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The Regime-Dependent Determination of Credibility: A New Look at European Interest Rate Differentials

by
Peter Tillmann

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Peter Tillmann*
University of Cologne

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Abstract

Abstract: Once you allow for persistence in macroeconomic variables, two aspects of exchange rate credibility emerge whose relative importance can vary over time. Hence, the effect of policy measures on interest rate differentials becomes ambiguous. In this paper, a Markov-switching VAR that allows for parameter shifts across regimes is employed to test the hypothesis of regime-dependent determination of credibility for major EMS countries. The model separates two regimes that are distinct with respect to the time series properties of the interest rate spread. Regime-dependent impulse response functions reveal substantial differences in the response of spreads to macroeconomic shocks across regimes.

Keywords: Regime-switching, VAR, interest rate differentials, regime-dependent impulse response functions, credibility

JEL classification: E4, F3

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Address: Peter Tillmann, Wirtschaftspolitisches Seminar, University of Cologne, Robert-Koch-Str. 41, 50931 Cologne, Germany, phone: +49+221-4703449, fax: +49+221-4705187, e-mail: peter.tillmann@uni-koeln.de

1 Introduction

In the years following the crisis of the European Monetary System (EMS) in 1992/93, an extensive literature emerged trying to establish reliable links between certain policy measures and market expectations of future exchange rate policy. These models shed light on the strategic interaction between policy makers, the central banks and financial markets under uncertainty. Since the true weights in the government's loss function are assumed to be unknown, individuals form expectations about the policymaker's "type" by using currently available observations to update their subjective prior probabilities about the government's true degree of commitment to a monetary or exchange rate target. This procedure, that is known as Bayesian updating of beliefs, describes a continuous learning process. A certain set of policy measures is assumed to signal a tough type of government whereas an other set of policies is assumed to signal a weak type. For example, observations of a high growth rate of the money supply or of the budget deficit are frequently seen as an indicator of a weak type of government, since this would result in depreciation pressure on the domestic currency. In a fixed exchange rate regime, beliefs about a weak commitment to the exchange rate parity must eventually lead to a higher interest rate spread with respect to the anchor currency.

A new strand of theoretical research focuses on ambiguous effects of policy measures on government's credibility and hence on interest rate differentials. Drazen and Masson (1994) allow for persistence of tough policies on unemployment in a two-period model that blends elements from the Barro-Gordon-type literature (Barro and Gordon 1983a, 1983b) and from the escape-clause literature initiated by Obstfeld (1997). They show that these policies can lead to a higher instead of a lower interest rate differential if agents notice that a future relaxation of these policies becomes inevitable: "If tough policies constrain the room to manoeuvre in the future, then following a tough policy may actually harm rather enhance credibility" (Drazen and Masson 1994, 744). As a result, the impact of the policies observed on interest rate spreads strongly depends on the state or regime assumed to prevail. This hypothesis stimulated substantial empirical research, that is, however, not entirely convincing.

In this paper a new way to test the hypothesis of regime-dependent determination of interest rate differentials is proposed. After a short summary of the literature building on Drazen and Masson (1994) in section two, section three develops a regime switching VAR model. In section four, regime-dependent impulse response functions are calculated showing remarkable differences in the response of European interest rate differentials to shocks in the real exchange rate and the unemployment rate across regimes. Section five finally concludes.

2 The determination of interest rate differentials

In a world of uncertainty economic agents are forced to make inferences about unknown parameters. Following the work of Backus and Driffill (1985) an extensive literature treats the type of governments, central bankers or policymakers as an unknown variable. Individuals use currently available information to form rational expectations about the true type employing the learning algorithm known as Bayes Rule. Applied to exchange rate policy, this framework implies that a policymaker who tolerates high levels of unemployment in order to maintain the given peg of the currency is seen as being a tough type. Hence, the observation of a higher unemployment rate would lower the probability of facing a weak policymaker and would reduce devaluation expectations. If the public observes that higher unemployment leads the policymaker to ease monetary policy in order to stimulate demand, this would serve as evidence of a weak type. If interest rate differentials reliably track down devaluation expectations, that is, as uncovered interest rate parity holds, these expectations and the unemployment rate are positively related.

Drazen and Masson (1994) challenge these results. They modify the model's set-up by allowing for persistence of policy effects. A tough policy on unemployment, e.g., is on the one hand seen as signaling a tough type of government but does on the other hand constrain the government's future room to manoeuvre. In other words, tolerating high levels of unemployment in order to hold the exchange rate parity raises the costs of these policies. As a result, even a tough policymaker eventually decides to devalue.

What does this modified analysis mean for the empirical relationship between policy decisions and measures of devaluation expectations? With persistence of macro-policies there does not necessarily exist a unique positive correlation between unemployment rates and interest rate differentials. Credibility of the peg might not increase with every percentage point of employment the policymaker is willing to sacrifice.

In essence, Drazen and Masson (1994) show that the interpretation of tough policies by the public is to a certain extent dependent on the prevailing state or regime of the economy. If the so-called ”signaling factor” dominates, higher unemployment means higher credibility of the peg and, hence, lower interest rate differentials. If, however, the ”external circumstances factor” dominates, unemployment and interest rate spreads are positively related. The state-dependency of the effect of policies on interest rates is the key point that motivates the application of regime-switching models in the following sections. The central question is whether we can identify different regimes empirically and whether we can track down different patterns of adjustment dynamics of the interest rate following shocks in unemployment that are depending on the regime prevailing in the economy.

The empirical research investigating the credibility of exchange rate target zones under the EMS was largely initiated by the work of Svennson (1991, 1993). He develops various techniques to extract devaluation expectations from interest rate differentials. Since he uses daily data, he can only shed light on the time series properties of this credibility measures. To measure credibility in target zones like the Exchange Rate Mechanism (ERM) of the EMS, Svennson (1993) proposes the drift-adjustment method that splits the interest differential in the expected change of the exchange rate within the band and the expected change of the central parity of the exchange rate system. However, due to practical problems, this measure is not applied widely. In accordance to the literature on fundamental determinants of credibility, this paper uses long-term interest rate differentials to approximate credibility.[1].

Following the pioneering work of Svennson (1991, 1993), later research tackles the interdependence of credibility measures with other macroeconomic variables. Caramazza (1993), Thomas (1994), Rose and Svensson (1994) and Chen

[1]De Grauwe (1994) points to the fact that the difference between raw interest rate differentials and the drift-adjusted spreads becomes smaller when long-term interest rates are used. He also applies simple interest rate spreads to measure credibility. Interestingly, Knot (1998) finds that fundamentals are as relevant for the drift-adjusted measure as they are for the raw interest rate differentials.

and Giovannini (1997) provide broad empirical studies of the determination of exchange rate credibility. Tronzano, Psaradakis and Sola (2000) extend the work of Svennson to a Bayesian framework and compute simple correlation coefficients between a credibility indicator and some macroeconomic variables. Bernhardsen (2000) explores the determination of interest rate spreads in a panel study. He finds significant effects of inflation differentials, current account positions and, which is relevant to this study, positive effects of the unemployment rate.[2] However, these studies rely on linear estimation techniques and are not able to detect the dynamics in the determination of credibility.

Closer to the present investigation are studies that look on the relationship between the unemployment rate and a measure of exchange rate credibility in more detail. De Grauwe (1994) explores correlations between the average of long term interest rate differentials of major EMS countries and the unemployment rate. He finds a positive relationship between unemployment and the credibility of the exchange rate system as a whole. Knot (1998) estimates a VAR and computes variance decompositions to show the adjustment of the interest rate differential to a shock in the unemployment rate. He finds that shocks in the unemployment rate explain a significant part of the dynamics of the spread. Hence, both authors see structural parameters of the economy as forces driving exchange rate credibility. Nevertheless, assuming a linear relationship might be misleading.

Among the first who tried to shed some light on the hypothesis of regime-dependency were Drazen and Masson (1994, 744) themselves: "...our model suggests that the partial effect of unemployment on the interest rate differential should be quite different in the different periods." Their empirical technique can give some preliminary evidence on the assumed non-linearities. They basically regress long-term government bond yield spreads between France and Germany on a measure of competitiveness of the French economy and on the French unemployment rate using OLS. Splitting the sample into three phases they incorporate dummy variables that can indicate structural breaks in the regression equation. The authors calculate the timing of breakpoints by looking at known policy changes. In addition, standard testing procedures, i.e. Andrew's (1993) likelihood ratio test, support the dates the structural breaks are assumed to occur. It is shown that the

[2] Favero, Giavazzi and Spaventa (1997) examine the time series properties of European interest rate differentials using high-frequency data.

three subperiods exhibit different relationships between the three variables. In the first period (1979:05-1982:12) and the third period (1987:01-1991:12), higher unemployment is associated with higher interest rates relative to Germany. Following the terminology of the authors, the ”external circumstances factor” dominates peoples’ perception. In the second phase (1983:01-1986:12), however, higher unemployment is associated with lower interest rate differentials indicating the dominance of the ”signaling factor”. Similar results were obtained from instrumental variables estimation.

Agénor and Taylor (1993), Masson (1995) and Agénor and Masson (1999) test the hypothesis of Drazen and Masson (1994) using more advanced empirical techniques. In employing state-space models like the Kalman filtering algorithm they lay the ground for the application of regime switching models in this paper.[3] Masson (1995) uses Kalman filter estimates for an analysis of credibility of the Pound Sterling parity under the EMS. Agénor and Masson (1999) use Kalman filtering to determine the factors that drove the credibility of the Mexican exchange rate regime in the years preceding its final collapse in December 1994.

Existing empirical research on policy credibility has so far established links between policy variables and interest rate spreads for a number of countries and various episodes. However, more empirical research is required to test the model developed by Drazen and Masson (1994). In order to derive convincing evidence in favour of time-varying effects of policy measures on credibility, an empirical model is needed that explicitly recognizes the regime-dependency of the effects presented above.

The prima-facie evidence presented by Drazen and Masson (1994) can be criticized on several grounds. First, by using appropriately specified dummy variables to account for structural breaks they separate regimes using ex-post knowledge. However, the central role of market expectations that are conditional on a certain information set at each point in time is ignored. In fact, they are not able to show the evolution or the dynamics of credibility. The regime-switching model presented below endogenously separates two regimes and tracks down regime-shifts by simulating a Bayesian learning process.

Second, their estimation can only yield single OLS coefficients that are de-

[3]Weber (1991, 1992) introduces Kalman filtering techniques to the empirical analysis of governments’ reputation and problems of exchange rate credibility.

pendent on the period specified by various dummy variables. The model used in this paper allows for a richer set of features that characterize distinct regimes with respect to various time series properties of the interest rate spreads

Third, single equation OLS estimates cannot account for the interaction of the variables. Here, the interaction of the variables is modelled as a Vector Autoregression (VAR). In this framework, the dynamic adjustment of the variables can be detected that goes beyond mere impact effects incorporated in changing OLS parameters.

The Kalman filtering technique with time-varying coefficients used in previous investigations is closely related to the regime-switching models employed in this study. However, rather than assuming a continuous state variable, the regime-switching specification formulated below assumes a discrete state variable. The model can therefore select different regimes that closely match the theoretical notion of ”regime” used by Drazen and Masson (1994). Since the application of switching VAR processes allows the adoption of standard impulse response analysis, it is particularly helpful for analyzing the evolution of credibility in response to various shocks.

3 Interest rates and regime-switching models

Modelling interest rates by means of switching regressions is still a rather novel technique that is mostly applied in assessing interest rates as indicators of business cycles, to test for term structure issues or to improve the forecasting performance of interest rate modelling.

A small number of studies applies regime-switching models to study the evolution of credibility. Gómez-Puig and Montalvo (1997) propose the conditional regime probabilities that result from the estimation of an univariate regime switching model of realignment expectations as an indicator of EMS credibility. They clearly demonstrate the deterioration of EMS-credibility surrounding the exchange rate crises. Valente (1998) shows the impact of fiscal policies in Italy on the term structure of Lira-Deutschemark interest rate differentials in a Markov-switching VAR whereas Amato and Tronzano (2000) extend the basic regime-switching autoregressive model by allowing for time-varying transition probabilities that are regressed on variables describing Italian fiscal policy. Dahlquist and Gray (2000) model weekly short-term interest rates of EMS countries in a Markov-switching framework that allows for non-linearities in the mean and the variance of the series to show the existence of different regimes that correspond to different degrees of exchange rate credibility. In addition, they extend the model and allow for time-varying transition probabilities to shed some light on the determinants of regime-shifts. It turns out that the inclusion of the level of the interest rate differentials relative to Germany and of variables that describe the position of the currency within the target zone into the regression equation of the transition probabilities significantly improves the performance of the model. However, they do not go beyond univariate analysis and, in particular, do not analyze whether a regime-sensitive relationship between interest rate spreads and macroeconomic variables can be identified. The work of Jeanne (1997) and Jeanne and Masson (2000) on the French Franc is closely related to the present study. They regress devaluation expectations on macroeconomic variables like unemployment and the real exchange rate in a regime switching specification as it is done in the model presented below. However, they allow only the intercept term of the single-equation regression to shift between two states. Since they are interested in giving evidence of the self-fulfilling nature of currency crises, they interpret these parameter shifts as sudden switches in expectations that drive devaluation expectations in a way that cannot be explained by fundamentals alone. The contribution of the present paper is to extend these models by setting up a VAR model where various parameters are allowed to switch. In this framework, it is possible to show the regime-dependent impact of macroeconomic fundamentals on devaluation probabilities.

3.1 Markov-switching models

In Markov-switching (MS) models the parameters to be estimated are allowed to switch between states. Whereas conventional time series models assume stable parameter representations over the entire sample period, regime-switching models can easily deal with structural breaks in the variables. Hamilton (1988, 1989, 1990) popularized Markov-switching models by offering convenient filtering algorithms. Since Markov-switching models can be interpreted as applying a discrete version of the Kalman filter, the basic structure follows the tradition of state-space modelling. Hence, these models are structured along the lines of a data-generating process that is supplemented by a regime- or state-generating process.[4]

Here, the Markov-switching assumption is build into a vector-autoregressive (VAR) model following Krolzig (1997, 1998). Since the vector of intercept terms $v(s_t)$, the autoregressive parameters $A(s_t)$ and the variance-covariance matrix of the innovations $\Sigma\left(s_t\right)$ of the VAR structure in equation (1) are allowed to switch, this class of models is classified as MSIAH-VAR models of order q :

$$y_t = v(s_t) + A_1(s_t)y_{t-1} + A_2(s_t)y_{t-2} + ... + A_q(s_t)y_{t-q} + u_t \qquad (1)$$

$$\text{with} \qquad u_t \sim NID(0, \Sigma(s_t)).$$

The realization of each observation y at date t depends on the regime $s_t = m$ with $m \in \{1,2\}$, hence the existence of two regimes is assumed. All realizations of the sample are collected in the vector Y_T.

As explained earlier, the Markov-switching model as a discrete version of Kalman filter state-space models relies on a law that governs the realization of the discrete state variable. Hamilton (1988, 1989, 1990) proposes the application of unobservable Markovian chains as regime-generating processes:

$$prob(s_t = j|s_{t-1} = i, s_{t-2} = k, ...) = prob(s_t = j|s_{t-1} = i) = p_{ij}. \qquad (2)$$

If the state would be observable at each point in time, dummy variables could be used to condition the parameter-estimates on a certain state. In the class of models applied here, however, the regime that prevails at time t is unobservable. The Markov property described in equation (2) simply says that the probability of a state m at time t, i.e., $s_t = m$, only depends on the state in the previous period, s_{t-1}. The probability p_{ij} is called transition probability and says how likely state i will be followed by state j. Collecting the transition probabilities in a (2×2) matrix gives the so-called transition matrix **P**:

$$\mathbf{P} = \begin{bmatrix} p_{11} & 1 - p_{22} \\ 1 - p_{11} & p_{22} \end{bmatrix}, \qquad (3)$$

where the element of the i-th row and the j-th column describes the transition probability $p_{ij.}$ If exactly one eigenvalue of the transition matrix is equal to unity

[4]See Kim and Nelson (1999) and Franses and van Dijk (2000) for accessible textbook treatments of Markov models

and all other eigenvalues lie inside the unit circle, the Markov-process is said to follow an ergodic Markov chain. A special case are chains that end in an absorbing state after which the chain stops (a so-called reducible Markov chain). In this case one element of the (2×2) Matrix is equal to zero so that the matrix can be reduced to a triangular matrix. For the use of Markov chains as underlying regime-generating processes of stationary VAR models the property of ergodicity and non-reducibility are essential.

The transition probabilities are assumed to be constant. In an interesting exercise these probabilities could be formulated as time-varying transition probabilities that could then be regressed on a set of macroeconomic fundamentals following Diebold, Lee and Weinbach (1994). However, this extension is not pursued here but is left for another study.

Since the state variable is assumed to be unobservable, the estimation procedure is based on the iterative Baum-Lindgren-Hamilton-Kim-filter (BLHK-filter, see Krolzig (1997)) that infers the regime-probabilities at each point in time. At the centre of this estimation procedure lies a Bayesian learning procedure. Agents are assumed to update their probability assessment using the information entailed in each subsequent observation. Since the Bayesian updating of individual beliefs incorporated in the estimation closely matches agents' rational expectations inferences, Markov-switching models are particularly useful to study the dynamics, that is, the evolution of credibility.

As a by-product of the filter-inferences, a likelihood function can be set up and maximized in order to obtain parameter estimates of the MSIAH-VAR model. The likelihood function $L(\theta|Y_T)$ is given by the sum of the densities $f(.)$ of the observation y_t conditional on the history of the process Y_{t-1}

$$L(\theta|Y_T) = \sum_{t=1}^{T} f(y_t|Y_{t-1};\theta) \tag{4}$$

with

$$\begin{aligned} f(y_t|Y_{t-1};\theta) &= f(y_t, s_t = 1|Y_{t-1};\theta) + f(y_t, s_t = 2|Y_{t-1};\theta) \\ &= \sum_{m=1}^{2} f(y_t|s_t = m, Y_{t-1};\theta)\ prob(s_t = m|Y_{t-1};\theta) \end{aligned} \tag{5}$$

where the second part of this expression follows from applying the rules of conditional probabilities which say that $f(y_t, s_t = m|...) = f(y_t|...)\ prob(s_t = m|...)$. The non-linear EM algorithm (Expectation-Maximization algorithm) is applied to solve the problem:

$$\widehat{\theta}_{ML} = \arg\max \ln L(\theta|Y_T), \tag{6}$$

where the vector θ includes the VAR-parameters to be estimated. What is most important to keep in mind is that the attractiveness of this class of Markov-switching models lies in its ability to estimate the dating of the regimes and the regression parameters simultaneously. Since the model endogenously selects two distinct regimes and tracks down conditional regime probabilities at each point in time, no a priori knowledge about the dates of the regime shifts is necessary. Rather, the model lets the data determine which realization of the series was generated under what regime.

3.2 EMS spreads in a MSIAH-VAR(q) model

In this section, interest rate spreads of major EMS countries are modelled using regime-switching processes. In particular it is assumed that each country's experience can be described by a VAR that allows for shifts in all parameters as described by equation (1).

For each country i, the MSIAH-VAR system consists of three variables:

$$y_t^i = \left(tcreer_t^i, tcur_t^i, \Delta(i_t^i - i_t^{DM})\right)' \tag{7}$$

where the temporary component of the real effective exchange rate ($tcreer_t^i$), the temporary component of the unemployment rate ($tcur_t^i$) and the first difference of the interest rate differential relative to Germany ($\Delta(i_t^i - i_t^{DM})$) are included. The model comprises three variables in order to remain comparable to the original results of Drazen and Masson (1994) presented above and to limit the number of parameters to be estimated. An obvious extension of the approach taken here would be to include other fundamentals like budget deficits, the level of the public debt etc.

In analogy to Drazen and Masson (1994) long term interest rates are used that are obtained from the IMF's *International Financial Statistics* database. Here the interest rates on long-term government bonds are included for reasons of compatibility since the existing literature mostly focuses on long term rates rather than on short term spreads, see Drazen and Masson (1994) and Masson (1995). The real effective exchange rate was provided on J.P. Morgan's website (http://www.jpmorgan.com). A higher exchange rate means an appreciation of the domestic currency. Here, the real exchange rate is meant to be a measure of competitiveness as an overvalued exchange rate is frequently seen as dampening exports and making an expansionary nominal devaluation more likely. The unemployment rate used in this model is either the standardized unemployment rate or the percentage of unemployed relative to the total labour force as given by the OECD's database. Since the hypothesis of a unit-root cannot be rejected by applying standard augmented Dickey-Fuller tests, estimating the VAR in levels is not appropriate. Hence, each series is modified in order to get stationary variables. For the interest rate spread, simple difference-stationarity is assumed. The unemployment rate and the real effective exchange rate are modelled as trend-stationary series. A standard Hodrick-Prescott filter is used to extract a linear trend and to identify the temporary components of both variables. The raw data used in this study is displayed in figure 1, figure 2 and figure 3 in the appendix.

The model is estimated separately for major EMS countries over a sample period that runs from the early days of the ERM to the first months of EMU in 1999.[5] All data is at monthly frequency.[6] Hence, this study covers the whole period the EMS was operating or exchange rates were actively managed, respectively. This is particularly interesting since the convergence process induced by the Maastricht treaty and the scheduled start of European Monetary Union in 1999 is likely to alter the structural relationships between the variables in favour of a more credible exchange rate policy. In order to secure a sufficient number of degrees of freedom and to reduce the already heavy computational requirements, a parsimonious modelling approach is helpful. Therefore, the number of lags included in the VAR model is set to $q = 3$. For each estimation the number of regimes is restricted to two. Trials with more than two regimes failed due to the large number of param-

[5]In particular, the sample covers the following periods: 1979:01-1999:03 for France and the UK and, for reasons of data availability, 1982:01-1999:03 for Italy.

[6]The BLHK filter, the maximum likelihood estimation and the EM algorithm are implemented using the MSVAR procedure which is written in Ox 2.0 and was developed by Krolzig (1998).

eters to be estimated or gave highly implausible results.

If the hypothesis of Drazen and Masson (1994) is correct, one would expect the VAR model to determine two states that are characterized by distinct time series properties of the interest rate spread and different patterns of interaction between the variables included. Regime-dependent determination of exchange rate credibility would imply that shocks in the unemployment rate or the real exchange rate affect the spread in different ways depending on the regime prevailing. The most interesting results and some diagnostic tests are reported in table (1).

The non-linear specification of the three-variable VAR system yields a higher maximum of the log-likelihood function for each sample country than the linear VAR does. This max. ln $L(.)$ can be interpreted as a measure of the model's goodness of the fit for the maximum likelihood estimator represents the value of the model's parameters for which the sample is most likely to have been observed. To test the quality of the nonlinear model against the corresponding linear VAR, likelihood-ratio tests (LR tests) are usually applied that are asymptotically $\chi^2(r)$ -distributed with r degrees of freedom. However, the LR test under normal conditions does not apply here due to the existence of unidentified nuisance parameters under the alternative (the transition probabilities are not identified under the linear model). The use of the standard χ^2 distribution would therefore cause a bias of the test against the null.[7] To circumvent the problem of unidentified nuisance parameters, an overly cautious approach is used in this study. This means that the LR test statistic is compared to a $\chi^2(r+n)$ distribution where n stands for the number of nuisance parameters. Since the test statistic exceeds the critical value even under this overly conservative benchmark, the null-hypothesis can be rejected at high significance levels. An alternative to this procedure is offered by Davies (1977) who adjusts the test statistic by deriving an upper bound for the significance level, see Krolzig (1997). The results remain highly significant under this modified test. Thus, a non-linear regime switching specification seems to be not only appropriate but rather superior to conventional linear models.

Figures (4), (6) and (8) in the appendix show the conditional regime probabilities obtained from Hamilton's estimation method. Here, only the smoothed regime probabilities are reported, that is, the probability assessment based on the information set y_T that comprises all observations in the sample. In all countries,

[7] See Hansen (1992), Andrews and Ploberger (1994) and Garcia (1998) for this problem.

the model detects regime shifts that occur at dates that do not always correspond -at least in the case of France- to the splited sample period in the specification of Drazen and Masson (1994). These conditional regime probabilities as well as the reported transition matrices and the expected durations (see table 1) make obvious, that the regimes are characterized by different degrees of persistence. Whereas, e.g., regime one and regime two in France last for more than two years each, they last for one and a half years and nine months in the UK.

Each country's results will be interpreted in more detail below after the framework for the analysis of the dynamic adjustment following shocks is elaborated in the next section.

How are the regimes characterized and in what respect do they differ? It is important to note that the model selects the regime characteristics in order to fit the specification efficiently to the data. However, it turns out that the regime properties can be interpreted reasonably well. Due to the large number of parameters only the vector of intercept terms is reported here as an example. Take, for example, the United Kingdom. The model selects regime one as being marked by an undervalued real exchange rate (relative to the trend), an unemployment rate that lies below its trend and a narrowing interest rate spread, whereas regime two shows an overvalued real exchange rate, above-trend unemployment and a widening interest rate differential:

$$\begin{aligned} v^{UK}(s_t = 1) &= (-0.215, -0.008, -0.012)' \\ v^{UK}(s_t = 2) &= (0.282, 0.002, 0.043)'. \end{aligned}$$

Similar characteristics result for the other sample countries.[8] To gauge the dynamic interaction between the variables in each regime and especially between macroeconomic variables and the credibility proxy, the following section computes regime-dependent impulse response functions.

[8]The numbering of regimes is done rather arbitrarily by the algorithm. Notwithstanding the classification of being regime one or two, the time series properties remain the decicive characteristics of the regimes.

4 Regime-dependent impulse response functions

The interpretation of the set of the 74 VAR parameters to be estimated is not straightforward. For this reason, the profile of the system's response to shocks is usually derived to visualize the dynamics represented by the VAR model. These so-called impulse response functions are a very convenient way to track down the magnitude and the persistence of each variable's response to economic shocks over time. In the VAR literature shocks are understood as being unexpected changes or innovations in one of the system's variables.

For the purpose of this paper the response patterns within a certain regime are most important. Here, the technique developed by Ehrmann, Ellison and Valla (2000) is used. They derive impulse response functions that are regime-dependent. Since each regime is characterized by distinct time series properties of the observable variables, these sets of parameters can be used to study the dynamic adjustment of the system to an unexpected Gaussian innovation.[9]

The MSIAH-VAR can be written in moving-average (MA) form as:

$$y_t = v(s_t) + u_t + \psi_1 u_{t-1} + \psi_2 u_{t-2} + ... \quad (8)$$

$$\text{with} \quad \psi_k = \frac{\partial y_{i,t+k}}{\partial u_{jt}} \,|_{s_t=...=s_{t+k}=m} \text{ for } \; k \geq 0. \quad (9)$$

A plot of the row i, column j element of $\psi_k$as a function of the response's time horizon k is called impulse response function. It describes the response of $y_{i,t+k}$ to an orthogonal one-time impulse in u_{jt} in regime m with all other variables dated t or earlier held constant.

In order to interpret the impulse functions reasonably, the structural shocks that drive the VAR dynamics must be exactly identified, see Sims (1980). For this reason, a triangular identification scheme known as Choleski decomposition is employed. This is done by imposing an order of the variables onto the system which implies that each variable has contemporaneous effects only on itself and on variables ordered below it. Ordering the variables as given by the vector y_t in

[9]Ehrmann (2000) uses the same technique to study asymmetries in the monetary transmission process. He models business cycle phases as the realization of an unobservable Markov-switching process. Kakes (2000) also shows asymmetric effects of monetary policy over the course of the business cycle. However, he does not provide evidence of the significance of the effects.

(7) means that shocks to the real exchange rate affect both the unemployment rate and the interest rate differential, but that shocks to the unemployment rate affect the interest rate spread only. Let Ω denote a lower triangular matrix such that $\Sigma(s_t) = \Omega\Omega^{-1}$. The innovations are then given as a function of the underlying structural shocks ε_t by $u_t = \Omega\varepsilon_t$. As a result of this ordering, the matrix $\Sigma(s_t)$ is exactly identified.

In this regime-dependent VAR set-up the impulse response functions are calculated separately for each regime m. It is assumed that the given regime prevails throughout the duration of the response, that is $s_{t+k} = m \, \forall k > 0$. Here, shocks in $tcreer$ and $tcur$ are considered. Depending on the regime prevailing when the shock occurs, the system should respond in different ways to these macroeconomic shocks. All impulse response patterns show the effect of a single shock of one percentage point over a time horizon of $k = 18$ months. The confidence bands are computed employing standard bootstrapping algorithms following Ehrmann, Ellison and Valla (2000) who kindly provided their software. Artificial histories for the variables and for the unobservable state variable are created by drawing random numbers from appropriate distributions and inserting these numbers into the MS-VAR system using the estimated parameter values. The model is then estimated with these artificial data. Replicating this procedure 1000 times gives a sufficiently precise distribution of the estimates that allows an assessment of the significance of each impulse response.

The regime-dependent impulse response functions are shown in figures (2), (4) and (6) in the appendix. As the most striking result from this analysis it occurs, that the responses of the interest rate spread to various shocks are indeed different across regimes. It is important to keep in mind that what is important in this study is not so much the significance of each country's reaction to shocks but rather the difference of these reactions across regimes. The insignificance of some impulse response patterns when analyzed in isolation is in line with the literature since most authors including Drazen and Masson (1994), Jeanne (1997) and Jeanne and Masson (2000) only find insignificant relationships between spreads and macroeconomic variables. However, as non-monotonic relations between spreads and macroeconomic data are the focus of this paper, the regime-dependency of the impulse responses supports the hypothesis of regime-dependent determination of exchange rate credibility.

In France, regime one corresponds to the Drazen-Masson notion of a signaling regime explained above. The interest rate spread relative to Germany decreases in response to shocks in the real exchange rate and the unemployment rate, hence, credibility is strengthened, see figure 5. In regime two on the contrary, which is the external circumstance regime in the terminology of Drazen and Masson (1994), the interest rate spread reacts positively to these shocks. Hence, credibility deteriorates because individuals perceive a devaluation as more likely. As figure 4 makes clear, the signaling regime prevailed during the 1990s. During that period, exchange rate stability was a prerequisite for the realization of European Monetary Union. Only during the turbulences in 1992/1993 that forced the devaluation of several EMS currencies, regime two prevailed. With the exception of 1982/1983, the 1980s saw regime two dominating with credibility deteriorating following higher unemployment and a stronger real exchange rate. The shift in 1982/1983 corresponds to the results of Drazen and Masson (1994). They refer to the announcement of a *"politique de rigueur"* [italics in original] of the newly elected socialist government in France, after its initial Keynesian macroeconomic policy had failed.

In Italy, regime one is characterized by a negative reaction of interest rate differentials to macroeconomic shocks, whereas regime two shows a positive reaction (figure 7). However, the shapes of the impulse response functions are not as striking as in the case of France. The switches occur very often reflecting the uncertainty of financial markets with regard to the stance of Italian economic policy (figure 6). In the second half of the 1990s, regime one clearly prevailed as the Maastricht convergence process required exchange rate stability and monetary policy discipline.

In the UK, regime two can be clearly identified as being the signaling regime (figure 9). Regime two was most likely to govern the VAR process in the beginning and the end of the 1990s and in some episodes in the 1980s (figure 8).[10] Since the UK decided to stay out of EMU for the time being, the shift towards a signaling regime occurring around 1997 cannot be explained as being induced by the Maastricht convergence process. An alternative explanation could be the decision

[10]Estimating the model through 1992 when the UK left the exchange rate mechanism results in comparable regime probabilities and similar impulse response functions. However, the likelihood ratio test statistic is significant only at the 10 percent level.

of the newly elected Labour government to grant the Bank of England independence in setting monetary policy. This shields monetary policy from political influence and forces the central bank to gain reputation. As in the case of France in 1982/1983, the newly elected government in the UK in 1997 and the newly formed government in 1990 are facing a high degree of uncertainty and, hence, try to establish some reputation for being tough on unemployment and monetary policy. As a result, financial markets considered a signaling regime most likely to govern the interest rate differential relative to Germany.

In general, the hypothesis of regime-dependent determination of exchange rate credibility can be supported by the MS-VAR model. It turns out that these results are fairly robust with respect to the ordering of the variables, the number of lags included and the choice of the technique employed to obtain stationary series. In addition, running the model with short term (three months) euro-rates instead of long-term government bond yields leads to similar results although the conditional regime probabilities show an implausible degree of persistence of each regime. An applications of the present model to smaller EMS countries is left for further research.

5 Conclusion

In this paper a three-dimensional regime switching VAR system is estimated in which the vector of intercept terms, the matrices of the autoregressive parameters and the variance-covariance matrix of the innovations are allowed to switch across states. In order to test the hypothesis of regime-dependent interaction between interest rate spreads and various macroeconomic shocks, the VAR model is fitted to the interest rate differential, the real exchange rate and the unemployment rate for France, Italy and the UK. The model endogenously selects the dates of the regime shifts and distinguishes between two remarkably different regimes. Regime-dependent impulse response functions that are obtained from a Choleski-decomposition of the variance-covariance matrices reveal substantial differences in the response patterns of the interest rate differential.

For all countries, the model identifies one regime in which the interest rate spread reacts positively to a shock in the real exchange rate and the unemployment rate and one regime in which the interaction has a negative sign. These regimes correspond to the notion of Drazen and Masson (1994) who refer to an external circumstance regime in the first case and a signaling regime in the latter. Hence, this paper provides some evidence on the regime dependent nature of exchange rate credibility.

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A Tables and figures

Table 1: Results from maximum-likelihood estimation of the MSIAH-VAR(q) model

	France	Italy	UK
max. ln $L(\theta\|Y_T)$:			
linear VAR(q)	-16.2	-221.8	-197.9
MSIAH-VAR(q)	49.9	-163.9	-162.2
LR test statistic	132.3	115.9	71.4
p values:			
adjusted χ^2	0.00	0.00	0.00
bounded LR	0.00	0.00	0.00
average expected duration:			
$s_t = 1$	35.4	6.5	17.7
$s_t = 2$	26.5	5.8	8.8
p_{11}	0.97	0.85	0.95
p_{22}	0.96	0.83	0.82

Notes:The LR test statistic is computed as $LR = 2(\ln L(\theta|Y_T) - \ln L(\theta^{restr}|Y_T))$ where θ^{restr} denotes the set of parameters obtained from an estimation of the restricted (linear) VAR model. To circumvent the problem of unidentified nuisance parameters, the test statistic is compared to an adjusted $\chi^2(74 + 2)$ distribution where the number of restrictions is supplemented by the number of nuisance parameters. Alternatively, the method of Davies (1977) is used who derives a bounded test statistic. The expected duration of regime m is equal to $\frac{1}{1-p_{mm}}$; the reported durations are the exact numbers whereas the probabilities are rounded.

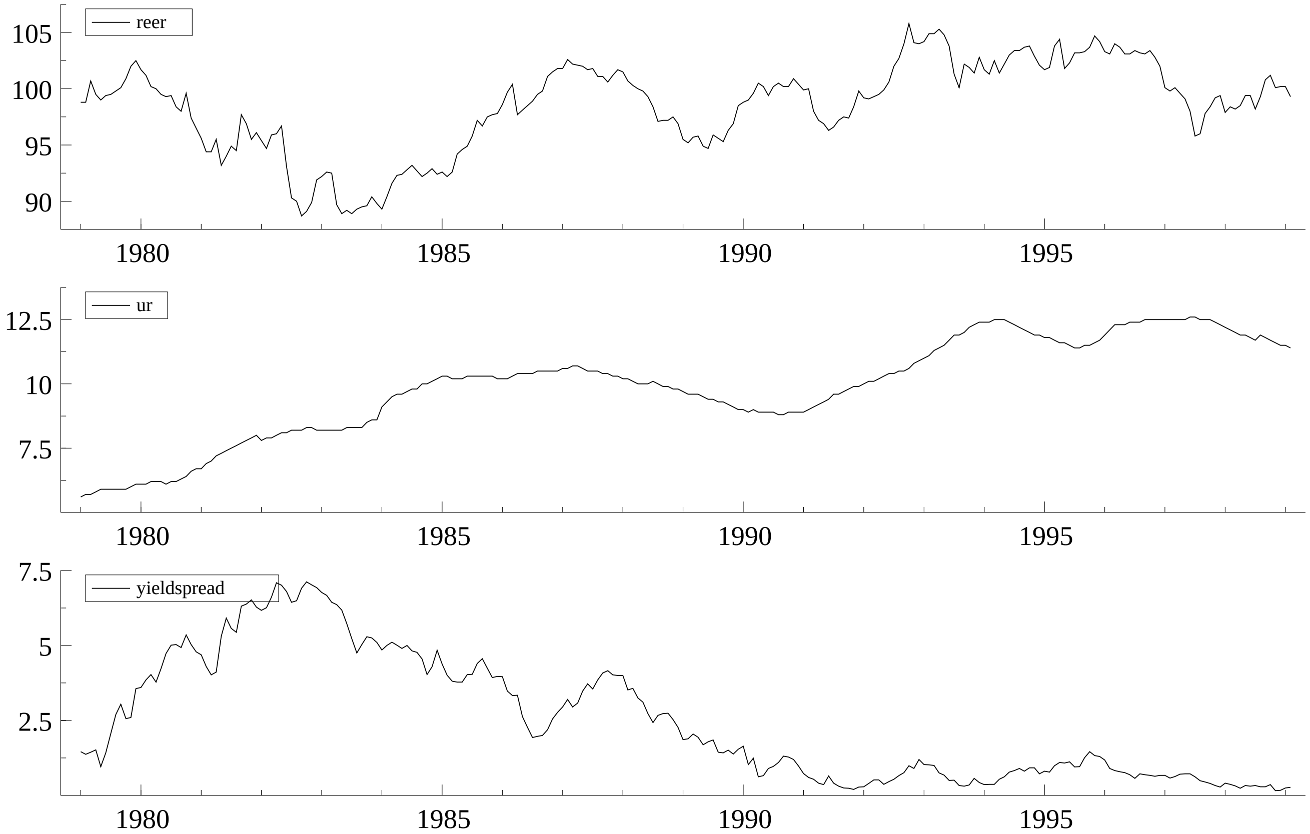

Figure 1: Italy-Data series

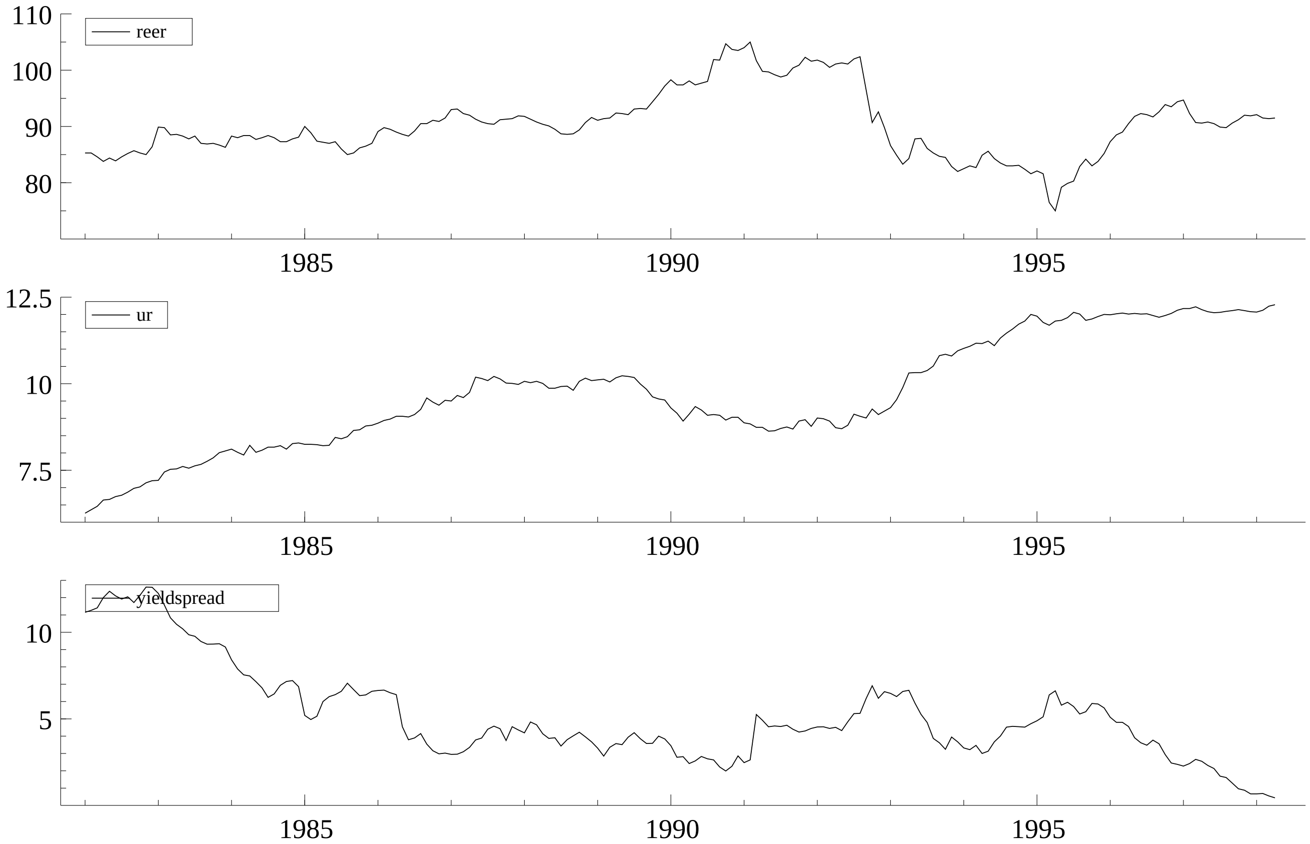

Figure 2: Italy-Data series

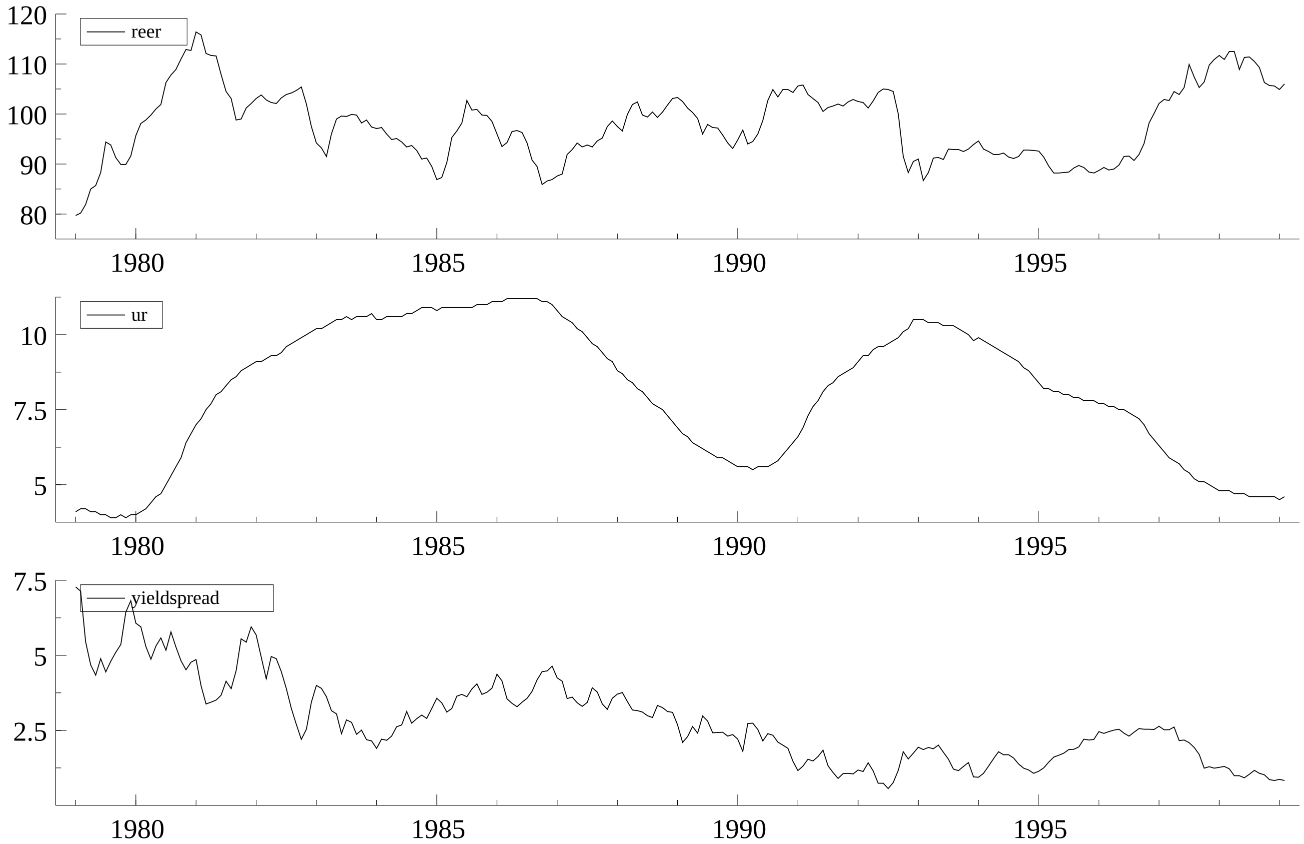

Figure 3: UK-Data series

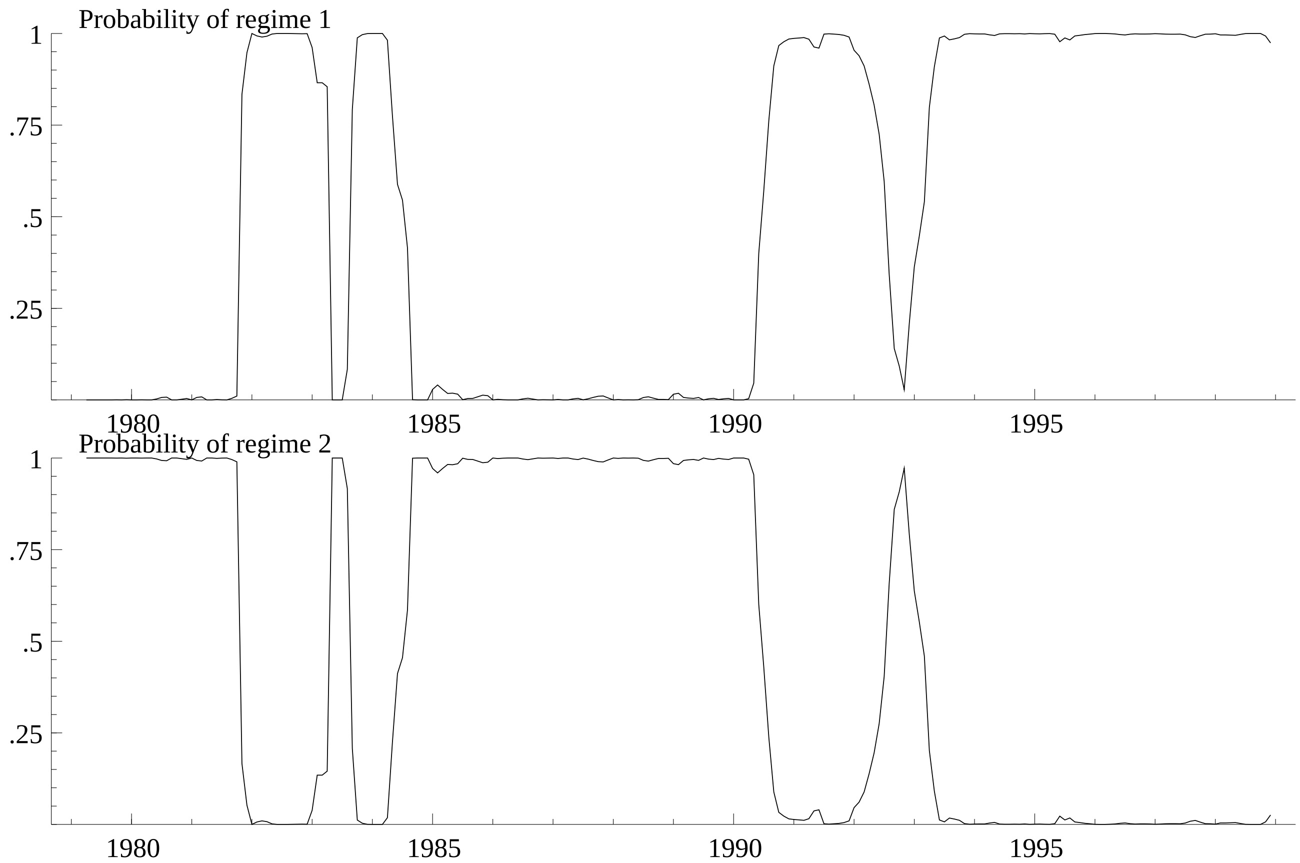

Figure 4: France — Smoothed regime probabilities $prob(s_t = m|Y_T), m \in \{1, 2\}$

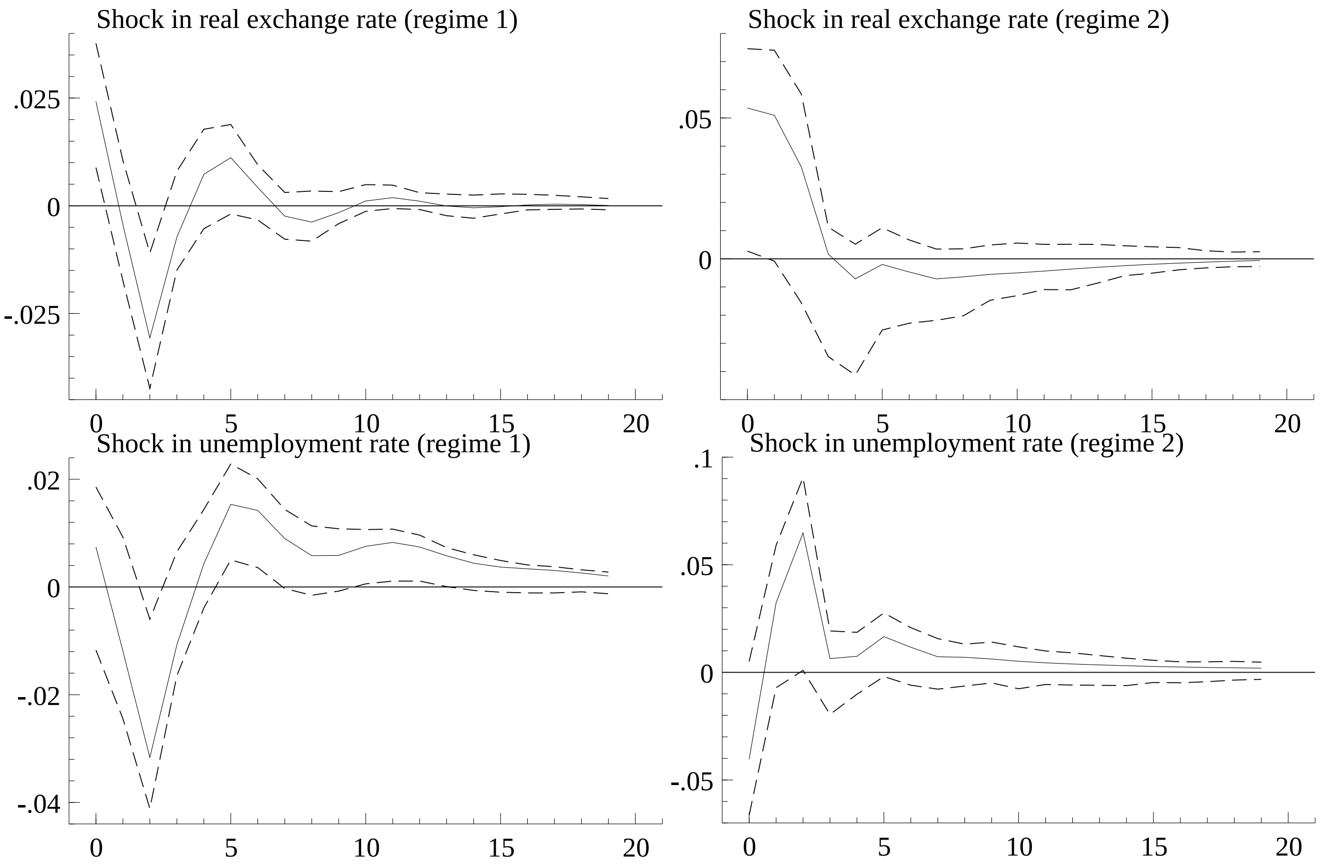

Figure 5: France - Response of interest rate differential to shock of one percentage point in the real exchange rate (upper panel) and the unemployment rate (lower panel) $\pm$ 1 standard deviation

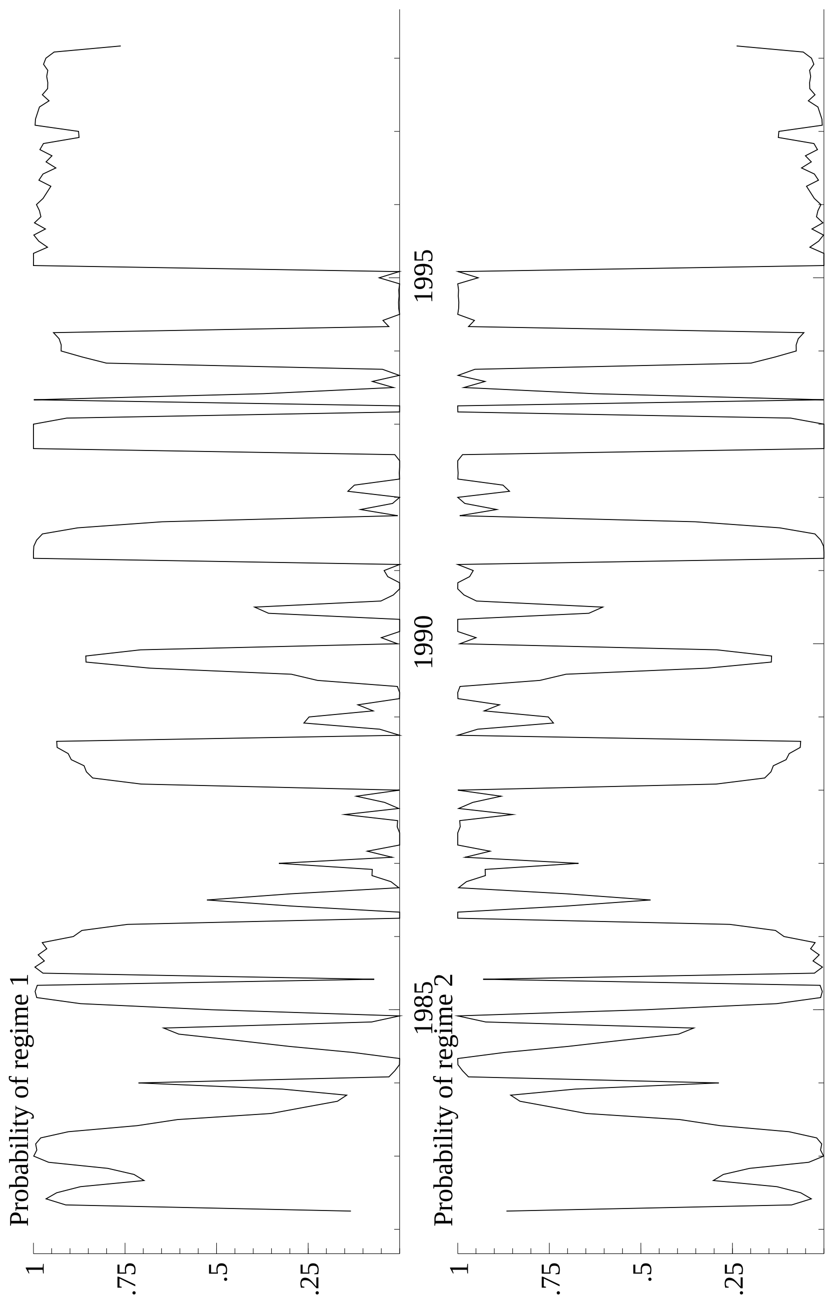

Figure 6: Italy - Smoothed regime probabilities $prob(s_t = m|Y_T), m \in \{1, 2\}$

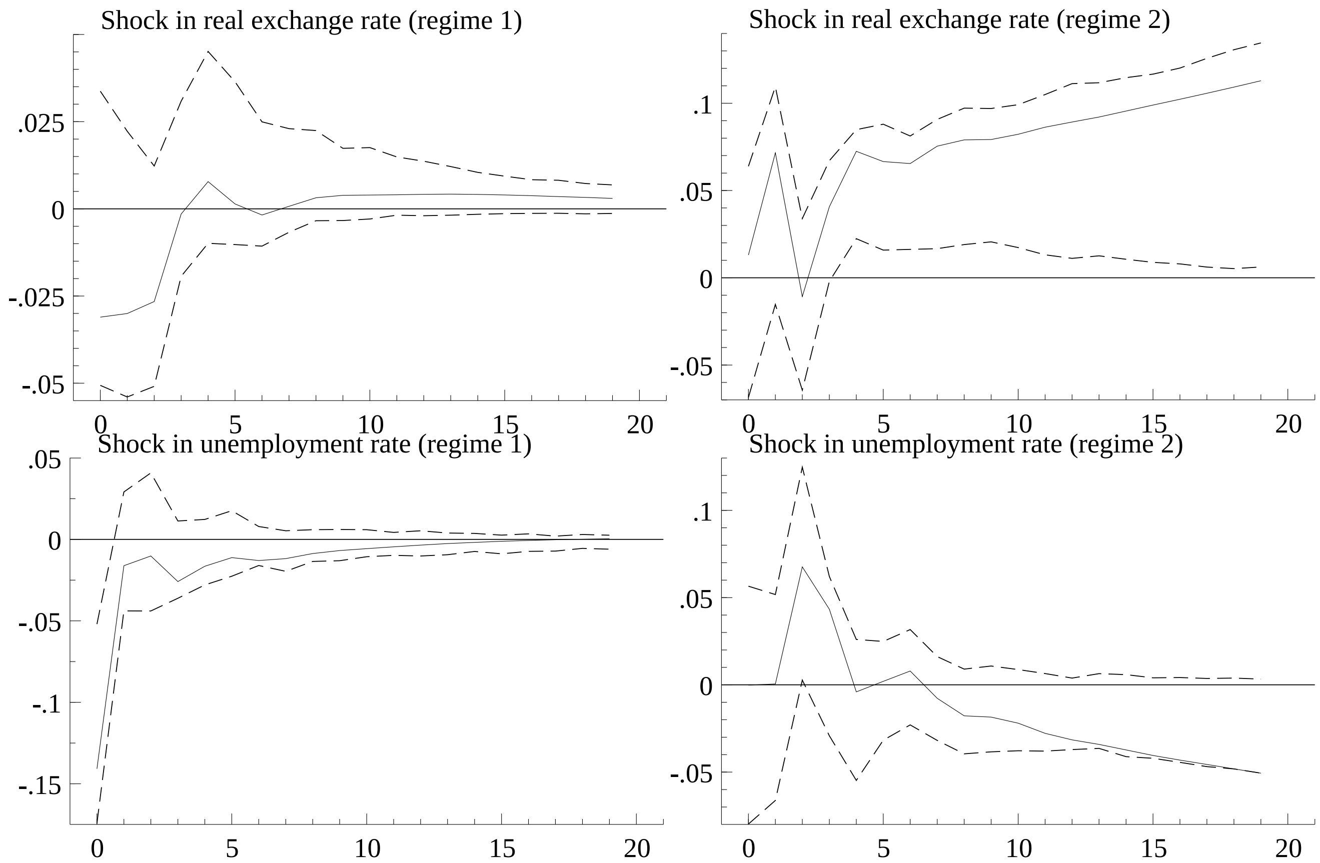

Figure 7: Italy - Response of interest rate differential to shock of one percentage point in the real exchange rate (upper panel) and the unemployment rate (lower panel) $\pm$ 1 standard deviation

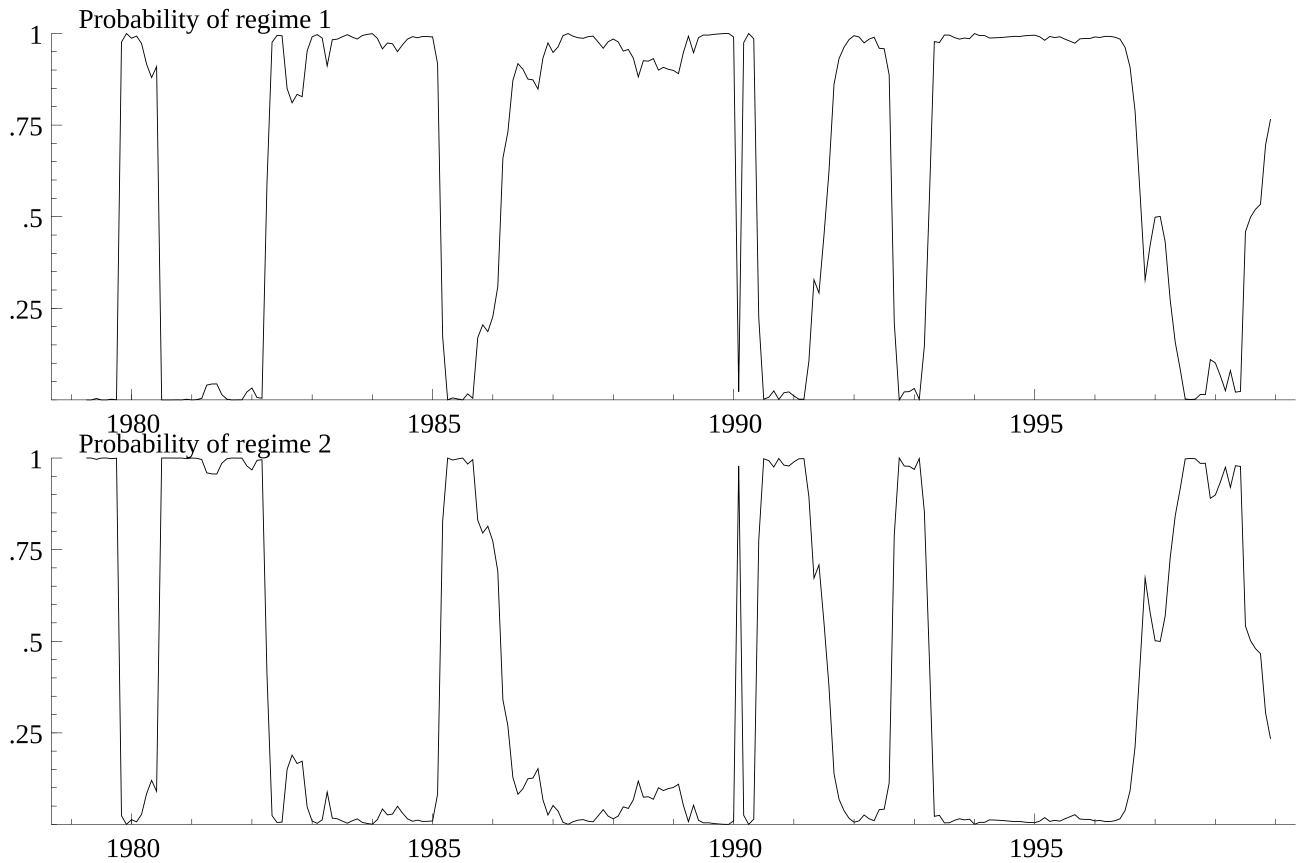

Figure 8: UK - Smoothed regime probabilities $prob(s_t = m|Y_T), m \in \{1, 2\}$

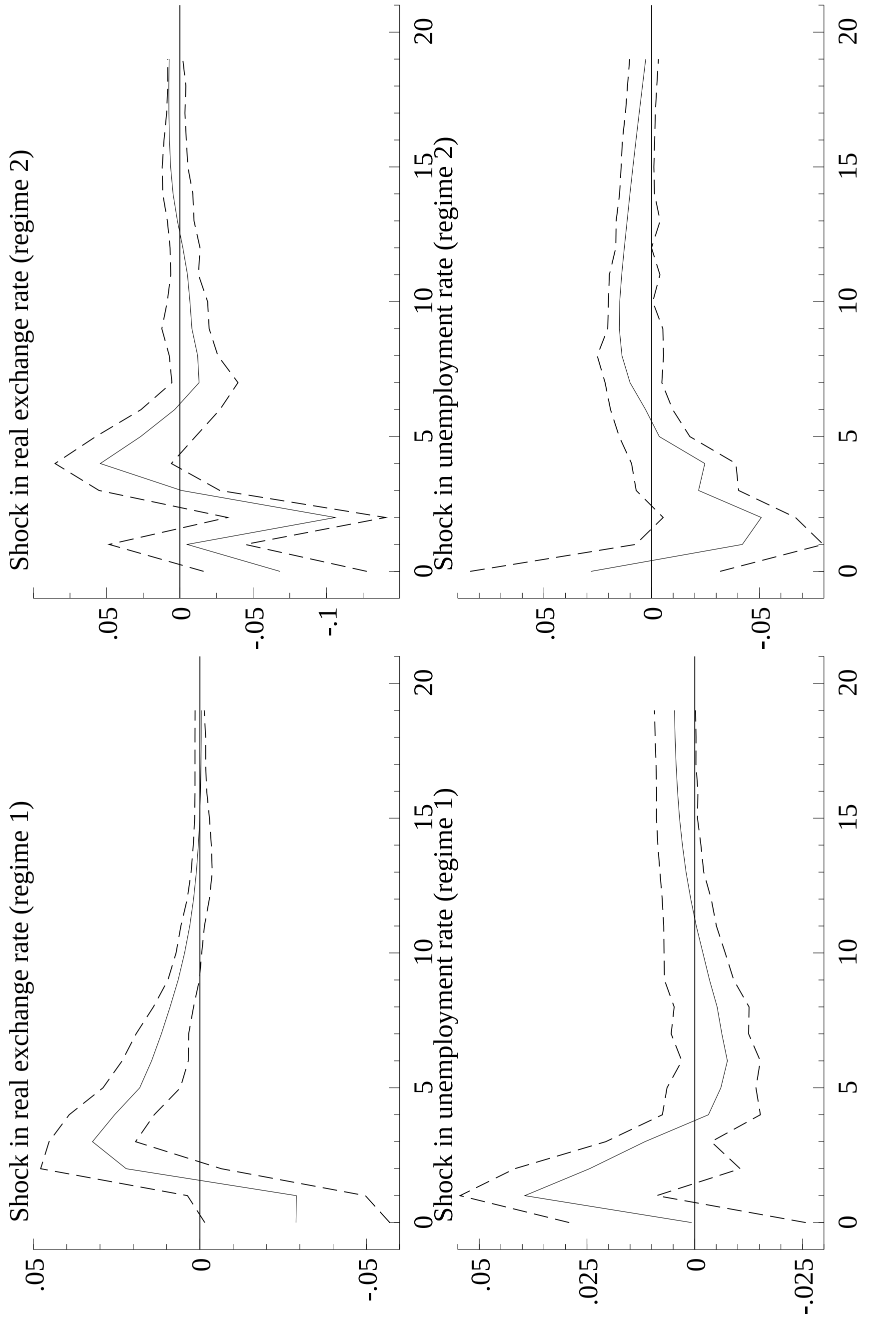

Figure 9: UK - Response of interest rate differential to shock of one percentage point in the real exchange rate (upper panel) and the unemployment rate (lower panel) $\pm$ 1 standard deviation