (1.000)	0.916 (0.033)	0.991 (0.010)	0.924 (0.029)
p_{22}	0.977 (0.025)	0.680 (0.118)	0.866 (0.125)	0.932 (0.041)
ζ_1	−0.683 (1.934)	5.732 (1.150)	5.617 (1.937)	6.597 (1.187)
ξ_1	1.653 (0.350)	0.128 (0.197)	−0.830 (0.645)	0.103 (0.197)
η_1	1.827 (0.289)	0.903 (0.190)	0.675 (0.492)	0.143 (0.202)
ζ_2	3.661 (1.373)	0.659 (0.974)	0.677 (0.122)	5.370 (1.654)
ξ_2	0.766 (0.230)	1.806 (0.422)	0.816 (0.198)	1.792 (0.771)
η_2	0.558 (0.362)	1.087 (0.176)	0.453 (0.149)	0.951 (0.341)
ρ	0.963 (0.008)	0.912 (0.016)	0.904 (0.021)	0.961 (0.015)
γ_x	−0.020 (0.010)	−0.049 (0.028)		−0.044 (0.017)
γ_π	−0.045 (0.016)	−0.062 (0.022)		0.007 (0.015)
σ_1^2	0.008 (0.004)	0.063 (0.014)	0.115 (0.016)	0.020 (0.005)
σ_2^2	0.341 (0.073)	1.897 (0.372)		0.413 (0.145)
p_{11}^u	0.759 (0.080)	0.989 (0.015)		0.927 (0.035)
p_{22}^u	0.809 (0.110)	0.986 (0.016)		0.872 (0.066)
α_0	2.928 (0.230)	2.519 (0.148)	2.296 (0.621)	2.984 (0.318)
$\beta_{0.1}$	1.247 (0.134)	1.836 (0.181)	1.048 (0.222)	0.008 (0.207)
$\beta_{0.2}$	0.229 (0.138)	0.216 (0.071)	−9.003 (1.969)	1.065 (0.345)
α_1	4.819 (0.526)	0.428 (0.362)	2.210 (0.442)	1.521 (0.104)
$\beta_{1.1}$	1.000 (0.244)	−1.330 (0.829)	1.224 (0.912)	1.000 (0.076)
$\beta_{1.2}$	−0.974 (0.420)	0.593 (0.216)	0.309 (0.230)	0.052 (0.073)
σ^2	2.242 (0.321)	1.433 (0.206)	2.409 (0.351)	0.675 (0.094)

Maximum Likelihood estimates of the Model in Equation 15. Asymptotic standard deviations in parenthesis.

Table 11: Results for $\beta_1 - \beta_2$. $\xi_1 - \xi_2$ and $\eta_1 - \eta_2$

	USA	UK	GER[a]	FRA
Results for the model in Equation (14)				
$\beta_1 - \beta_2$	1.013 (5.322)	1.453 (5.708)	0.448 (1.229)	0.977 (9.619)
$\xi_1 - \xi_2$	0.951 (1.856)	−1.664 (−3.584)	−1.128 (−2.230)	−2.467 (−1.377)
$\eta_1 - \eta_2$	1.537 (1.257)	−0.124 (−0.366)	1.087 (1.984)	0.750 (0.979)
Results for the threshold model in Equation (15)				
$\beta_{0.1} - \beta_{0.2}$	1.017 (5.095)	1.620 (8.043)	−9.680 (−4.883)	−1.057 (−2.856)
$\beta_{1.1} - \beta_{1.2}$	1.974 (3.650)	−1.923 (−2.529)	0.915 (1.042)	0.948 (9.928)
$\xi_1 - \xi_2$	0.887 (3.019)	−1.678 (−4.211)	−1.646 (−2.312)	−1.689 (−2.219)
$\eta_1 - \eta_2$	1.268 (2.841)	−0.183 (−0.789)	0.223 (0.435)	−0.808 (−2.105)

Differences of estimates of the joint Markov-Switching models in Equations (14) and (15). Compare Table 9 and 10. Values of the t-statistics in parentheses.

C Figures

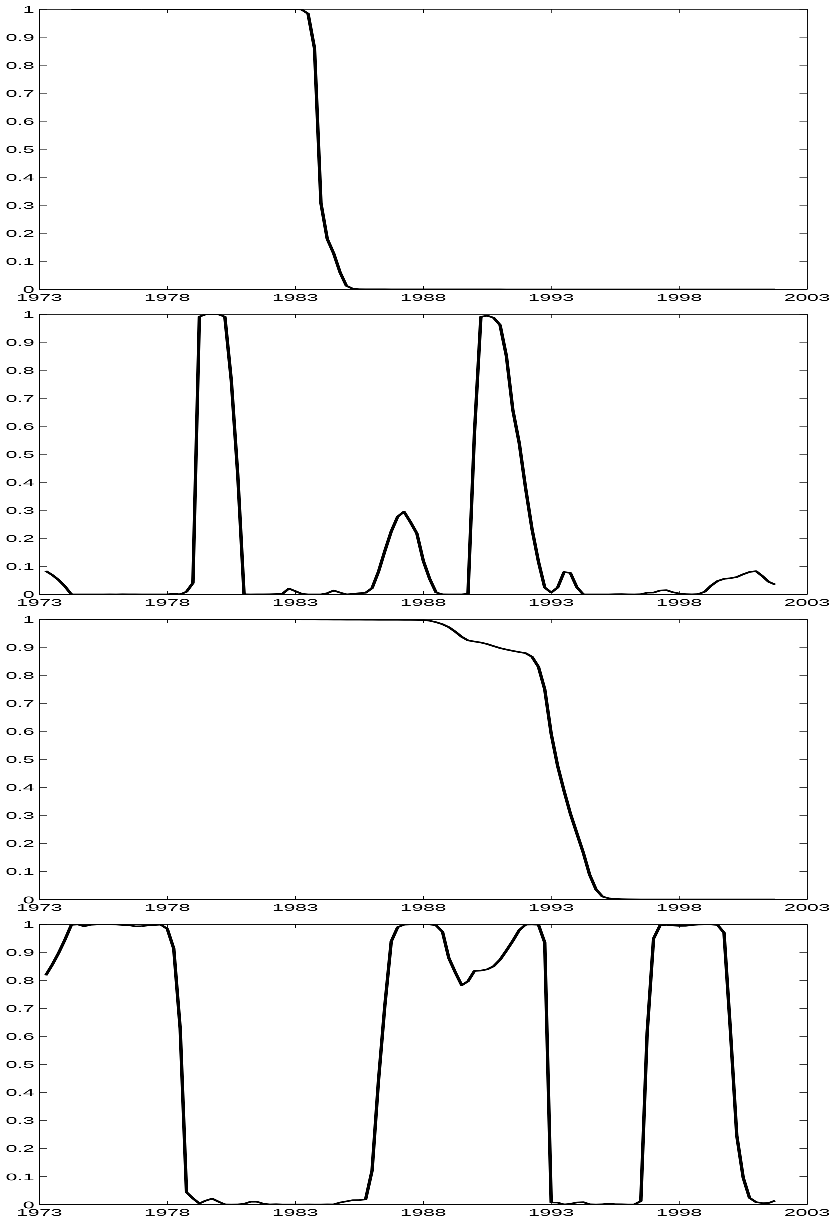

Fig. 1. Smoothed probabilities for the high state ($v_t = 1$) of the Markov-Switching model in Equation (3): USA (top panel), UK (second panel), Germany (third panel), France (bottom panel)

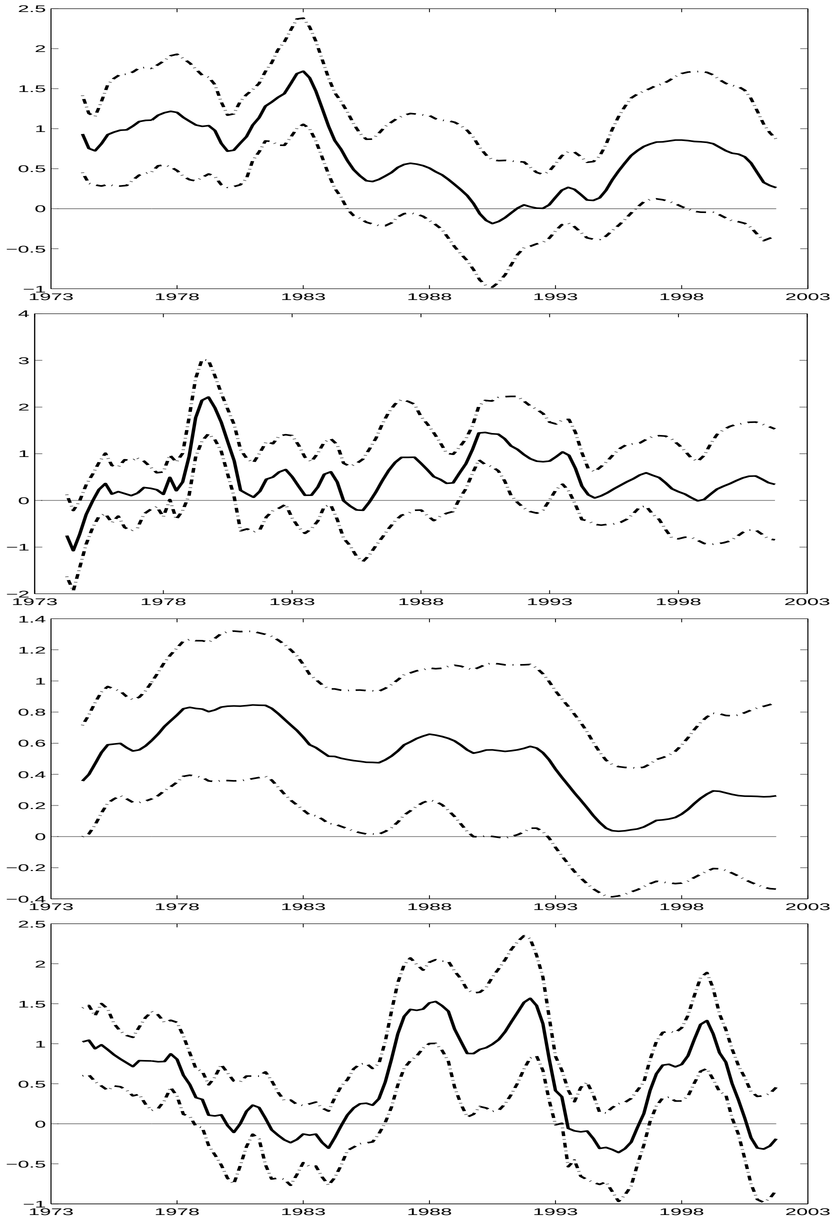

Fig. 2. Time varying coefficient β_t according to Equation (4) [smoothed: $E(\beta_t|I_T)$], the dashed lines show the confidence bands with two times the standard-deviation: USA (top panel), UK (second panel), Germany (third panel), France (bottom panel).

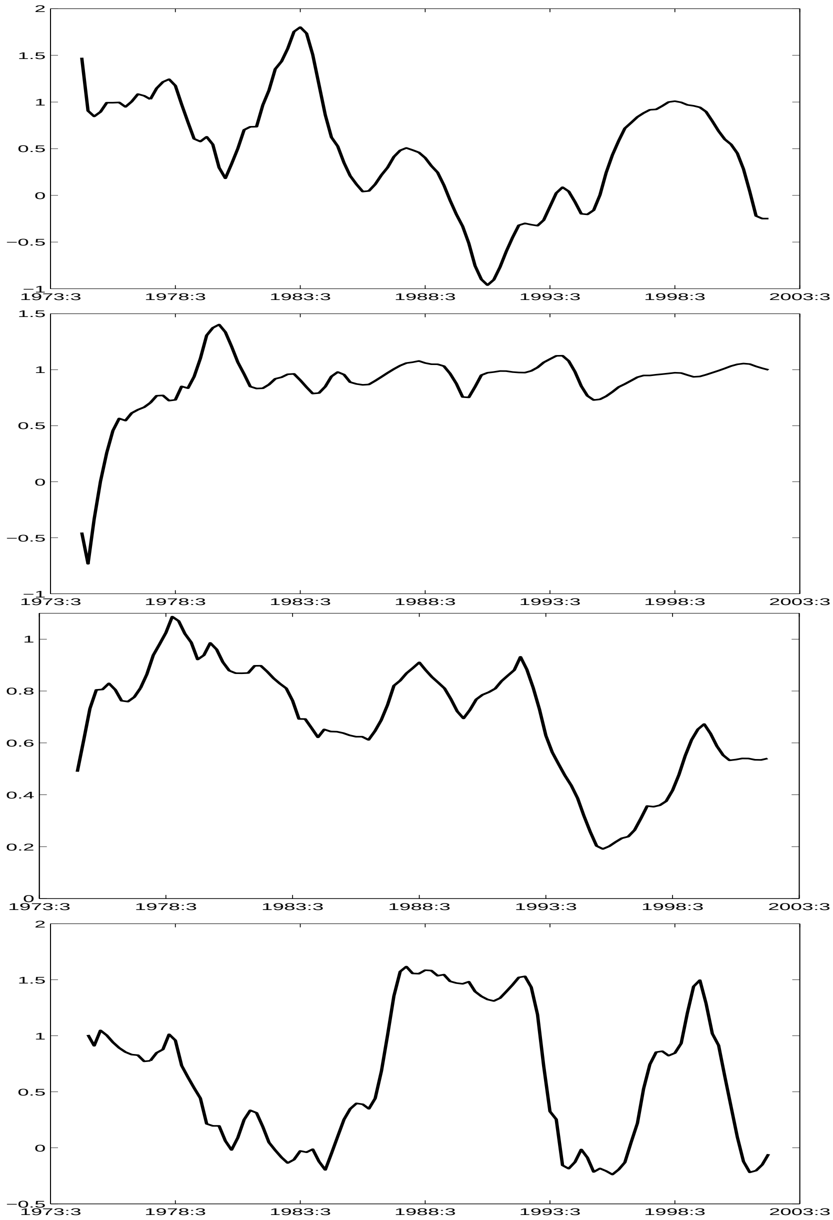

Fig. 3. Variation in time of coefficient β according the time varying coefficient model with threshold, see Equation 6, [smoothed: $E(\beta_t|I_T)$]: USA (top panel), UK (second panel), Germany (third panel), France (bottom panel)

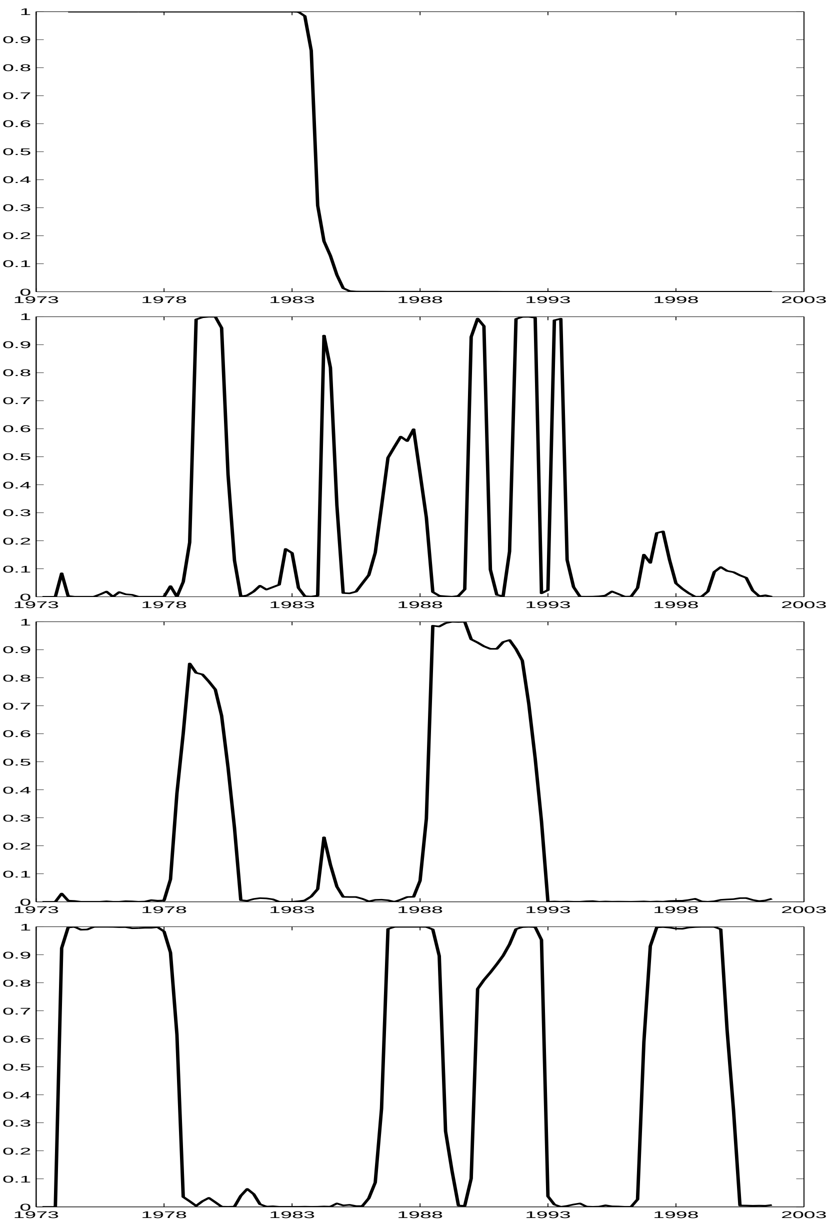

Fig. 4. Smoothed probabilities of state 1 of the joint Markov-Switching model (Equations 14): USA (top panel), UK (second panel), Germany (third panel), France (bottom panel)

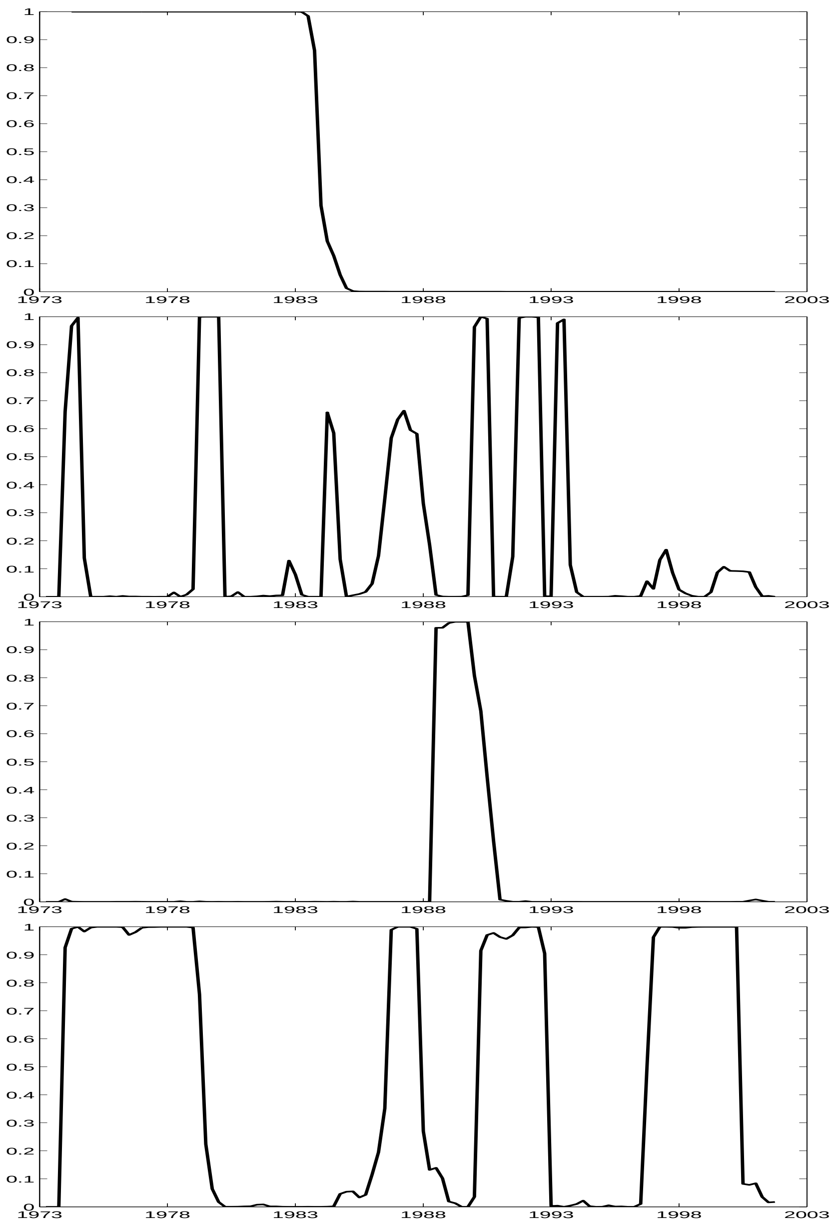

Fig. 5. Smoothed probabilities of state 1 of the joint Markov-Switching model with threshold (Equations 15): USA (top panel), UK (second panel), Germany (third panel), France (bottom panel)

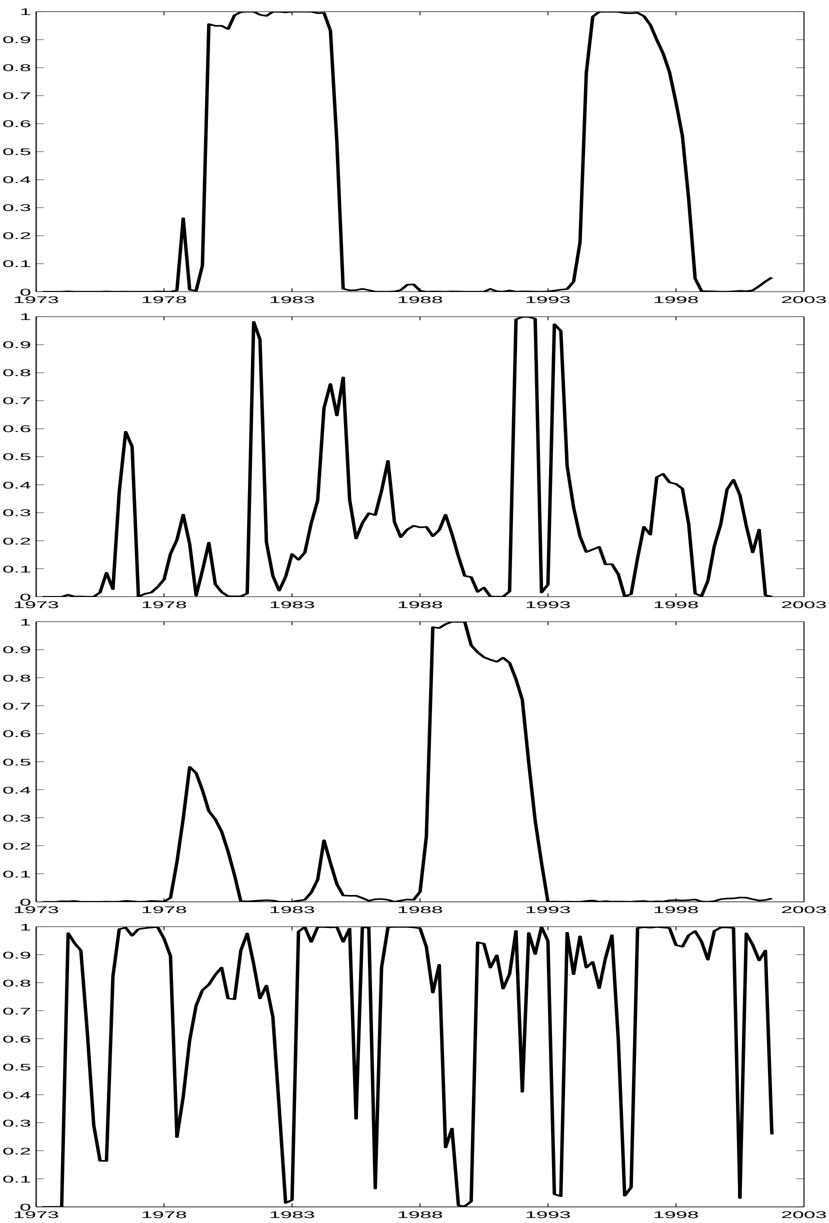

Fig. 6. Smoothed probabilities of state 1 of the Markov-Switching model of the Taylor rule according the model in Equation 12: USA (top panel), UK (second panel), Germany (third panel), France (bottom panel)