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November 2007

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Abstract

This paper analyzes the transmission of inflation across the five largest economies in the European Monetary Union, i.e. France, Germany, Italy, Netherlands and Spain. We use monthly CPI inflation rates for the period 1970-2006. Given the long observation period and the continuing economic integration of Europe's economies, we first try to investigate, if there were changes in inflation dynamics in these countries using univariate Markov-switching models. To assess the inflation transmission mechanism, we first establish a long-run relationship between the five countries using cointegration methods. As implied by the results of the univariate models, we allow for changes in the adjustment coefficients of the cointegrating relationships and the short-run dynamics. Using a Markov-switching vector error correction model we find evidence for multiple regime switches during the early 1970s till the mid 1980s. Exactly during this period we find evidence for Germany being weakly exogenous, which highlights the dominance of German monetary policy at this time. Since the mid-1980s we find evidence for a stable transmission mechanism both in the long- and the short-run characterized by a low degree of inflation persistence.

Keywords: Inflation transmission, monetary integration, MS-VECM, cointegration, euro area

JEL classification: E30, E31, E50

*This research was supported by the Deutsche Forschungsgemeinschaft through the SFB 649 "Economic Risk".

†I am grateful to Jürgen Wolters, and Enzo Weber for their help and comments. All remaining errors are mine, of course.

1 Introduction

1.1 Motivation

A huge strand of economic literature has dealt with the question of optimal monetary policy, particularly since the introduction of the euro as the common currency in the European Monetary Union (EMU). One result of this field of research is that price rigidities across regions determine the optimal behavior of monetary policy in a currency union such as the EMU. Price rigidities in turn determine inflation persistence or more generally inflation dynamics. Recent research has focused on estimating country-specific price rigidities within the EMU using various methods ranging from simple univariate models for aggregate data to disaggregated models measuring product- and country-specific frequencies of price adjustments with heterogeneous results.

Besides the knowledge of price-setting behavior across countries and sectors a central bank like the ECB needs to know how inflation dynamics differ across regions within the monetary union. Further, a central bank in a monetary union needs to have sufficient knowledge about the transmission of inflation across its member states in order to conduct monetary policy optimally.

This paper analyzes the transmission of inflation across the five largest economies in the EMU, i.e. France, Germany, Italy, Netherlands and Spain. We use monthly CPI inflation rates for the period 1970-2006. Given the long observation period and the continuing economic integration of Europe's economies, we first try to investigate, if there were changes in inflation dynamics in these countries. Using a univariate Markov-switching model, we obtain evidence for an almost simultaneous regime change in all five countries during the mid-1980s. To assess the inflation transmission mechanism, we first establish a long-run relationship between the five countries using cointegration methods. As implied by the results of the univariate models, we allow for changes in the adjustment coefficients of the cointegrating relationships and the short-run dynamics. Using a Markov-switching vector error correction model (MS-VECM) we find evidence for multiple regime switches during the early 1970s till the mid 1980s. Exactly during this period we find evidence for Germany being weakly exogenous, which highlights the dominance of German monetary policy at this time. Since the mid-1980s we find evidence for a stable transmission mechanism both in the long- and the short-run. Further, the analysis shows that the decrease of inflation persistence in the euro area, which is found in recent studies such as Altissimo, Ehrmann and Smets (2006), has its origin in French and German inflation dynamics and their importance for the transmission process[1].

1.2 Literature Review

Country-specific inflation dynamics and persistence are basically determined by country-specific price rigidities. For this reason a thorough understanding of the patterns and determinants of inflation persistence is important for policy-makers, as inflation persistence has immediate consequences for the conduct of monetary policy, e.g. the appropriate response to a rise in inflation depends on the degree to which the shock itself is persistent (cf. Altissimo, Ehrmann

[1] In contrast to Altissimo, Ehrmann and Smets (2006) we find evidence that the decline in inflation persistence has already taken place during the 1980s.

and Smets 2006). Country-specific inflation dynamics are in turn affected by inflation dynamics in other countries, particularly in a currency union such as the EMU with highly integrated economies. From this it follows that a country's inflation is not only affected by its own inflation persistence but also by other countries' inflation persistence.

Univariate models using the sum of autoregressive coefficients as a measure of persistence, as proposed by Andrews and Chen (1994), offer a simple way to estimate inflation persistence both at an aggregated and disaggregated level. Applications of this approach to EMU data, including both aggregated and country-specific estimations, may be found in Gadzinski and Orlandi (2004), Batini (2006) and Levin and Piger (2004). The results of these contributions substantially differ depending on the inflation definition chosen, i.e. annual, monthly, the proxy chosen for the price level, i.e. CPI, HCPI or GDP deflator, and the time frame of the analysis. Another reason for the heterogeneity of the results may be due to country-specific restrictions of the estimates, as univariate models do not allow for interactions of inflation dynamics across countries. Instead it is more appropriate to assume, as mentioned above, that a persistent inflation shock in country A will lead to a persistent run of inflation in country B, if both A and B are highly integrated. Univariate models would then suggest that inflation is persistent in both country A and B, although the origin of inflation persistence arises from country A.

The transmission of inflation is analyzed in few empirical papers. Motivated by the period of high inflation in Western countries during the 1970s Darby et al. (1983) and Darby and Lothian (1989) were among the first who analyzed the transmission of inflation across countries empirically. In particular, they investigate the sources and origins of high inflation in the industrialized West with a focus on the economic linkages between these countries, while the quantification of the inflation transmission mechanism itself is not considered. Yang, Guo and Wang (2006) investigate the international transmission of inflation among G-7 countries using a VECM approach. They show that U.S. inflation has a less dominant role than is usually assumed. Generally, they find a broad linkage of inflation among G-7 countries for the period 1973-2003. Cheung and Yuen (2002), who assess inflation dynamics across the U.S., Hong Kong and Singapore, find evidence for the inflation rates in the small economy being caused by the large economy.

In this paper we analyze how inflation is transmitted across euro area's member states. Starting with univariate models, which offer a simple way to measure inflation persistence[2], which in turn determine inflation dynamics, we first investigate, if one can find evidence for parameter instabilities in the sample. We then try to establish a long-run relationship between the five countries' inflation rates, as implied by the relative purchasing power parity. Given that there exists at least one cointegrating relationship, we formulate a VECM in the five inflation rates. In contrast to the existing literature, we allow for changes in the short-run dynamics and the adjustments coefficients. We then compute impulse response for each of the regimes of the MS-VECM.

[2] Of course, this simpleness has its costs, as interactions in inflation dynamics between countries cannot be considered. More sophisticated models such as structural ones, would probably deliver estimated parameter that differ from the ones given in the next section. Nonetheless, we think that for the purpose of assessing parameter changes, it is sufficient to use a reduced form univariate model.

2 Analysis

2.1 Evidence for Changes in Inflation Dynamics from Linear Univariate Models

2.1.1 Methodology

As a first step in the analysis we start with estimating simple univariate autoregressive (AR) models for each of the five countries. We do this to investigate, if regime shifts have possibly played a role in describing the inflation persistence in the euro area, as this would in turn have an impact on the inflation dynamics within a country and most likely for the transmission process in the entire euro area.

We follow the line of Andrews and Chen (1994) and use the sum of AR coefficients as a measure for persistence of a time series. That means we start with an AR(p) model of the form

$$\pi_{i,t} = c + \alpha_{1,i}\pi_{i,t-1} + \alpha_{2,i}\pi_{i,t-2} + \ldots + \alpha_{p,i}\pi_{i,t-p} + \nu_t, \tag{2.1}$$

where $\pi_{i,t}$ denotes the rate of inflation in country i at time t, c represents a constant and ν_t is an *i.i.d.* error term. (2.1) may be easily rewritten in the augmented Dickey-Fuller form as

$$\pi_{i,t} = c + \alpha_i^*\pi_{i,t-1} + \psi_{1,i}\Delta\pi_{i,t-1} + \ldots + \psi_{1,p-1}\Delta\pi_{i,t-p+1} + \nu_t, \tag{2.2}$$

where Δ denotes the difference operator. One can show that α_i^* equals the sum of the p AR coefficients, which is the measure of persistence proposed in Andrews and Chen (1994).

2.1.2 Data

We use monthly non-seasonally adjusted CPI data for France, Germany, Italy, the Netherlands and Spain for the period 1970-2006. The data is seasonally adjusted with the help of an auxiliary regression on a set of seasonal dummies[3]. To model the rate of inflation we use the first difference of the logged CPI series. All data is taken from OECD statistics. In the case of Germany the data relates to West Germany before 1991 and to unified Germany afterward.

2.1.3 Results

Table 1 gives the results for the estimated sum of AR coefficients for the five countries in different periods[4]. One can basically see that when considering the entire sample from 1970-2006 we find values for the sum of AR coefficients which are all larger than 0.7. In none of the countries we are able to reject the null of a unit root in the monthly inflation series. When splitting the sample right in the middle, we obtain slightly lower values for the inflation persistence in the period 1970-1988 than for the entire sample. But still we are not able to reject the unit root hypothesis in any of the cases. The results change for the period after 1988. Except for France and the

[3]We decided to apply this procedure instead of a filtering method, as the seasonal pattern of the series is quite stable over the sample period. We compared the results with those from a Census X12 method. The results reported in the following sections were not affected by the choice for the one or the other method.

[4]The lag length was chosen such that the residuals were uncorrelated.

Sample	France	Germany	Italy	Netherlands	Spain
1970-2006	0.84	0.87	0.95	0.93	0.92
	(0.08)	(0.06)	(0.03)	(0.03)	(0.06)
1970-1988	0.84	0.74	0.88	0.86	0.92
	(0.09)	(0.11)	(0.05)	(0.06)	(0.08)
1989-2006	0.67	**0.41**	**0.02**	0.83	0.53
	(0.18)	(0.06)	(0.07)	(0.08)	(0.22)

Table 1: *Inflation persistence as the sum of AR coefficients in different periods. Numbers in bold are able to reject the unit root hypothesis at the 5% level. Numbers in brackets indicate the standard errors.*

Netherlands we find sums of AR coefficients which are considerably lower than during the first half of the sample. For Germany and Italy we are even able to reject the null of the unit root test at the 5% level.

Of course, the choice for the split of the sample is purely arbitrary and not economically founded. Furthermore, the model considered here is kept very simple and neglects the structural elements arising from economic theory. Nonetheless, these results may be interpreted as a hint for a change in inflation persistence during the sample. A change in inflation persistence does in turn indicate a change in inflation dynamics in these countries.

2.2 Evidence for Changes in Inflation Dynamics from Non-Linear Univariate Models

2.2.1 Methodology

Empirical evidence suggests that many macroeconomic variables behave differently during upswings and downturns, i.e. the underlying data generating process (DGP) is subject to non-linearities (Hamilton 1989). The previous results suggest that this may also be the case for euro area's inflation rates. For this reason we consider a non-linear version of (2.2) given by

$$\pi_{i,t} = c(S_t) + \alpha_i^*(S_t)\pi_{i,t-1} + \psi_{1,i}(S_t)\Delta\pi_{i,t-1} + \ldots + \psi_{1,p-1}(S_t)\Delta\pi_{i,t-p+1} + \sigma(S_t)\nu_t, \tag{2.3}$$

where S_t denotes the state of the parameters at time t. σ^2 is the variance of the error term, which is generally assumed to be state-dependent[5]. Since the parameters depend on the regime S, which is assumed to be unobservable and stochastic, it is requires for the data generating process to be complete that we formulate a regime generating process. For notational reasons we define θ as the vector of all parameters in (2.3).

[5]This should be a reasonable assumption in the case of inflation rates, which are among others determined by monetary policy behavior. Sims and Zha (2006) have emphasized the importance of heteroscedastic error terms in describing U.S. monetary policy.

In Markov-switching models the states are assumed to follow a Markov chain with transition probability p_{uv}, where p_{uv} is defined as the probability for state v at time t given state u in $t-1$. The number of states is finite, i.e. $S_t = 1, \ldots, M$. We may write this as

$$p_{uv} = Pr(S_t = v | S_{t-1} = u), \quad \sum_{v=1}^{M} p_{uv} = 1 \quad \forall u, v \in \{1, \ldots, M\}.$$

The estimation is carried out with the Expectation-Maximization (EM) algorithm introduced by Dempster, Laird and Rubin (1977) and described in Krolzig (1996). The EM algorithm is an iterative maximum-likelihood estimation technique. It allows to estimate models with the observed time series depending on some unobservable stochastic variables, as given by the states S here in our case. In each iteration of the EM algorithm two steps are carried out:

1. The expectation step
2. The maximization step.

For notational purposes let us define an indicator variable for the states

$$I(S_t = u) = \begin{cases} 1 & \text{if } S_t = u \\ 0 & \text{otherwise,} \end{cases}$$

for $u = 1, \ldots, M$. We then may summarize the vector of all indicator variables at a given time t as

$$\xi_t = \begin{bmatrix} I(S_t = 1) \\ \vdots \\ I(S_t = M) \end{bmatrix}.$$

The expectation step in iteration j estimates the unobserved states ξ_t by their smoothed probabilities conditional on the data and the estimated vector of parameters of the previous iteration, which we denote by λ^{j-1}. Thereby, λ includes the parameters in (2.3) that need to be estimated as well as the parameters determining the Markov process, i.e. the initial state ξ_0 and the transition probabilities p_{uv}. Formally, the expectation maximization gives

$$Pr(\xi | \pi_i, \lambda^{j-1}).$$

In the maximization step the parameter vector λ is estimated with the help of the first-order condition of the likelihood function given by

$$p(\pi_i, \lambda) = \int p(\pi_i, \xi | \lambda) d\xi, \tag{2.4}$$

This may be rewritten as

$$p(\pi_{i,t}, \lambda) = \int p(\pi_i | \xi, \theta) Pr(\xi | p_{uv}, \xi_0) d\xi. \tag{2.5}$$

Coefficient	France	Germany	Italy	Netherlands	Spain
c_1	0.11	0.11	-0.10	-0.05	0.14
	(0.01)	(0.02)	(0.04)	(0.01)	(0.02)
c_2	0.35	0.06	0.03	0.23	—
	(0.04)	(0.13)	(0.03)	(0.09)	(—)
α_1^*	0.24	0.61	0.57	0.84	0.24
	(0.06)	(0.05)	(0.19)	(0.04)	(0.11)
α_2^*	—	0.98	0.90	0.67	0.73
	(—)	(0.12)	(0.07)	(0.12)	(0.09)
p_{11}	0.99	0.99	0.99	0.99	0.97
p_{22}	0.97	0.98	0.99	0.97	0.95
σ_1	0.16	0.18	0.19	0.15	0.15
σ_2	0.28	0.51	0.22	0.44	0.28
ln$L$	91.71	-98.14	71.60	28.05	68.16
# regimes	2	2	2	2	2
# lags	1	9	11	11	2
Model	MSIH	MSIAH	MSIAH	MSIAH	MSAH

Table 2: *Inflation persistence as the sum of AR coefficients - evidence from non-linear models. Numbers in brackets indicate the standard errors.*

The conditional regime probabilities $Pr(\xi_t|\pi_i, \lambda)$ are replaced by smoothed regime probabilities obtained from the last iteration of the expectation step.

We start with two regimes of the parameters and the variance for each country i. We then test the non-linear model with two regimes against the linear alternative using a Davies test[6]. If the Davies test cannot reject the null, we will also consider more than two states in the regression. The number of differenced lags is chosen such that the residuals are uncorrelated. We start with the most general model (MSIAH), which allows for a state-dependent intercept, parameters and variance. Using LR tests[7] we investigate, if further restrictions are appropriate (i.e. MSIH, MSAH, MSH). The estimation is carried out with Hans-Martin Krolzig's MSVAR package.

2.2.2 Results

The numerical results of the estimation are reported in table 2. For all countries we estimated a model with two regimes, as proposed by the Davies test. We also tried specifications with a higher number of regimes, which did not deliver any reasonable results[8]. The temporal distributions

[6]The Davies test is a modified likelihood ratio (LR) test and has the null hypothesis $m-1$ regimes vs. $\geq m$ regimes. In the case when $m = 1$ we basically test for a linear vs. a non-linear specification. For more details see Davies (1987).

[7]The LR test can be based on the LR statistic $LR = 2(lnL(\lambda) - lnL(\lambda_0))$, where λ_0 denotes the restricted ML estimate of the parameter vector λ. Under the null, LR has an asymptotic χ^2 distribution with r degrees of freedom, where r is the number of restrictions.

[8]The usual information criteria did also favor a two-state specifications instead of more complex ones.

of the regime probabilities are given in section 4.1 of the appendix. We can basically observe a dominance of regime 2 during the first half of the sample in all five countries, while regime 1 is more appropriate in describing the data during the second half of the sample. This regime change is first observed in France, followed by the Netherlands, Spain and finally Germany and Italy. While we find evidence for lower inflation persistence in Germany, Italy and Spain in regime 1, as indicated by α_1^*, the opposite is true for the Netherlands. Using a LR test, France provides no evidence at all for a change in the autoregressive coefficients so that we estimated a MSIH model. The LR test does also suggest a regime-invariant intercept for Spain. For all countries we find a decrease in variance from regime 2 to regime 1.

Despite the differences in the model specifications and results we observe a simultaneity in the regime switches across countries, as the switch toward regime 1 occurs in all five countries around 1985. Given these results and the fact that all five time series provide evidence for non-stationary behavior, as shown in table 1, we analyze in the following, if we can find one or more cointegrating relationships between the variables that again turn out to be stationary. Given that there exists at least one cointegrating relationship, we estimate a MS-VECM, which allows for parameter changes in the adjustment coefficients and the short-run dynamics.

2.3 Evidence for Changes in Inflation Transmission from Non-Linear Multivariate Models

2.3.1 Economic Fundamentals

For an analysis of the dynamics and transmission of inflation in the euro area, it requires a theoretic framework on which the econometric model can be based. A simple but effective model framework is offered by the relative purchasing power parity (PPP). We define the real exchange rate ρ as

$$\rho = E\frac{P}{P^*}, \tag{2.6}$$

where E denotes the nominal exchange rate, P is the price of domestic goods in domestic currency and P^* the price of foreign goods in foreign currency. When we now consider rates of change in the real exchange, we end up with

$$\frac{\Delta\rho}{\rho} = \frac{\Delta E}{E} + \frac{\Delta P}{p} - \frac{\Delta P^*}{P^*}, \tag{2.7}$$

which is the sum of the change in the nominal exchange rate, i.e. nominal appreciation, plus the change in the domestic price level, i.e. domestic inflation, less foreign inflation. While absolute PPP demands the real exchange rate to be equal to 1 in the long run, the relative PPP requires it to be constant in the long run, i.e. $\rho = const$. From this it follows that $\frac{\Delta\rho}{\rho} = 0$, which finally gives

$$\frac{\Delta E}{E} = \frac{\Delta P^*}{P^*} - \frac{\Delta P}{p}. \tag{2.8}$$

Exactly this relationship will be helpful for the empirical analysis in the following, as it gives a theoretical relationship for the long-run behavior of the inflation rates of the countries under consideration. This means that (2.8) offers us a way to test for a long-run relationship between the rates of inflation using cointegration methods given that the inflation rates may be considered as $I(1)$ variables, which has already been analyzed in section 2.1.

2.3.2 Methodology

We formulate a simple VAR in the five inflation rates with lag length $q+1$, in which we should expect at least some positive relationship between the inflation rates, as implied by the relative purchasing power parity. The VAR takes the following form

$$y_t = c^* + \sum_{s=1}^{q+1} A_s^* y_{t-s} + u_t, \tag{2.9}$$

where y_t is a vector 5×1 vector in the rates of inflation, A_i^* is the 5×5 matrix of coefficients and u_t is a 5-dimensional vector of Gaussian errors. c^* denotes a 5-dimensional vector of constants.

Given that the five endogenous variables are cointegrated, we may use Granger's representation theorem to formulate the VECM

$$\Delta y_t = \alpha(\beta' y_{t-1} + c) + \sum_{s=0}^{q} A_s \Delta y_{t-s} + u_t. \tag{2.10}$$

The non-linear version of (2.10) would be given by

$$\Delta y_t = \alpha(S_t)(\beta' y_{t-1} + c) + \sum_{s=0}^{q} A_s(S_t) \Delta y_{t-s} + u_t, \tag{2.11}$$

where S_t indicates the state of the coefficients at time t. Further, we consider the covariance matrix of the error terms Σ to be state-dependent. In particular, we assume $u_t|S_t \sim N(0, \Sigma(S_t))$. The unobservable regime variable S_t is determined by a Markov chain with M states.

For the estimation of (2.11) we use the two-step procedure proposed in Krolzig (1996). The first step uses the Johansen maximum likelihood procedure to determine the cointegration rank r of the system (2.10) and the corresponding cointegrating relationships (Johansen 1995). In a second step we estimate the remaining coefficients in (2.11) given the cointegrating relationships obtained in the first step using the EM algorithm.

2.3.3 Cointegration Analysis

Given that the five inflation series were all found to be integrated of order one, we want to analyze in the following, if we are able to identify common stochastic trends among them, using a cointegration analysis. Table 3 reports the results of the trace and maximum eigenvalue test[9].

[9]We did not include a constant or trend in the test specification. We first allowed for a constant inside the cointegrating relationship, which turned out to be not significant in any cointegrating equation. We included five differenced lags, as suggested by the Akaike criterion.

H_0	Trace Stat.	5% Critical Val.	Max. Eigenvalue Stat.	5% Critical Val.
$r = 0$	207.54	59.46	78.87	30.04
$r = 1$	128.67	39.89	57.20	23.80
$r = 2$	71.47	24.31	42.96	17.89
$r = 3$	28.51	12.53	25.61	11.44
$r = 4$	2.90	3.84	2.90	3.84

Table 3: *Results Trace Test and Maximum Eigenvalue Test.*

Both tests indicate four cointegrating relationships[10]. In particular, we find[11]

$$z_{1,t} = \pi_{FR,t} - \underset{(0.06)}{0.28}\,\pi_{GER,t} \tag{2.12}$$

$$z_{2,t} = \pi_{IT,t} - \underset{(0.06)}{0.79}\,\pi_{GER,t} \tag{2.13}$$

$$z_{3,t} = \pi_{NL,t} - \underset{(0.09)}{1.11}\,\pi_{GER,t} \tag{2.14}$$

$$z_{4,t} = \pi_{ES,t} - \underset{(0.08)}{0.40}\,\pi_{GER,t}. \tag{2.15}$$

We decided to order Germany last, as it is the largest economy within the group of countries under consideration. This allows an easier interpretation and comparison of the estimated coefficients.

All coefficients head in the expected direction. Interestingly, we find coefficients for Germany, which are all smaller than one in absolute values except for the one in equation (2.14). This result contradicts the implications of the relative purchasing power parity (PPP), which would require a cointegrating vector of $(1, -1)$. A possible explanation for this strong deviations from the theoretically implied relationship could be the fact that the PPP is built on true prices and true inflation, while we use CPI inflation, which is of course just a proxy for true inflation. The underlying baskets of the five CPIs include different products with different weights, which may cause deviations from the theoretical relationships. Further, the cointegrating relations come from a linear model that may be regarded as a first-best approximation to the non-linear one, which we do think is more appropriate to describe the data. As we think that the relative PPP is a plausible mechanism for these five economies, we continue not only with the results given by equations (2.12)-(2.15), but also with the theory-consistent cointegrating vectors $(1, -1)$. In the following we will report the results of both the restricted and the unrestricted version.

[10] These results were also supported by pairwise country-specific analyses.

[11] Numbers in brackets indicate the standard errors.

2.3.4 Model Specification

We estimated a MS-VECM with two regimes[12]. We allowed for regime-dependent error variance, which we think is reasonable in the case of inflation rates and is also supported by formal LR tests. We included one differenced lag, which turned out to be sufficient for the residuals being uncorrelated.

2.3.5 Results

Figure 1 shows the temporal distribution of the two regimes for the restricted and the unrestricted version of the model. We can basically see that the sample was mostly dominated by regime 2 except for the period 1973-1983.

In particular, we obtain the following numerical results for the matrix of adjustment coefficients in the two regimes[13]:

$$\alpha_{unrestr}(S_t = 1) = \begin{bmatrix} \mathbf{-0.70} & 0.13 & 0.05 & 0.16 \\ -0.05 & 0.31 & 0.34 & -0.14 \\ \mathbf{0.41} & -0.19 & \mathbf{0.13} & -0.12 \\ \mathbf{0.55} & \mathbf{0.64} & \mathbf{-0.50} & -0.14 \\ 0.08 & \mathbf{0.32} & 0.12 & \mathbf{-0.94} \end{bmatrix}$$

$$\alpha_{unrestr}(S_t = 2) = \begin{bmatrix} \mathbf{-1.01} & -0.06 & \mathbf{0.25} & \mathbf{0.25} \\ \mathbf{-0.54} & \mathbf{0.42} & \mathbf{0.58} & \mathbf{0.51} \\ \mathbf{-0.35} & \mathbf{-0.73} & \mathbf{0.51} & \mathbf{0.30} \\ \mathbf{-0.29} & \mathbf{0.26} & \mathbf{-0.19} & \mathbf{0.17} \\ \mathbf{0.37} & -0.08 & \mathbf{0.27} & \mathbf{-0.68} \end{bmatrix}$$

$$\alpha_{restr}(S_t = 1) = \begin{bmatrix} \mathbf{-0.32} & -0.10 & -0.04 & \mathbf{0.47} \\ 0.20 & 0.15 & 0.24 & 0.09 \\ \mathbf{0.48} & \mathbf{-0.42} & 0.06 & -0.01 \\ 0.28 & \mathbf{0.67} & \mathbf{-0.50} & -0.30 \\ \mathbf{0.39} & 0.30 & 0.33 & \mathbf{-0.84} \end{bmatrix}$$

$$\alpha_{restr}(S_t = 2) = \begin{bmatrix} \mathbf{-0.64} & 0.18 & 0.15 & \mathbf{0.41} \\ \mathbf{-0.54} & \mathbf{0.44} & \mathbf{0.61} & \mathbf{0.52} \\ \mathbf{-0.16} & \mathbf{-0.62} & \mathbf{0.44} & \mathbf{0.37} \\ \mathbf{-0.12} & \mathbf{0.28} & \mathbf{-0.36} & \mathbf{0.17} \\ \mathbf{0.58} & 0.02 & 0.11 & \mathbf{-0.63} \end{bmatrix}.$$

Regardless of the specification chosen we find the adjustment coefficients for Germany being insignificant at the 5-percent level in regime 1. This implies a weak exogeneity for Germany within regime 1, which reveals the dominance of Germany's monetary policy. The cointegrating

[12]The Davies test could clearly reject the null hypothesis of a linear system. We also estimated the model with three regimes, which did not deliver any reasonable results. Further, one should note that the degrees of freedom substantially decrease with the number of regimes in a MS-VECM.

[13]Countries are ordered alphabetically, i.e. France, Germany, Italy, Netherlands, Spain. Numbers in bold face indicate significance at the 5-percent level.

equation between Germany and France has a negative and significant adjustment coefficient in the equation for France, which is a reasonable result. The same is true for the other countries respectively. The only exception is Italy in the unrestricted version of the model, where we obtain an insignificant adjustment coefficient for the cointegrating relation between Germany and Italy, while both the cointegrating relation between France and Germany and the one for the Netherlands and Germany are significant, but with a positive sign. When restricting those coefficients to zero that are not statistically significant at the 5-percent level, this would imply an unstable system. In the version with the restricted cointegrating relationships we find basically the same results for the adjustment coefficients in regime 1, but with Italy exhibiting a significant and negative adjustment coefficient for the cointegrating equation between Italy and Germany, which seems to be a more reasonable result than the one in the unrestricted version.

The most substantial result we may find in regime 2 is the fact that Germany is no longer weakly exogenous, regardless of how the cointegrating relationships are specified. All four adjustment coefficients become significant. Interestingly, we see that the overall number of significant adjustment coefficients has substantially increased, which would be consistent with a system of more strongly integrated economies. Again, we find that a cointegrating relationship between Germany's rate of inflation and a specific country does turn out to be significant for the specific country, which should be a reasonable result.

Figure 1 reports the temporal distributions of the two regimes for both the unrestricted and restricted version of the model. Again, we see that the results are hardly affected by the specification chosen. For the period 1970-1973 we find inflation transmission being described by regime 2 with Germany being endogenous. With the breakdown of the Bretton Woods system in 1973 we observe a shift toward regime 1 with Germany becoming weakly exogenous. Regime 1 played the dominant role in the years 1973-1983. Except for some brief changes in 1986, 1987 and 1989 regime 1 played basically no role after 1983. Finally, from 1989 onwards we see that the inflation transmission mechanism is solely described by regime 2.

Hence, regardless of how the cointegrating relations are specified, we find a similar distribution of the regimes over time. Further, the regime shifts seems to be consistent with the evidence obtained from the univariate models in section 2.2. With respect to the adjustment coefficients we also obtain consistent results over the two specifications except for Italy, where we find more plausible estimates in the case of the restricted version of the model. For this reason, we will continue the analysis on the basis of the restricted model.

The results indicate that the dominance of German monetary policy already ended around 1984 with Germany being no longer weakly exogenous.

2.3.6 Impulse Responses

Impulse responses offer a fairly simple way to analyze the inflation dynamics and inflation transmission for the five countries under consideration. The impulse response analysis describes the response of $\pi_{i,t+\tau}$ to a one-time impulse in $\pi_{j,t}$ with all other variables held constant. One problematic assumption in the impulse response analysis is that a shock in t occurs only in one variable. This demands the error terms summarized in the vector u_t in (2.11) to be independently distributed. This assumption will be hardly fulfilled, when considering the inflation rates of the five largest economies in the euro area. Therefore, we need a way to orthogonalize the

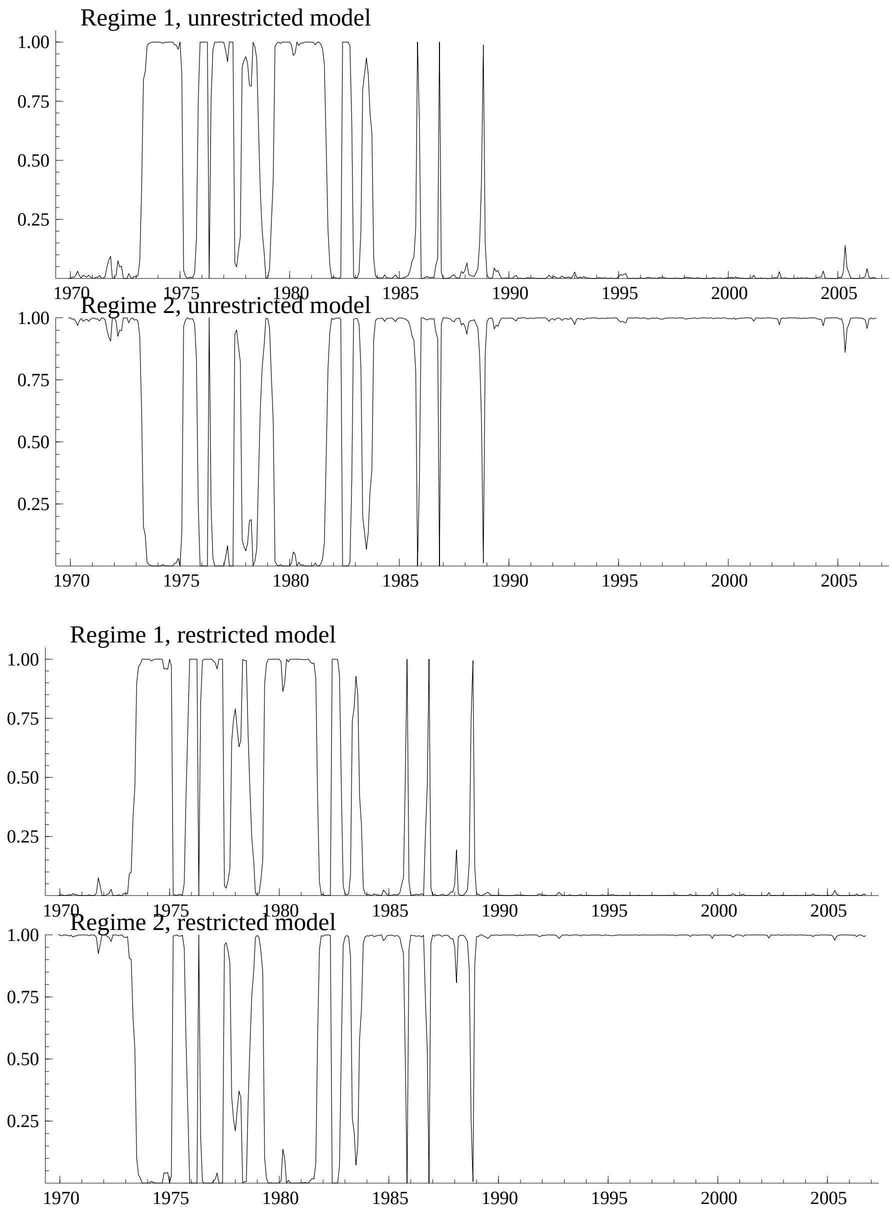

Figure 1: *Temporal Distribution of Regime Probabilities in the Unrestricted (Upper Panel) and Restricted Model (Lower Panel), 1970-2006.*

impulse responses. We do this using a Cholesky decomposition of the covariance matrix. The Cholesky decomposition requires an appropriate causal choice of the variables' ordering, as the impulse responses may be substantially affected by it. We decided here to order the inflation rates by the economic size of the countries, i.e. we put Germany first followed by France, Italy, Netherlands and Spain. A covariance that turned out to be insignificant at the 5-percent level was restricted to zero[14].

Figure 7-11, given in section 4.2 of the appendix, show the country-specific impulse responses to a one-unit shock. For the impulse response analysis we restricted those coefficients in the short-run dynamics to zero that turned out to be insignificant at a 5-percent level. For Germany it then follows that its rate of inflation during regime 1 is described by a random walk, as both long-run and short-run dynamics are not significantly different from zero in the VECM representation. Therefore, it should not be surprising in the following, when we do not find any response of Germany in regime 1 to inflation shocks taking place in one of the other countries.

Generally, one can see that the initial shock has become less persistent in regime 2 than it used to be in regime 1 in case of the two largest countries France and Germany, while the shocks in Italy, the Netherlands and Spain exhibit a persistence of almost equal size across the two regimes. With the initial shocks in France and Germany in regime 1 being more persistent than in regime 2, we find that the responses of the other countries show a higher degree of persistence. Further, one can see that the three smaller countries have gained in importance for the transmission mechanism in regime 2, particularly Italy and the Netherlands.

In regime 1 in figure 8 we observe an initial negative impact of the German inflation shock on France, Italy and the Netherlands. The same is true for regime 2 in figure 7 for the reaction of Germany to the French inflation shock. This result is rather unexpected in that an inflationary shock can actually elicit a deflationary reaction. Other studies on inflation transmission have found similar results[15]. Eun and Jeong (1999) explain this result by an overshooting depreciation of the country's currency, in which the initial shock takes place. This then leads to lower import prices in the other countries despite the inflation shock. A theoretical foundation for this may be found in Dornbusch (1976) known as Dornbusch's overshooting result.

Due to the lower persistence of the inflation shock itself in France and Germany, we may say that inflation persistence has reached a lower level since late 1980s, when regime 2 occurred for the last time in the sample period. This result is particularly remarkable, as it offers an explanation for the decrease in inflation persistence in the euro area, as described in Altissimo, Ehrmann and Smets (2006). According to the impulse responses presented here, the fall in inflation persistence within the euro area during the last decade is substantially related to a fall in the persistence of inflation shocks in Germany and France.

Further, one can easily see that the impact of an inflation shock in Germany and France on the other countries is almost negligible in regime 2, particularly in the case of Germany, while the three smaller countries, i.e. Italy, Netherlands and Spain, gain in importance within regime 2.

[14]The impulse response are computed with a Matlab code that is available upon request.

[15]Cf. Yang, Guo and Wang (2006) and Eun and Jeong (1999).

3 Conclusions

This paper has empirically analyzed the inflation transmission mechanism in the euro area for the period 1970-2006 using Markov-regime switching models.

The analysis offers evidence for stable inflation dynamics and a stable transmission mechanism in the euro area since the mid-1980s. Interestingly, it turns out that the period in which Germany is weakly exogenous already ended 1983. This indicates that the period of Germany's monetary policy being dominant for today's euro area already ended earlier than is usually assumed. Using impulse responses, we find that inflation persistence across countries has substantially decreased since that time, which is in accordance with recent research on inflation persistence. Our analysis finds the reason for this decline in the decrease of inflation persistence in France and Germany, leading to less persistent responses of the other countries to a shock in inflation in one of these countries. For the period since the mid-1980s it is found that country-specific inflation shocks are less strongly transmitted across the euro area than in the period before. Instead we find that the country in which the shock takes place returns quite rapidly back to the long-run equilibrium, which contradicts conventional findings that inflation in the small country is caused by the large country[16].

[16] Cf. Cheung and Yuen (2002).

4 Appendix

4.1 Regime Probabilities - Univariate Model

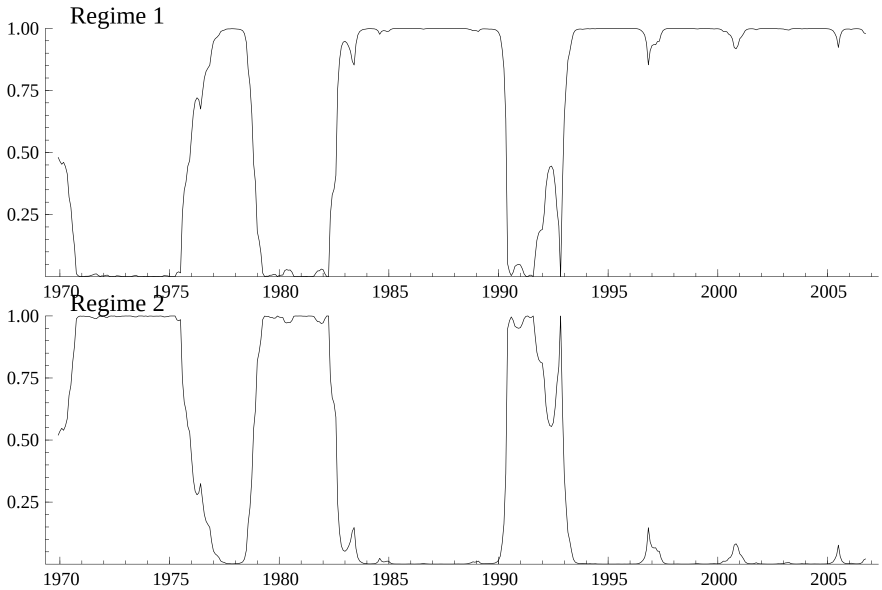

Figure 2: *France, Regime Probabilities, 1970-2004.*

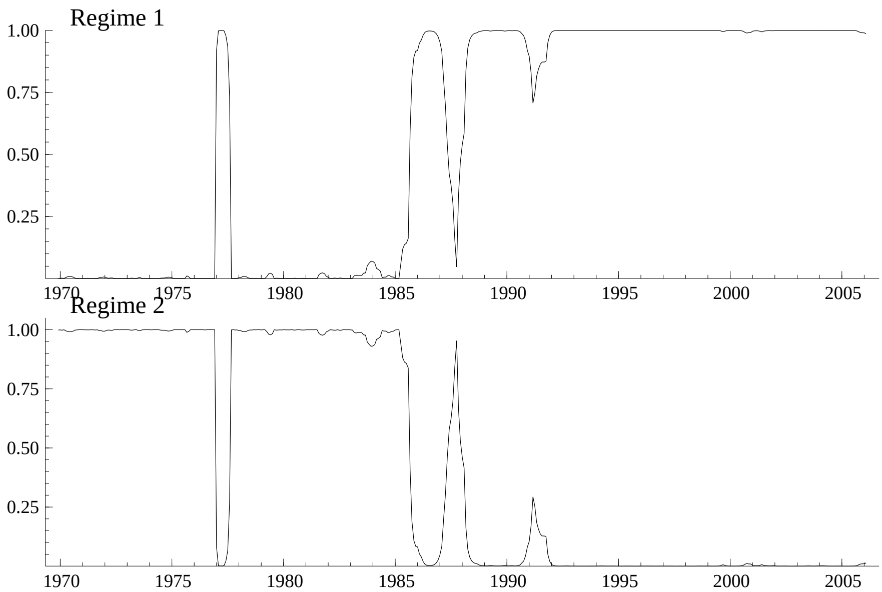

Figure 3: *Germany, Regime Probabilities, 1970-2004.*

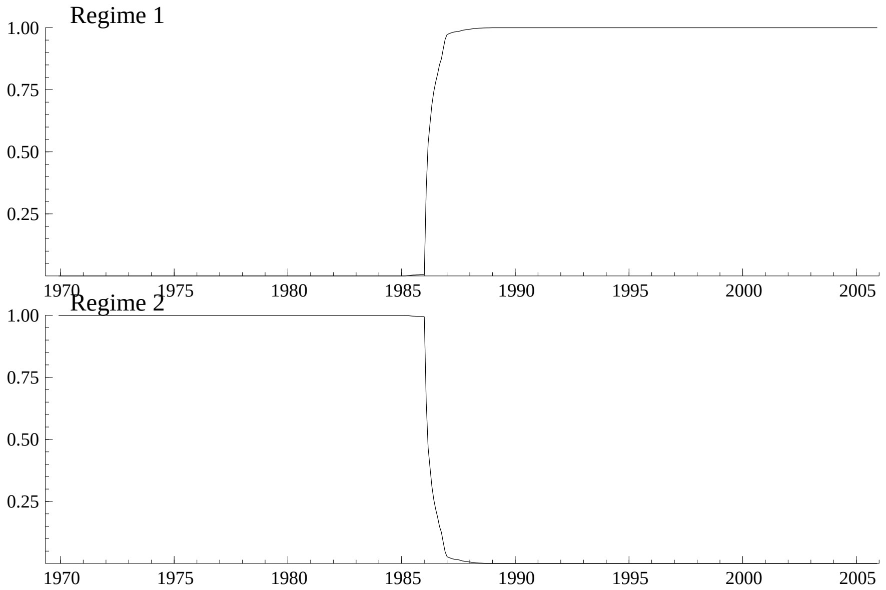

Figure 4: *Italy, Regime Probabilities, 1970-2004.*

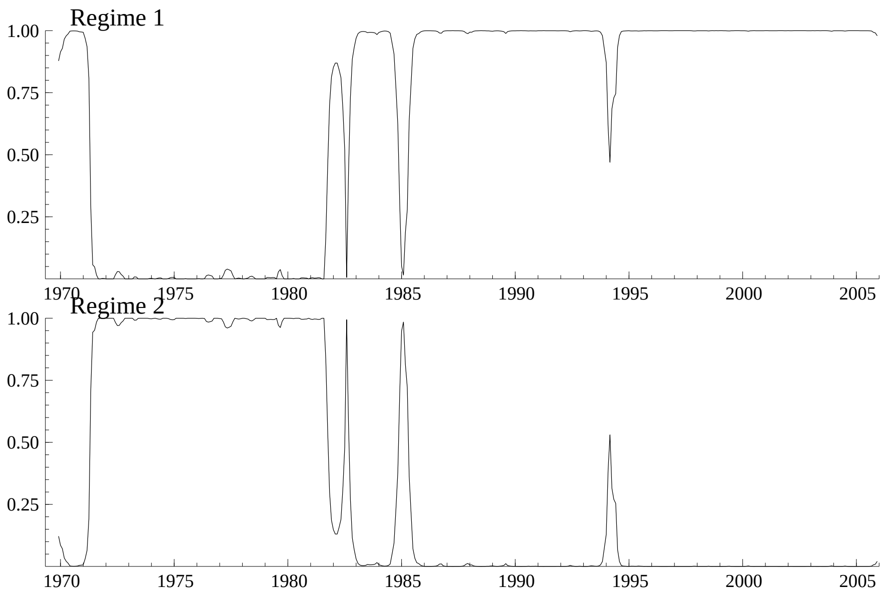

Figure 5: *Netherlands, Regime Probabilities, 1970-2004.*

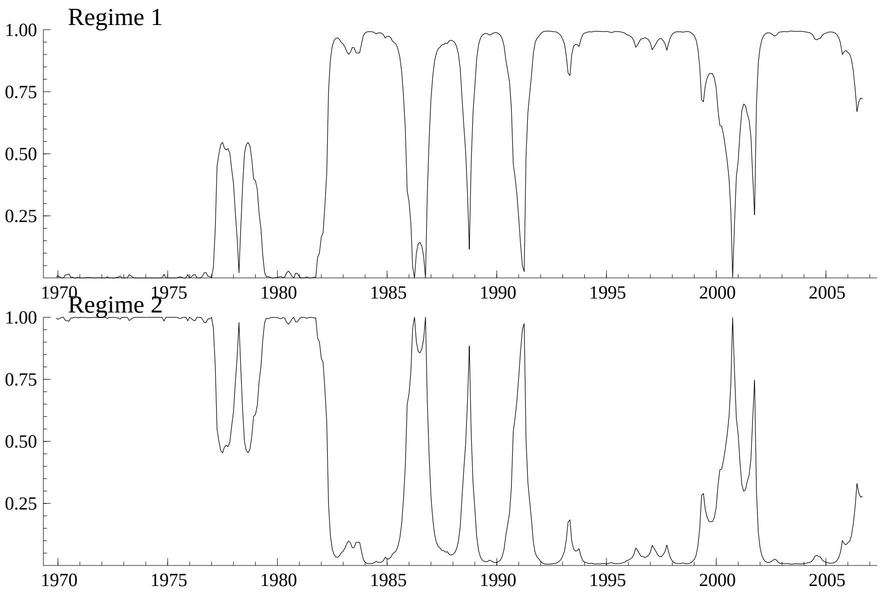

Figure 6: *Spain, Regime Probabilities, 1970-2004.*

4.2 Impulse Responses

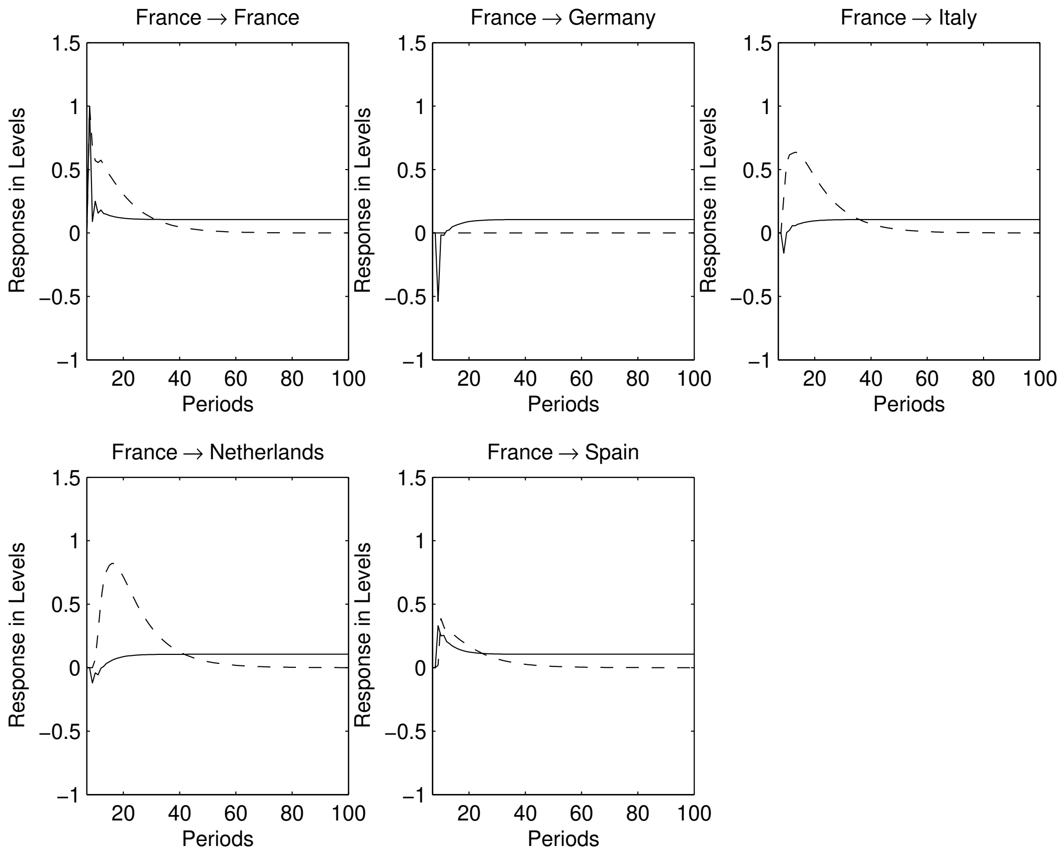

Figure 7: *Impulse Responses to a One Unit Shock in France (Dashed Line: Regime 1, Solid Line: Regime 2).*

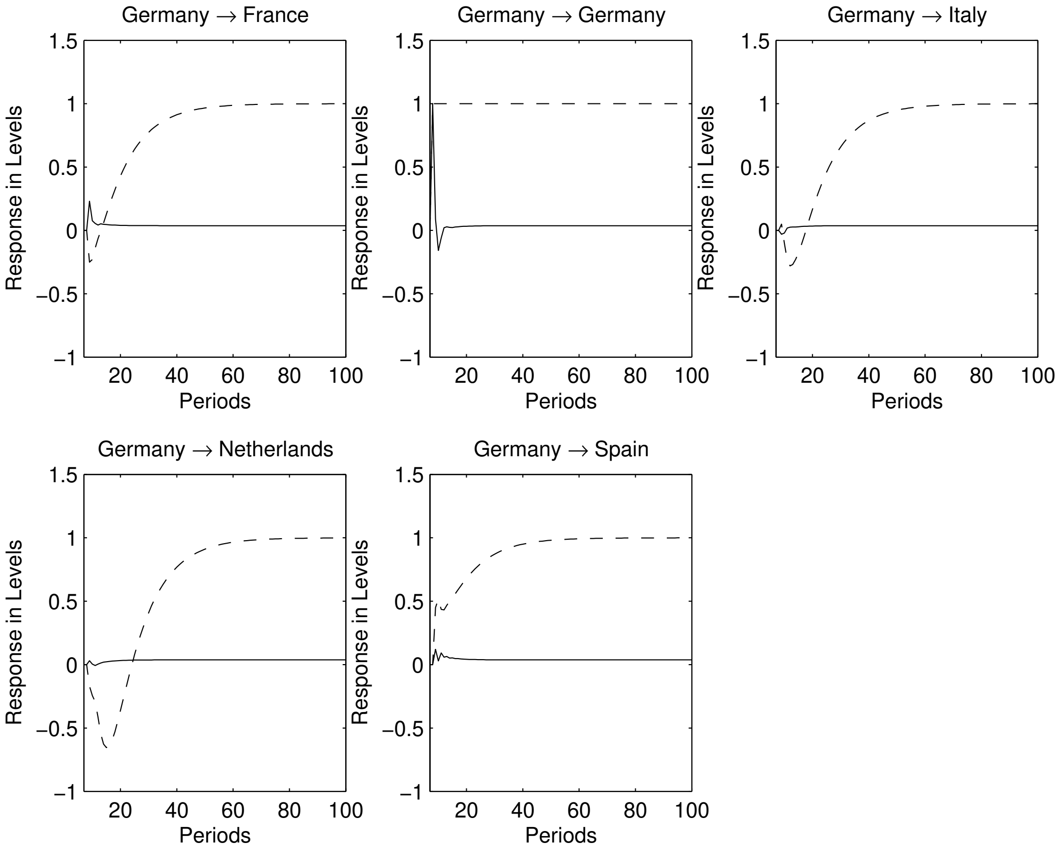

Figure 8: *Impulse Responses to a One Unit Shock in Germany (Dashed Line: Regime 1, Solid Line: Regime 2).*

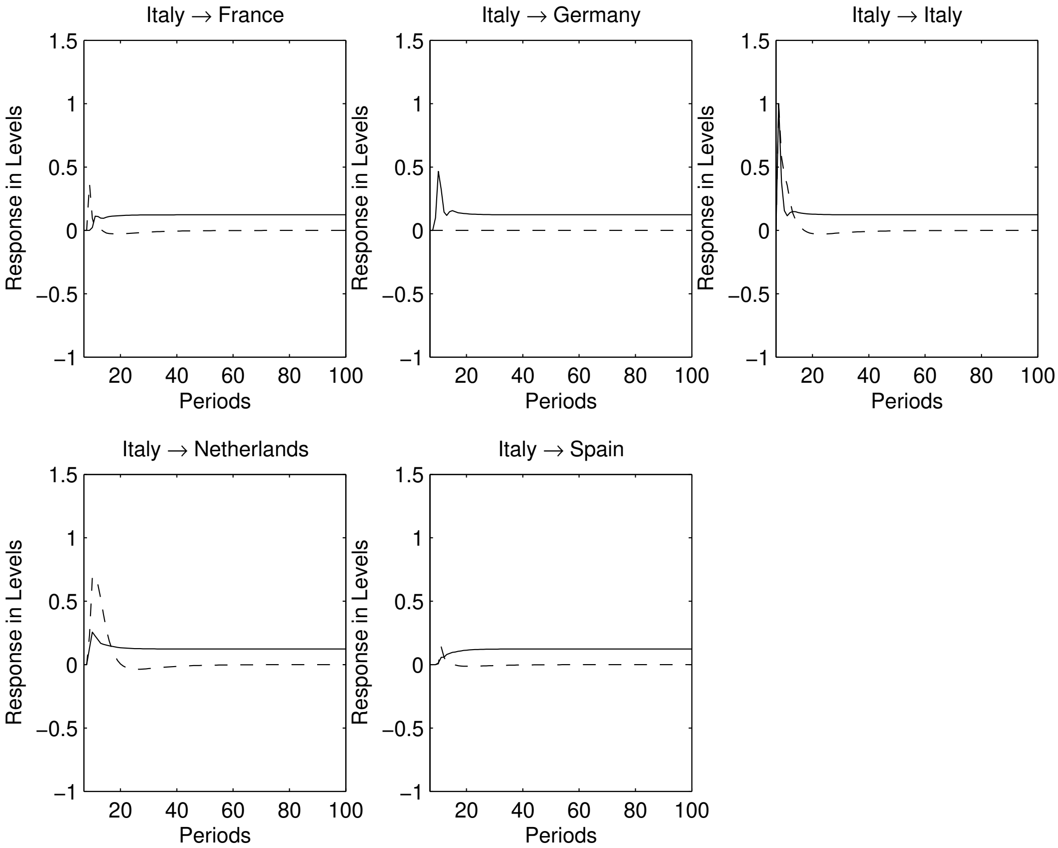

Figure 9: *Impulse Responses to a One Unit Shock in Italy (Dashed Line: Regime 1, Solid Line: Regime 2).*

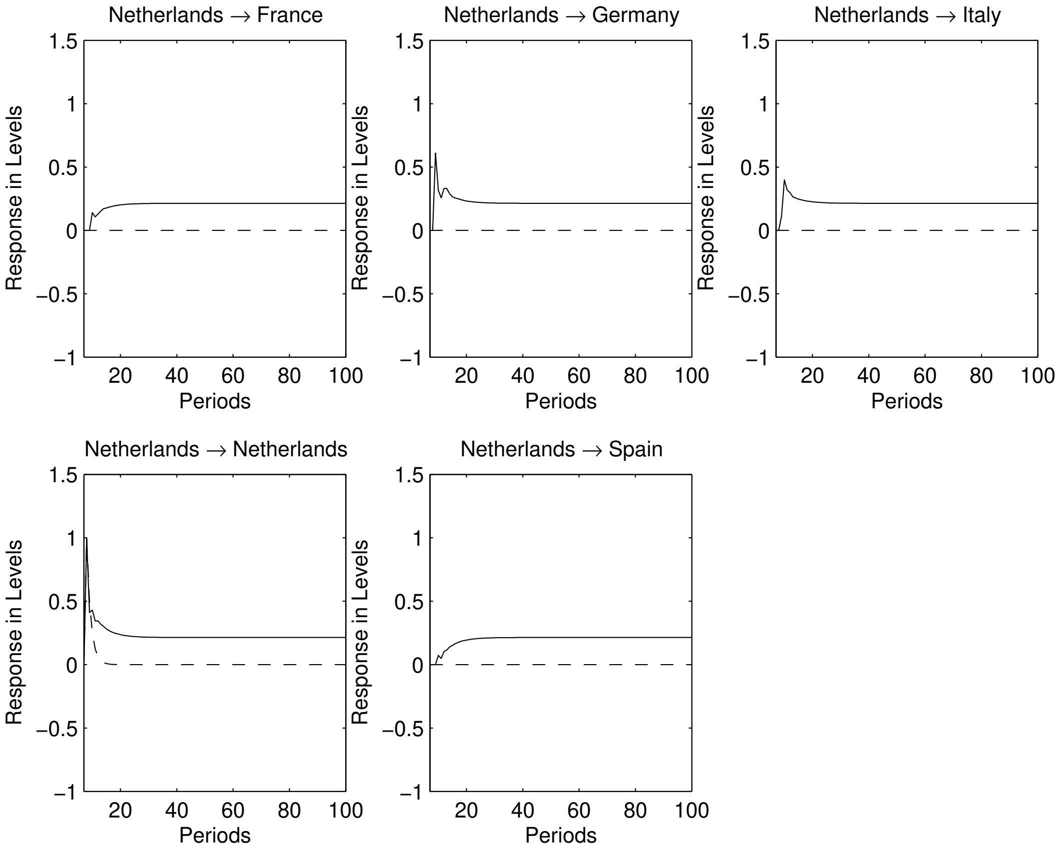

Figure 10: *Impulse Responses to a One Unit Shock in the Netherlands (Dashed Line: Regime 1, Solid Line: Regime 2).*

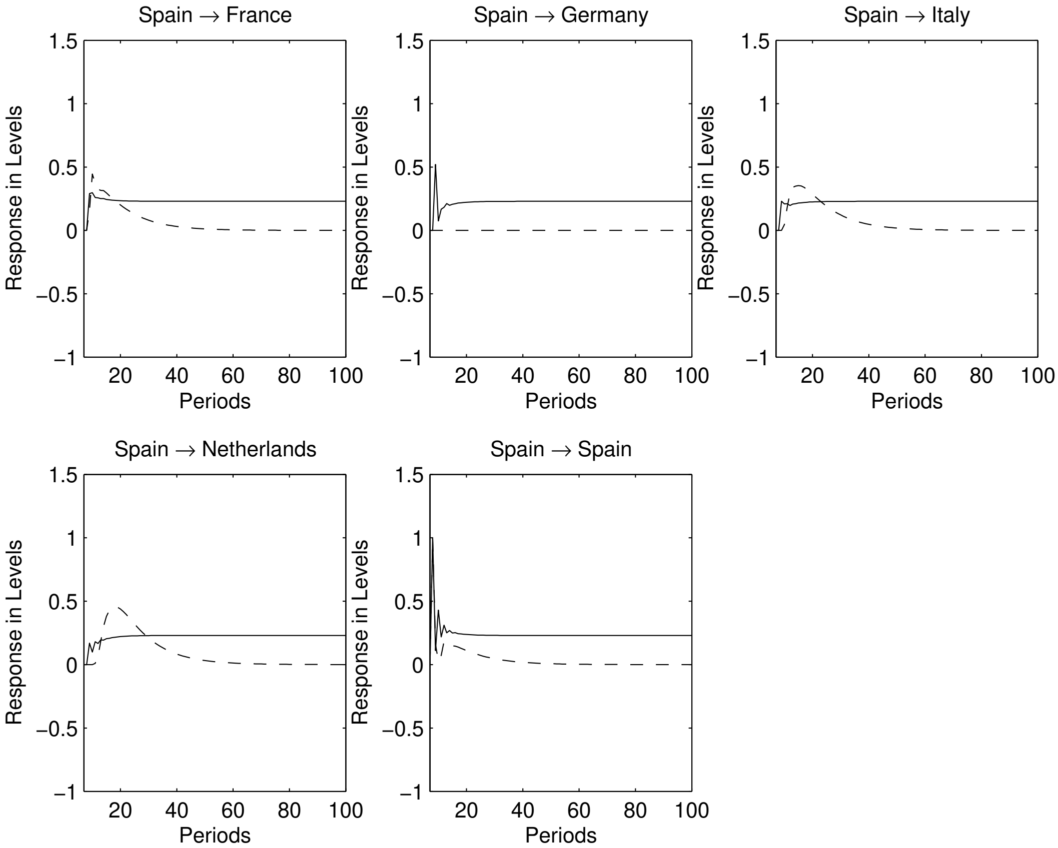

Figure 11: *Impulse Responses to a One Unit Shock in Spain (Dashed Line: Regime 1, Solid Line: Regime 2).*

4.3 Data

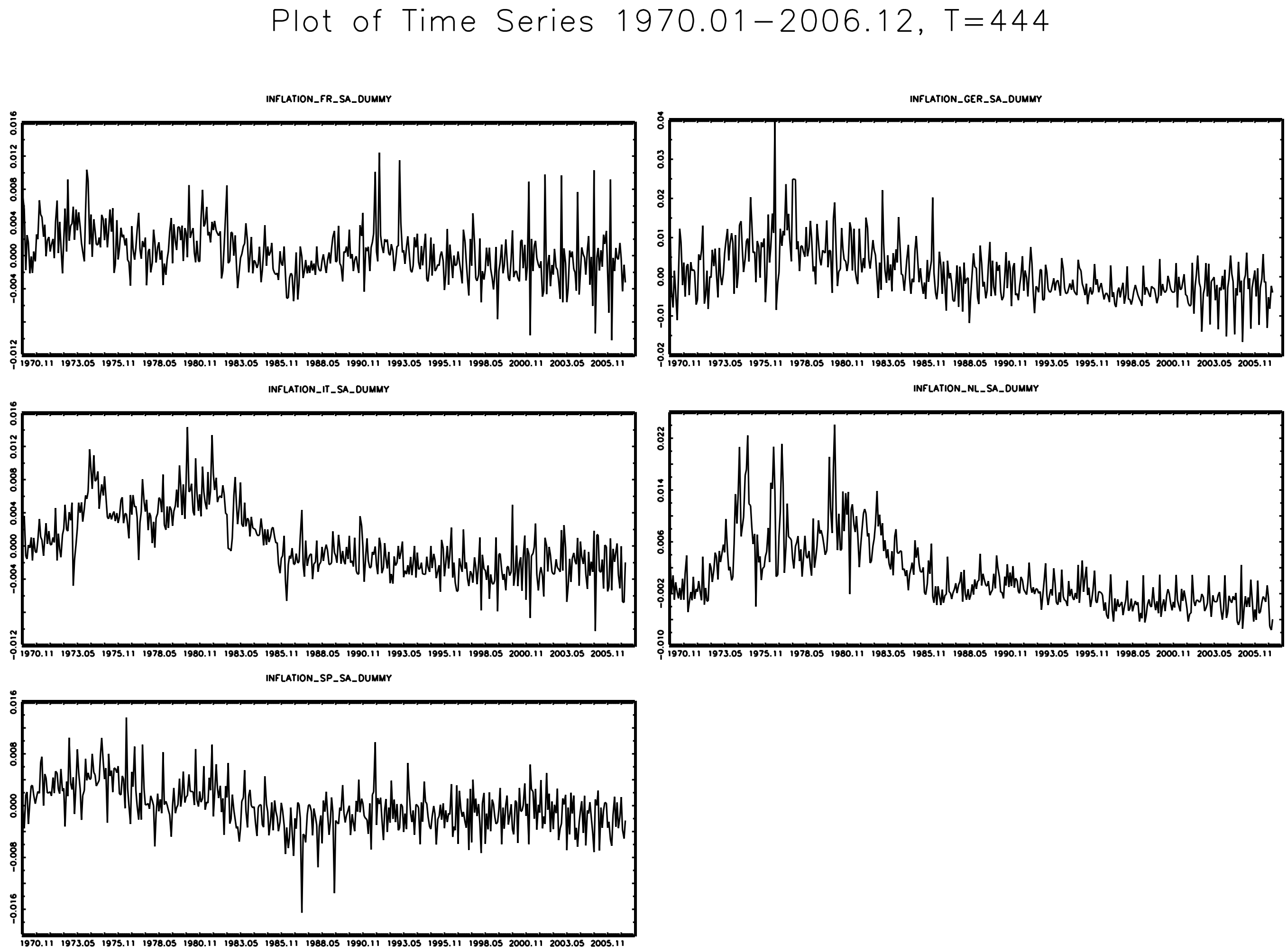

Figure 12: *Monthly Inflation Rates, Seasonally Adjusted.*

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