Monetary Transmission in the Euro Area: A Factor Model Approach

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Abstract

This paper studies the transmission of common monetary shocks across European countries by using a dynamic factor model (Forni-Reichlin (1998)). This technique allows to extract the common European monetary shock and to compute country-specific responses. Our identification employs rotations of the shocks space and a loss function (as in Uhlig (1999)). European countries display responses in line with a broad set of theoretical models and are characterized by quantitatively different responses. Spain and Germany are the most sensitive countries to common monetary shocks, while France, the Netherlands and especially Italy are the least. The interest rate channel is significant in explaining these asymmetries while we find no role for the credit channel.

JEL classification: C33, F42, E52, E58.

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1 Introduction[1]

When countries decide to participate to a Monetary Union, they fix the nominal exchange rate with respect to the other participants and lose the control of their national monetary policy. There are three different ways in which a common currency managed by a Central Bank with area-wide objectives may create conflicts across countries about the preferred conduct of monetary policy.

On one side, different countries in the monetary union may be hit by asymmetric shocks. The "Optimal Currency Area" literature (see Mundell (1961)) identifies conditions under which a monetary union can overcome the impact of these asymmetric shocks: the costs of forming a CA are lower the higher the flexibility of wages and prices, the higher the mobility of factors of production and the higher the automatic redistribution generated by a centralized fiscal system.

A second channel of possible disagreement concerns the differences in the cyclical conditions. If one country is experiencing a boom, with output above potential and inflationary pressures and another is experiencing a recession, with a positive output gap and rising unemployment, they would require different monetary policies that of course cannot be managed by a central bank with at its disposal only union-wide instruments. Many papers have studied whether European countries have become more synchronized (for a recent survey, see Mihov (2001)) and they conclude that in the last 10 years business cycle have converged[2].

In this paper we focus on a third issue that has received less consideration in the literature since recently and is concerned with the asymmetric effects of a common monetary policy.

Even if there are no asymmetric shocks and if business cycles are perfectly aligned, differences in the transmission mechanism (in the propagation mechanism of a *common* monetary policy) may create tensions in the management of monetary policy. Some countries may be characterized by faster or slower speeds of adjustment or by bigger or smaller dynamic responses to the same shock and this may itself generate asymmetries and distributional differences, again undermining the stability and the political consensus towards the union.

Country-specific institutional and economic characteristics can interplay with a common monetary policy to create and magnify asymmetries across the components of the union.

Important differences are those in the industry mix, through the interest rate channel. Countries with a large percentage of interest-sensitive industries (for instance, those producing durable or investment goods) will in turn be very sensitive to monetary actions.

Differences in the financial systems (Giovannetti-Marimon (1999)) and in the debt structure of firms will create asymmetries through the *credit channel*, as those countries in which banks are the main source of financing for firms and those in which banks are less healthy will be more sensitive (see Bernanke-Gertler (1998)).

The importance of the different channels in the EMU is discussed in Dornbush-Favero-Giavazzi (1998), Schmidt (1999) and Guiso et al. (1999)).

Other works measure empirically the extent to which the transmission differs across European countries and US states. A small subset of these works goes one step further and evaluates which are the most relevant channels to explain these differences. Different techniques have been employed to identify asymmetries in the transmission. The empirical literature can be divided in three main categories:

1.1 Country-specific studies

(a) Country-specific VARs (Gerlach-Smets (1995), Barran-Coudert-Mojon (1996), Ramaswamy-Sloeck (1997), Ehrmann (1998), Kieler-Saarenheimo (1998), Leichter-Walsh (1998), Kouparitsas (1999), Mihov (2001))

These studies estimate separate VARs for each country and conduct impulse response analysis: how an unexpected shock to one of the variables (typically, a shock to the interest rate, assumed to be the monetary policy instrument) affects the dynamics of the other variables in the system[3].

We have many doubts concerning country-specific VAR models and we discuss them in detail as we will propose a method able to overcome these limitations.

First, country-specific VARs employed in the context of the EMU are affected by an endogeneity problem.

The European Monetary System was designed to allow a certain degree of symmetry among countries, but in practice, it became an asymmetric system in which Germany was playing the role of the nominal anchor. The European-wide monetary policy was run by the Bundesbank with the other countries struggling to maintain the exchange rate parity with the Deutsche Mark. Given these premises, it is easy to see that country-specific VARs that do not control for the leading role of Germany can easily mix up endogenous responses of, say, French monetary policy to a German policy shock with the truly exogenous French monetary policy shock. The fact of not controlling for the German leading role in the EMS causes the system to be misspecified. As shown in Clarida-Gali-Gertler (1998) for instance, the policy rule for France and Italy contains as a significant explanatory variable the German policy rate. If this is the case, inference drawn from a French-specific VAR will be misleading, bringing to wrong conclusions about the French transmission mechanism. Even if the country-specific transmission would be correctly identified by controlling for all the exogenous effects, this is not what we are interested in. The final objective is to measure the effects of a *common* monetary policy on heterogeneous units.

Second, country-specific VARs typically do not include any foreign variables. The covariance structure and the interdependences across countries are not exploited. As the objective of these exercises is to measure the effects of a common monetary policy across interrelated economic units, it seems sensible to have methods suited to analyze all the countries together, allowing for a reduction in the number of parameters to be estimated, but at the same time allowing for heterogeneity in the propagation mechanism of shocks.

A third criticism comes from the observation that the size of the Impulse Response Functions (IRFs) and the comparison among them depend on the size of the shock. IRFs have been typically computed in response to a one-percent shock equal across countries, or to a country-specific one standard error shock. The shape of the responses will be the same, but the magnitudes will change. If the VAR is used to study the transmission in a single country, this does not create any problem: it is normal that changing the size of the impulse, will also modify the size of the responses. What happens when country-specific VARs are combined to draw cross-country inference? Suppose one uses consistently across countries the one standard error shock IRFs. As the standard errors of monetary shocks are in general different across countries, it will not be possible to tell apart the differences in the shocks from the differences in the transmission. A comparison based on these premises will be flawed, because it will not be able to isolate the origin of the movements from the induced movements themselves. Is the bigger response in country A than in country B caused by a bigger initial shock or is it a consequence of underlying structural differences in the transmission? Suppose alternatively that the researcher decides to use the one-percent shock IRFs. In this case the shocks will be equalized across countries, but the relative magnitudes will be modified. Again, this will not allow a meaningful comparison: a one-percent shock may be a "normal" episode in country A, while it may be a completely unprecedented episode in country B[4].

Finally, the identification assumptions for different countries are often difficult to compare and to justify[5].

(b) BIS simulations with national central bank models (BIS (1995)).

This is one of the most important study for the assessment of the differences in the transmission mechanism. Each European country was asked to use the Central Bank model to simulate the effect of an increase in the interest rate. Some countries performed the simulation under the restriction of constant exchange rates between EMU countries; others left the exchange rate endogenous, thus limiting the comparability of the results.

The modelling assumptions differ across countries. On one side, this is a positive feature, because it allows to capture the specific features of each country's economy. On the other side, the differences in the effects of monetary policy reported in this study may be due precisely to the specific features of each model.

1.2 Multi-country models

(a) Output equations (Dornbush-Favero-Giavazzi (1998), Peersman-Smets (1998)).

The first of these studies estimates jointly an output equation, by specifying for each country an equation where output growth is regressed on its own past, on present and past values of the output growth in the other countries, of the *expected* interest rates (identified as the fitted values of a policy reaction function) and of the exchange rates with the DM and with the dollar.

The experiment performed is different from what typically done with VARs. The effects of monetary policy are measured as the responses of output growth to an increase in the *expected*[6]component of interest rates. This component is not the exogenous policy shock, but it is the endogenous response of interest rates to developments in the output gap, in expected inflation and in the exchange rate, as specified in the reaction function they estimate. Their experiment is thus the "complement" of the VAR impulse response analysis.

The specification in Peersman-Smets (1998) adds as regressors the German real interest rate, the spread with the German interest rate and replaces the national exchange rate with respect to the dollar with that between the DM and the dollar. In this case, the focus is on the response of output growth to an increase in the German real interest rate. This experiment once more does not measure the exogenous monetary shock, but captures a mixture of endogenous reactions and exogenous shocks.

It must be said that these two works are the only that can replicate the environment in which the ECB will work: a common monetary action and a fixed exchange rates across countries.

(b) Multi-country VARs (Kieler-Saarenheimo (1998)).

By recognizing the existence of interdependencies across European countries, this work nests together 3 VARs for Germany, France and UK and estimates the system by SURE. A very careful identification exercise is then performed, monetary shocks are identified in each country, but they are left correlated across countries, so that it is impossible to disentangle a German shock from a French shock and to identify the *common* part.

(c) Large models (Fed Model, IMF Model, European Commission Model).

1.3 Evidence from the US

Carlino-DeFina (1998b) base their analysis on previous work (Carlino-DeFina (1998a)) on the state-specific effects of monetary policy in the US. They first estimate separate VARs for each state and identify the response of state personal income to a US-wide monetary shock. Second, they perform cross-sectional regressions of the state-specific responses on a set of explanatory factors, proxies for different channels of the monetary transmission. Third, an index for the sensitivity of European countries to monetary shocks is computed by multiplying the regression coefficients estimated for the US by the values of the corresponding explanatory factors for each European country.

To sum up, many techniques have been employed, but the empirical evidence available does not allow to draw firm conclusions. There is no consensus on what strategy is better suited to answer to the relevant question and different modelling philosophies give very different results (summarized in Dornbush-Favero-Giavazzi (1998) and Guiso et al. (1999)).

1.4 A New Proposal: Factor Models

In this paper we propose a new approach to the study of common monetary policy. By taking an agnostic view concerning the structural model of the different European economies, we ask what is the effect of a common monetary policy shock on different countries by using a Dynamic Factor Model, developed by Forni-Reichlin (1996 and 1998). The Factor Model we use presents some advantages with respect to the widely employed country-specific VARs:

1. It permits the analysis of large panels, pooling different countries together and overcoming the limited amount of information and the interdependence issues.

2. It allows the identification of the *common* monetary shock, and not only of the country-specific, overcoming the endogeneity and the shocks vs. transmission issues.

3. It reduces the number of restrictions needed for identification of the *structural common* shocks, thanks to the reduction of dimensionality of the *common* shocks space, limiting partially the identification assumptions issue.

Concerning the first point, the Factor Model requires the estimation of a limited amount of parameters and at the same time allows for complete heterogeneity in the IRFs to a common shock.

Concerning the second point, we identify the *common* European monetary policy with German monetary policy, by recalling that Germany was the only country able to conduct a completely independent monetary policy during the EMS phase and that the other European countries had to follow the decisions of the Bundesbank: as reported in Dornbush-Favero-Giavazzi (1998), "Europe was on the Buba standard".

Concerning the third point, we will show below that the problem of identification of *structural* shocks in a Factor Model generates the same kind of indeterminacy found in VAR models. Given that there will be few shocks driving the system, less restrictions will be needed. We employ an identification criterion that does not impose any *a priori* zero restriction (the kind of restrictions typically used in the VAR literature) on the parameters of the model. As stressed by Lippi-Reichlin (1994a,b), Faust (1998) and Uhlig (1999), the zero-restriction schemes are only a subset of all the admissible identification schemes and any identification can be generated by the appropriate choice of a "rotation" matrix. Our approach consists in choosing the rotation matrix that minimizes the distance between the IRFs estimated for Germany using the Factor Model and the IRF estimated from a "benchmark" German-only VAR, so to give a precise formal meaning to the assumption that German monetary policy was the *common* European monetary policy[7].

Our method shed light on the pre-EMU period and infer what will happen under the EMU regime[8]. The responses of 8 European countries participating to the Euro project to a *common* monetary shock are consistent with what predicted by a large class of theoretical models. We show that under the EMU regime, Germany, acting as the nominal anchor, put the adjustment costs on the other countries while under the symmetric regime, as the one managed by the ECB, it appears more sensitive to common monetary shocks, as reported in Wieland (1996).

Under the ECB regime, Germany and Spain display strong sensitivity to monetary shocks, while the contrary is true for France, the Netherlands and especially for Italy.

We relate our results to the previous literature and conclude that the interest rate channel is significant in explaining the asymmetries, while the credit channel does not play a significant role, as shown by Carlino-DeFina (1998a) for US states.

Section 2 introduces the econometric model and the estimation strategy. Section 3 discusses the results for the Euro area countries and deals in detail with the identification strategy. Section 4 studies the new environment in which the ECB operates and relates the results to the previous literature. Section 5 concludes.

2 The Model

In this section we briefly discuss the econometric model[9].

Suppose we have a number of sectors, countries or regions, indexed by i. Suppose that each of them can be represented by a *structural* equation of the form:

$$y_t^i = A^i(L)u_t + \varepsilon_t^i \tag{1}$$

where y_t^i is a $(m \times 1)$ zero-mean covariance-stationary vector stochastic process $y_t^i = (y_t^{1i}, y_t^{2i}, ..., y_t^{mi})'$.

$\varepsilon_t^i = (\varepsilon_t^{1i}, \varepsilon_t^{2i}, ..., \varepsilon_t^{mi})'$ is a $(m \times 1)$ vector of country-specific, idiosyncratic shocks, possibly autocorrelated, but mutually orthogonal at all leads and lags across countries, with variances bounded above by the real numbers σ_s^i, $s = 1...m$.

$u_t = (u_t^1, u_t^2, ..., u_t^q)'$ is a $(q \times 1)$ vector of unit variance white noises, the *common structural shocks*, mutually orthogonal and orthogonal to ε_t^i for all i, and with $q < m$.

$A^i(L)$ is a $(m \times q)$ matrix of rational functions in the lag operator L. $A^i(L)u_t$ is the common component. Though it is driven by the common shocks (the u_t's), its dynamic is still specific to the individual unit (note the index i in the $A^i(L)$ filter). This is the component that allows for the heterogeneity in the transmission of the common shocks.

The intuition behind the method is that by using a Law of Large Numbers argument, thus by aggregating across sectors, the idiosyncratic component vanishes with respect to the common. This implies that we will be able to find weighted averages Y_t generated only by the common shocks u_t:

$$Y_t = \begin{pmatrix} \sum_{i=1}^n \omega^{1i} y_t^{1i} / \sum_{i=1}^n \omega^{1i} \\ \sum_{i=1}^n \omega^{2i} y_t^{2i} / \sum_{i=1}^n \omega^{2i} \\ \vdots \\ \sum_{i=1}^n \omega^{mi} y_t^{mi} / \sum_{i=1}^n \omega^{mi} \end{pmatrix} = A(L)u_t \tag{2}$$

where Y_t is a $(m \times 1)$ vector, u_t is a $(q \times 1)$ vector, the matrix $A(L)$ have dimension $(m \times q)$ and the ω^{mi} are weights that will be described in greater detail below.

2.1 Determining the Number of Factors

Define the $(m \times 1)$ vector Y_t of aggregates, as in equation (2), $Y_t = A(L)u_t$, where the weights ω^{hi} are set equal to $1/\sigma^{hi}$, $\sigma^{hi} = Var(y^{hi})$[10].

The dimension of the common shocks space can be recovered as follows. From equation (2), recall that the $(m \times 1)$ vector Y_t is generated by q shocks $(q < m)$.

This in turn means that the $(m \times m)$ spectral density matrix of Y_t , $f_Y(\lambda)$ (with $\lambda \in [0, \pi]$) will have reduced rank q.

We can then decompose the matrix $f_Y(\lambda)$ in terms of its Dynamic Eigenvectors and Eigenvalues[11,12]:

$$f_Y(\lambda) = P(\lambda)\Lambda(\lambda)P(\lambda)' \tag{3}$$

$\Lambda(\lambda)$ is a diagonal matrix containing the Dynamic Eigenvalues (DE) sorted according to their magnitude from the biggest to the smallest, for each frequency λ. The columns of $P(\lambda)$ contain the dynamic eigenvectors associated with each eigenvalue.

The rank of $f_Y(\lambda)$ and the number of common shocks q correspond to the number of eigenvalues different from zero at each frequency[13].

In practice, the dimension of the vector u_t can be recovered by a informal test, by checking graphically the rank of the spectral density matrix, or in other words, by checking how many DE are needed to capture most of the trace of $f_Y(\cdot)$ across frequencies[14].

2.2 Estimation of the Common Component

Having determined the number of factors q, we can estimate the common component as follows. Consider q of the m weighted averages[15]in Y_t and call this vector Y_t^q. We can then rewrite: $Y_t^q = A^q(L)u_t$, where $A^q(L)$ is $(q \times q)$. From $Y_t^q = A^q(L)u_t$, under the assumptions of non-singularity of Y_t^q and fundamentalness[16]of u_t, there exists a finite VAR representation of the form $u_t = A^q(L)^{-1}Y_t^q$. We can then substitute in (1) to obtain:

$$y_t^i = A^i(L)A^q(L)^{-1}Y_t^q + \varepsilon_t^i \tag{4}$$

It is immediate to see that the common components of all the series y_t^i lie in the space spanned by the present and past values of Y_t^q. The common component can then be consistently estimated equation-by-equation by OLS, using as regressors Y_{t-k}^q, $k = 0, ..., K$.

It can be easily proven that the weights ω^{hi} minimizing the variance of the aggregate local component are given by $1/Var(\varepsilon^{hi})$[17]. As ε^{hi} is unknown before estimation, we can follow an iterative procedure:

1. Start with $\omega^{hi} = 1/Var(y^{hi})$, and compute Y_t^q as in (2).

2. Regress y_t^{hi} on Y_{t-k}^q, for $k = 0, ..., K$, get the regression residuals $\hat{\varepsilon}^{hi}$, compute $\omega^{hi} = 1/Var(\hat{\varepsilon}^{hi})$, and a new Y_t^q.

3. Iterate until convergence of all Y_t^q is achieved. At the end of the procedure final estimation of common and idiosyncratic components will be obtained.

2.3 Recovering the Structural Shocks

So far we have estimated two unobserved components: the common and the idiosyncratic. The last step is to identify the *structural* common shocks u_t.

As we have seen, we can represent Y_t^q as a finite order VAR: $A^q(L)^{-1}Y_t^q = u_t$.

We can then estimate it with OLS and apply the identification techniques developed for VARs. The main advantage is that now this techniques are to be applied to a "small" system of dimension q.

More precisely, we can follow the standard steps:

1. Estimate the *reduced form* $B^q(L)^{-1}Y_t^q = v_t$, where $Cov(v_t) = \Sigma$.
2. Orthogonalize the shocks, using the Cholesky decomposition[18]: $\tilde{u}_t = Cv_t$, where $Cov(\tilde{u}) = I_q$
3. Pick among the orthogonal rotation matrices R such that $RR' = I$ the one that satisfies our identification assumptions: $u_t = R\tilde{u}_t$, where $Cov(u) = I_q$

Once the matrix R has been chosen, we can construct the IRFs of all the variables y_t^i to the *common* shocks u_t.

3 European Monetary Policy

We are now ready to use the tools discussed above to study the transmission across European countries.

We use monthly data on the sample 1985:01-1998:12.

The choice of this sample is motivated by the fact that only in the mid-eighties barriers to capital flows, which might have affected the transmission of monetary policy, were lifted. We consider eight European countries: Austria, Belgium, France, Germany, Italy, the Netherlands, Portugal and Spain. For each country we use the following set of variables: $x_t^i = \{IP_t^i, CPI_t^i, INT_t^i, NOM_t^i\}$[19].

IP_t is the log of the Industrial Production Index: $\log(IP_t)$, CPI_t is the consumer price index[20], INT_t is the short-term nominal interest rate and NOM_t is the nominal exchange rate with respect to the US dollar, defined in units of national currency per one dollar.

Being our methodology based on the estimation of the spectral density matrix, which is defined only for $I(0)$ stochastic process, we have to be careful in analyzing the stationarity of our data[21].

According to Augmented Dickey-Fuller unit-root tests, the variables IP_t^i, INT_t^i and CPI_t^i are difference-stationary, while NOM_t^i is $I(0)$ in the sample considered. Table 1 reports the ADF tests for NOM_t^i[22].

[Insert Table 1 about here]

For each country we then consider the vector: $y_t^i = \{\Delta IP_t^i, \Delta INF_t^i, \Delta CPI_t^i, NOM_t^i\}$.

3.1 The Number of Factors

We assume that the model (1) can be rewritten as:

$$y_t^i = B^i(L)A(L)u_t + \varepsilon_t^i \tag{5}$$

with $B^i(L)$ of finite order, so that:

$$y_t^i = B^i(L)Y_t^q + \varepsilon_t^i \tag{6}$$

The first step is to determine the number of factors. We then construct the 4 aggregates Y_t, where Y_t is defined as in equation (2), using weights $\omega^{hi} = 1/Var(y^{hi})$. The cumulated Dynamic Eigenvalues (DE) of the spectral density matrix of the standardized version[23]of Y_t are shown in Figure 1.

[Insert Figure 1 about here]

The frequencies $[0, \pi]$ are reported on the horizontal axis, the percentage of variance explained by the first n DE is reported on the vertical axis.

As there is no formal test for the rank of the spectral density matrix, we have to use a graphical and heuristic criterion: we conclude that our data are characterized by q factors if q DE are sufficient to explain 90% of the variance across frequencies[24].

We conclude that our dataset is characterized by 3 common factors.

3.2 The Common Component

As we have seen above, the common components can be consistently estimated by OLS equation-by-equation. We use the iterative procedure explained above, using as regressors the components of the vector $Y_t^q = \{Y_t^{IP}, Y_t^{INT}, Y_t^{NOM}\}$.

A finite VAR representation for Y_t^q exists only under the assumption of its non-singularity. We check if this assumption is satisfied by the components chosen, by running a Johansen cointegration test on the levels of Y_t^q. The results, presented in Table 2, show that the three aggregates chosen are not cointegrated.

The cointegration relation appearing in the top part of Table 2 is due to the stationarity of Y_t^{NOM}: there exists a trivial cointegration vector: $\{0, 0, 1\}$.

[Insert Table 2 about here]

We perform the test on the two non-stationary components $\{Y_t^{IP}, Y_t^{SHORT}\}$ and we were not able to reject the null of no cointegration among the two I(1) series. Our Y_t^q vector is thus non-singular at frequency zero[25,26]. We use a constant, the contemporaneous value and 2 lags of Y_t^q in the estimation of the common component $B^i(L)$. The initial weights are set to $\omega^{ih} = 1/Var(y^{ih})$. Convergence is achieved after few iterations and the final weights, reported in Table 3, have interesting economic interpretations.

[Insert Table 3 about here]

The only countries with a significant positive weight on the exchange rate are Germany, Austria and the Netherlands.

Austria and the Netherlands were the countries with an exchange rate closely pegged to the Deutsche Mark during the whole sample or, in other words, with a "more common" exchange rate. The weight of the other countries is zero because of the various realignments of the exchange rate that increased their idiosyncratic variance. It is true that in some historical episodes, namely during the September 1992 crisis, the realignments have been "common", as many countries went out of the EMS at the same time and one would expect the weights to reflect this "commonality". It has to be taken into consideration that the weights are constructed not only as cross-sectional averages. They take also into account the time-series properties of the data by means of the OLS regression. They reflect the "typical", the average behavior during the whole sample period. It is by combining the two dimensions that the weights can be given an economic interpretation.

The common part of the exchange rates can be seen in Figures 2 and 3. We show the common components and the original series for NOM for the different countries.

[Insert Figure 2 about here]

[Insert Figure 3 about here]

In Table 4 we also report the variance of the idiosyncratic part. It is evident that the peripheral countries (Italy, Portugal and Spain) display a larger variance for the idiosyncratic part and thus less common exchange rates.

[Insert Table 4 about here]

3.3 The Structural Shocks

The last step is the identification of the monetary shock in the structural VAR: $A^q(L)^{-1}Y_t^q = u_t$, where $Cov(u_t) = I_q$. We estimate the reduced form $B^q(L)^{-1}Y_t^q = v_t$, with $Cov(v_t) = \Sigma$, using 12 lags and a constant.

The rank reduction $q = 3$ reduces the number of restrictions needed. We have to impose only $3 = q(q-1)/2$ restrictions to identify the whole model for all European countries[27].

As we do not have any clear view about how to identify directly the VAR for the aggregates and we do not want to impose any *a priori* zero-restriction, we use an agnostic criterion (for a similar approach, see Lippi-Reichlin (1994a) or Uhlig (1999)), based on orthonormal rotations of the shocks space and on a minimization criterion. Our assumption is that the common monetary shock during the EMS period in Europe can be identified with the German one, so we use this information in the identification strategy.

We fit a VAR for Germany alone, following the specification presented in Bernanke-Mihov (1998). The authors estimate a VAR with $Y_t = \{IP_t, log(WORLD_t), log(CPI_t), INT_t\}$, where the variable $WORLD$ is the Commodity Price Index[28]. We then search for the rotation matrix R that minimizes the distance between the IRF of the German interest rate INT^{GER} to the monetary shock identified from the German-only VAR *à la* Bernanke-Mihov and the IRF of the same variable INT^{GER} to one of the three common shocks u_t generating the Factor Model.

Let us explain in detail our procedure by going one step back. Recall that to identify a VAR we can proceed in the following way. First, we orthogonalize the shocks: $\tilde{u}_t = Cv_t$, where C is $s.t.$ $Cov(\tilde{u}) = I_q$ and second, we choose the orthonormal rotation matrices R, such that $RR' = I$, that satisfies our identification assumptions: $u_t = R\tilde{u}_t$, where $Cov(u) = I_q$

For an m-dimensional system, any rotation matrix R can be parameterized as function of $m(m-1)/2$ parameters.

In our 3 shocks case, any rotation matrix R can be parameterized as follows:

$$R(a,b,c) = \begin{pmatrix} \cos(a) & \sin(a) & 0 \\ -\sin(a) & \cos(a) & 0 \\ 0 & 0 & 1 \end{pmatrix} \begin{pmatrix} \cos(b) & 0 & \sin(b) \\ 0 & 1 & 0 \\ -\sin(b) & 0 & \cos(b) \end{pmatrix} \begin{pmatrix} 1 & 0 & 0 \\ 0 & \cos(c) & \sin(c) \\ 0 & -\sin(c) & \cos(c) \end{pmatrix} \tag{7}$$

with $(a,b,c) \in [0, 2\pi]$[29].

We look for the matrix $R(a,b,c)$ that minimizes the distance between the IRF of INT^{GER} to the German monetary policy shock obtained from the Bernanke-Mihov's VAR, call it $(u^{VAR} \rightarrow INT^{GER})(t)$ and the IRF of INT^{GER} to the shocks $u_i^{FACTOR}(a,b,c)$, $i = 1,..,3$, obtained from the Factor Model, call it $(u_i^{FACTOR}(a,b,c) \rightarrow INT^{GER})(t)$, by using a quadratic loss criterion:

$$\Pi(a,b,c,i) = \sum_{t=1}^{hor} [(u^{VAR} \rightarrow INT^{GER})(t) - (u_i^{FACTOR}(a,b,c) \rightarrow INT^{GER})(t)]^2 \tag{8}$$

The shock identified by the parameters (a,b,c,i) such that the function $\Pi(\cdot)$ reaches its minimum will be denoted as the *common European monetary policy shock*[30]. In Figure 4 we report the IRF obtained from the VAR and the one from the Factor Model such that $\Pi(a,b,c,i)$ reaches its minimum.

[Insert Figure 4 about here]

Having identified the *common shock*, we can now compute the responses to it of all the variables in the system.

Results for each of the 4 variables in the system are reported in Figures from 5 to 8. The response labeled EUROPE represents a weighted average[31] of the country-specific responses.

[Insert Figure 5 about here]

[Insert Figure 6 about here]

[Insert Figure 7 about here]

[Insert Figure 8 about here]

The responses of IP are in line with what predicted by theory: a monetary contraction causes a reduction in industrial production after about 6 months in all countries. It is interesting to note that the average European response to a *common* monetary shock virtually coincides with the German response, thus confirming indirectly our identification strategy concerning the leading role of Germany.

There are nevertheless significant differences in the responses across countries. In order to understand these wide differences we pass to analyze the response of the other variables.

Our experiment allows different endogenous responses of the interest rates. We can see that all the countries increase their interest rates and that all of them have to perform an endogenous monetary restriction stronger than Germany[32]. This behavior can be explained by the need for all countries to show their tightness in the conduct of monetary policy during the EMS period, in order to gain credibility and reduce inflation expectations. Italy and France are the countries characterized by the stronger increases.

The same kind of behavior is evident in the responses of prices. All countries display a deeper response of prices than Germany[33].

Interesting results come from the analysis of the exchange rates. Without imposing any predeterminedness assumptions between interest rates and exchange rates, we see that exchange rates appreciate on impact and they then depreciate to the old equilibrium, as predicted by the Dornbush (1976) overshooting model. We can notice a clear distinction between the peripheral countries, Italy, Portugal and Spain, the countries that adjusted more often the exchange rate parity during the EMS period, and the rest of Europe.

4 The New Environment for the ECB

In the new environment, monetary policy will not be decided taking Germany as the leader. It is likely that the ECB will target a weighted average of country-specific indicators.

In order to shed light on the conduct of monetary policy by the European Central Bank, we perform a counterfactual experiment in the spirit of Sims-Zha (1996), constraining the interest rate responses of all countries to follow the European average response.

Given this constrained interest rate response we recomputed the IRFs for the Industrial Production.

Results are shown in Figure 9.

[Insert Figure 9 about here]

We can notice that Italy and to a lesser extent, France and the Netherlands are now characterized by a smaller-than-average response, while Germany and Spain are characterized by a very large response in comparison to the other countries. The strong endogenous increase in the interest rate in France and Italy in the pre-EMU regime was responsible for much of the restrictive effect of monetary policy on output. The other countries will not be far away from the European average.

We follow Carlino-DeFina (1998a,b) and we relate the cumulated out-of-trend responses of IP to proxies of different theories of the transmission mechanism. Having only 8 observations, we cannot add more than one regressor at time. Given the limited amount of degrees of freedom, our results are to be interpreted with caution; nevertheless, they will provide interesting insights.

Our analysis is similar to the one performed by Mihov (2001). He finds that significant explanatory variables for the strength of monetary policy in his country-specific VAR analysis are the share of the manufacturing sector and some indicators of the credit channel. It is important to notice that his sample is composed of 10 countries, 5 of which are non-Euro countries. Some of his results may be driven by the inclusion of these countries in the analysis. It is unclear whether these results can be significant for the European Union. Second, he controls for the German interest rate, but uses country-specific VARs and computes IRFs in response to a 100-basis point shock to the national interest rate. All the caveats on country-specific VARs are still valid in this case.

We run univariate regressions of the three indexes discussed above, our IRFs computed 24 months after the shock (called Factor), the Mihov index and the Carlino-DeFina index, on various indicators of the interest rate channel and of the credit channel. We use as proxies: TOT, the share of the manufacturing+construction sector (from Carlino-DeFina (1998b)), LOANS1, the ratio of bank loans to total liabilities (from Mihov (2001)), LOANS2, bank loans as a percentage of all forms of finance (from Cecchetti (1999)), SMALL, the percentage of small firms (from Carlino-DeFina (1998b)) THOM, the Thomson Index of bank health (from Cecchetti (1999)), CONC, the concentration ratio of the three largest commercial banks (from Carlino-DeFina (1998b)) and EFFECT, the predicted effectiveness of monetary policy (from Cecchetti (1999)). The figures are reported in Table 5

[Insert Table 5 about here]

[Insert Table 6 about here]

In Table 6 we show the results of the univariate regressions. It is important to notice that regressions with the Mihov index contain at maximum 5 observations.

The only variable that turns out to be significant consistently across studies is the share of the (manufacturing + construction) sector. We thus find evidence for a role of the interest rate channel in explaining the asymmetries. We do not find any significant role for the credit channel and if any, the evidence goes in the opposite direction, as healthier banking systems (measured by low values of the Thomson Index) appear to be very sensitive to monetary policy and the same is true for the Carlino-DeFina index.

Our findings extend their results for US states to European countries. Once we control for the limitations in the country-specific VARs outlined above, the credit channel does not seem to play a first-order effect in explaining the asymmetries. In Table 7 we compare our results with those in CDF. It is evident that the ranking provided by the different methods are very similar, particularly for the "extreme" countries. Regression of Factor on CDF gives a coefficient of 0.45 with a t-statistic of 2.89.

[Insert Table 7 about here]

We can thus conclude with more confidence that in the new environment in which the ECB is operating Spain and Germany will be the countries more sensitive to monetary policy, while Italy and the Netherlands and France will be the less. Results for the other countries show that they will be in the European average.

Interesting observations can be drawn by comparing our results with those in Wieland (1996). In that paper the author uses a large-scale macroeconometric model and shows that an asymmetric monetary regime with Germany as leader, as the one that was functioning during the EMS period, allowed the Bundesbank to reduce output and inflation volatility. The other countries (France and Italy in his paper) had to bear the burden of the adjustment . On the other side, Wieland shows that a symmetric regime in which European averages are taken as targets will reduce variability in France and Italy at the expenses of a higher variability in Germany. This is precisely the same conclusions we can draw by comparing Figure 3.9 with Figure 3.10: the response of Germany is stronger than that of France and Italy under the symmetric rule.

5 Conclusions

In this paper we propose a new way to identify *common* European monetary policy shocks and to analyze the asymmetries in the transmission of the monetary policy managed by the ECB.

Our suggestion is to use a Factor Model. This econometric technique allows us to find the number of common factors in the data. We show that Europe is characterized by the presence of three common factors.

The identification assumption we employ is that European monetary policy followed Germany during the EMS period. We then develop a method to identify the common shocks suited to give formal content to the informal statement about the leading role of Germany in the conduct of monetary policy.

The qualitative results are in line with the literature on the effects of monetary policy in closed-economy. The dynamic responses of European countries show that they will be hit asymmetrically by the same shock.

We conclude that in the new regime, in which monetary policy is likely to be conducted on the basis of the "average" economic conditions in the Euro Area, Spain and Germany will be the countries the more sensitive to monetary policy, while the Netherlands, France and Italy will be the less.

The weight of the interest rate sensitive industries in the different countries appear to be a significant explanatory variable, while credit channel indicators does not seem to play an important role.

A caveat and a proposal for further research are necessary at this point. Our analysis relies, of course, on past data; given that the establishment of the ECB may represent a clear change in regime, we are fully aware of the risks of drawing conclusions regarding the future by looking backwards. We believe that some of the factors influencing the monetary transmission will not change immediately across Europe and that some of them have already changes as agents. A challenging research agenda should then try to study how agents and institutions' behavior will endogenously adapt to the new environment.

6 Appendix 1: Common Factors and Common Trends

This appendix provides a link between the Factor Model we are using and the Common Trends model *à la* Stock-Watson (1988).

Let us briefly recall the Stock-Watson model.

Suppose X_t is a $(n \times 1)$ vector of $I(1)$ processes, whose first differences have the Wold representation: $\Delta X_t = C(L)\varepsilon_t$, with $Cov(\varepsilon) = \Sigma$. This can be rewritten in terms of orthogonal shocks as: $\Delta X_t = C(L)\Sigma^{1/2}\nu_t$, where $Cov(\nu) = I$.

Stock-Watson proved that in presence of $n - q$ cointegration relations among the X_t, the system can be rewritten in the common trends representation:

$$X_t = A\tau_t + D(L)\nu_t \qquad \tau_t = \mu + \tau_{t-1} + \eta_t \tag{A1.1}$$

where A is a $(n \times q)$ matrix of loadings, τ_t is a $(q \times 1)$ vector of common trends, random walks generated by the η_t, which are function of the original shocks ν_t[34].

In this representation, the vector X_t is decomposed in two parts. The first, $A\tau_t$, is driven by q common factors τ_t, random walks with permanent effects on the X_t. The second, $D(L)\nu_t$, is driven by n white noises and has transitory effect on X_t.

The Factor Model we use can be compactly rewritten as:

$$\begin{pmatrix} x_t^1 \\ x_t^2 \\ \vdots \\ x_t^n \end{pmatrix} = \begin{pmatrix} A^1(L) \\ A^2(L) \\ \vdots \\ A^n(L) \end{pmatrix} u_t + \begin{pmatrix} \varepsilon_t^1 \\ \varepsilon_t^2 \\ \vdots \\ \varepsilon_t^n \end{pmatrix} \Rightarrow X_t = A(L)u_t + \varepsilon_t \qquad \text{A1.2}$$

where X_t is a $(nm \times 1)$ vector that stacks the n cross sectional units, each represented by the $(m \times 1)$ vector x_t^i, $A(L)$ is a $(nm \times q)$ polynomial matrix. Using the aggregation techniques we have seen before, we can rewrite it as:

$$X_t^m = A^m(L)u_t$$

where X_t^m is $(m \times 1)$ and $A^m(L)$ is $(m \times q)$ with $m > q$.

As we have seen, the common shocks u_t are not constrained to affect only the long-run behavior of the system, as in the common trends model, but they can generate transitory dynamics.

It is then possible that while q common shocks are present in the system, only $k < q$ of them have permanent effects and are equivalent to the τ_t's in equation (A1.2)[35].

The link between the two models can be then formulated by analyzing the spectral density matrix at frequency zero.

Cointegration can be studied in the time and in the frequency domain. The analysis of the rank of the spectral density matrix of $f_{X^m}(\lambda)$ at zero frequency provides the same information as a time-domain cointegration test on X_t^m (for a formal test, see Phillips-Ouliaris (1988)). The rank of $f_{X^m}(\lambda)$ at frequency zero will then equal the number of common stochastic trends driving the system[36].

Following the approach explained in Section 2.1, this implies that if the system is characterized by k common trends, then only k dynamic eigenvalues will be necessary to explain the trace of the spectral density at frequency zero, while q of them will be necessary at higher frequencies.

As in the presence of cointegration a finite VAR representation for the differences does not exists, the procedure requires the estimation of a VECM for the vector of the q aggregates spanning the space of the common shocks u_t with the imposition of $n-k$ cointegration relations.

7 Appendix 2: Rotations and "Partial" Identification Schemes

In Uhlig (1999), it is shown that an impulse vector[37]can be characterized as:

$$g = \sum_{i=1}^{m} (\alpha_i \sqrt{\lambda_i}) x_i \tag{A2.1}$$

where λ_i and x_i are respectively the eigenvalues and the eigenvectors[38]of Σ and α_i, $i = 1, ..., m$ are coefficients such that: $\sum_{i=1}^{m} \alpha_i^2 = 1$. It is immediate to see that there are $m-1$ degrees of freedom in the choice of g and that the choice of these $m-1$ elements identifies the responses to the shock associated with g.

Turning to the rotation approach, we showed in the text that a complete identification scheme is composed by the product of two matrices CR, where C is any matrix that orthogonalizes the variance-covariance matrix of the shocks (the Cholesky factor we use and the spectral decomposition of Σ employed by UHLIG are two equivalent methods) and R is an orthonormal matrix for $\Re^m$. In the $\Re^3$ case:

$$R(a,b,c) = \begin{pmatrix} \cos b \cos c & -\cos a \sin c + \sin a \sin b \cos c & -\sin a \sin c - \cos a \sin b \cos c \\ \cos b \sin c & \cos a \cos c + \sin a \sin b \sin c & \sin a \cos c - \cos a \sin b \sin c \\ \sin b & -\sin a \cos b & \cos a \cos b \end{pmatrix} \tag{9}$$

One can immediately see that the first column is function only of two parameters or, more generally, of $m-1$ parameters, as in the Uhlig "impulse vector" approach. If one wants to identify only the first shock, he or she has to choose only $m-1$ parameters. By noting that the spectral decomposition[39][40]of Σ gives: $C = X\Lambda^{1/2}$, it is immediate to see that the first column of CR can be parameterized as g in (A2.1) and that the first column of R satisfies the same restriction $\sum_{i=1}^{m} R_{i1}^2 = 1$. Note that as the dimension of the system increases from $m-1$ to m, the number of restrictions required to obtain a complete identification increases by $(m-1)$, while the number of restrictions needed to obtain a "one-shock" identification increases by 1.

We show how to implement a long-run restriction *à la* Blanchard-Quah (1989), using the rotation approach and what are the consequences. Suppose that our system is composed of difference-stationary time-series and that we want to identify only the first shock. Our assumption is that it has a transitory effect on the levels of the first variable. From the preceding analysis we have seen that we need $m-1$ restrictions to uniquely identify this first shock.

The Vector Moving Average (VMA) representation of the reduced form is: $Y_t = B(L)v_t$. The VMA of the structural system is: $Y_t = B(L)CRu_t$.

The identifying restriction is: $B_{1i}(1)(CR)_{i1} = 0$, where $B_{1i}(1)$ and $(CR)_{i1}$ are respectively, the first row of the long-run matrix $B(1)$ and the first column of the matrix product CR. Recall that $(CR)_{i1}$ is function only of $m-1$ parameters. In a bivariate system, as in Blanchard-Quah (1989), this restriction exhausts the degrees of freedom in the choice of g: we obtain identification of the first shock. At the same

[40]The same is true if one uses the Cholesky decompostion.

time, it corresponds to an complete identification scheme. For $m = 2$, $m(m-1)/2 = m-1$: a "one-shock" identification coincides with a complete identification.

In a three-dimensional system, this restriction is not sufficient to get a "one shock" identification and, of course, neither a complete identification, as we need $m(m-1)/2 > m-1 > 1$ restrictions. For the case of the "one-shock" identification, we can however reduce the dimension of the "admissible" rotations (those satisfying the restriction) to the choice of only one parameter, as the other has to satisfy a typically non-linear restriction of the kind $b = f(c)$ in (A2.2). For the case of the complete identification, we still have a reduction of the admissible rotations space. It will be defined by an equation of the form: $a = g(b, c)$.

8 Appendix 3: The Data

Industrial Production Index. OECD database, code: xx2027KSA (where xx denotes the country code).

Consumer Price Index. OECD database, code: xx5241K.

Interest Rate. Austria, Belgium, Germany, Italy, the Netherlands and Portugal: Call Money (Money Market) Rate, IMF database, code: xxx60B.., France, Spain: Call Money Rate ($<$ 24 hours), OECD database, code: xx6207D.

Exchange Rate. OECD database, code: xx7003D.

Footnotes

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2. Kalemli-Ozcan-Sorensen-Yosha (1999) find empirical evidence that financial markets integration across European countries generates more risk-sharing. In turn, this enhances specialization in production, as there is less production risk. This allows countries to exploit the economies of scale. The consequence is that cycles will become less correlated. These asynchronous cycles will affect GDP but not disposable income.
3. For a survey of the monetary VAR literature, see Christiano-Eichenbaum-Evans (1998).
4. If one considers a VAR as a linearized version of an underlying non-linear structure, then the IRFs exercises are meaningful only in a neighborhood of the steady state. If the shock is too big, then one may raise concerns about the reliability of the linear approximation.
5. One of the clearest example is the recursive structure often imposed between the interest rate and the exchange rate. For a discussion, see Bagliano-Favero-Franco (1999).
6. Dornbush-Favero-Giavazzi (1998) show that in their model the residuals of the reaction function, the component that should identify the shocks, capture also risk premia shocks associated with devaluation episodes. This unexpected component turn out to be non-significant in the output equation.
7. Some researchers, recognizing some of the drawbacks we have listed above, have recently suggested that an interesting direction for future research is the kind of modelling strategy that we are going to use. Guiso et al. (1998) write: "There has been very little work on [identifying the responses to monetary shocks] for multiple countries using a common framework. [...] Our reading of the literature is that this kind of study has yet to be done". Kieler-Saarenheimo (1998) write: "One could imagine decomposing the effects to, say, international monetary shocks, country-specific shocks, etc.. [...] We think that structural identification at the multi-country level might be an interesting path for future work".
8. Inferring future behavior from past information is of course subject to the Lucas' critique and this is the case for any modelling strategy. We nevertheless believe that there is still relevant information in historical data, as it is likely that agents adapt

their behavior gradually and that they already started to "discount" the change in the conduct of monetary policy well before the setup of the ECB in January 1999.

9. This section is based on Forni-Reichlin (1996 and 1998).

10. The problem of selecting the weights ω^{hi} is relevant when the cross-sectional dimension is not very large, as it will be in our case. The weights minimizing the variance of the local components are given by $\omega^{hi} = 1/Var(\varepsilon^{hi})$. As $Var(\varepsilon^{hi})$ is not known, then it is reasonable to assume $\omega^{hi} = 1/Var(y^{hi})$. This is similar to what is done in GMM estimation by using the "optimal weighting matrix", which is nothing else then the inverse of the variance-covariance matrix. See also footnote 17.

11. This is the standard decomposition of a matrix in terms of eigenvalues and eigenvectors. One more dimension is added here: the decomposition is performed frequency-by-frequency.

12. For a detailed discussion of the properties of dynamic eigenvalues and eigenvectors and for the related concept of dynamic principal components, see Brillinger (1981).

13. Recall that the rank of a matrix is given by the number of its eigenvalues different from zero. Recall also that: $Trace(f_Y(\lambda)) = Trace(\Lambda(\lambda))$ at any λ.

14. A similar test has been proposed by Phillips-Ouliaris (1988). They propose a formal cointegration test based on the rank of the spectral density matrix at frequency zero, conducted by using the same decomposition.

15. As the rank of Y_t^q has to be equal to q, the component have to be chosen in such a way as to differentiate the aggregates as much as possible.

16. If u_t are not fundamental, the common components are spanned by the past, present and future of the u_t. Only under the assumption of fundamentalness of u_t we can restrict our attention to the present and the past.

17. The intuition is clear: as the weighted average must be such that the idiosyncratic component vanishes, the series with the higher idiosyncratic component must receive a small weight in the aggregation.

18. The spectral decompositon will do the same job. See Appendix 2 for details.

19. Data are described in Appendix 3

20. The raw data are not deseasonalized. The filter $(1 - L^{12})$ is applied to $\log(CPI_t)$ to remove the seasonal component.

21. In VAR analysis, a very important result by Sims-Stock-Watson (1990) simplifies things with this respect. They prove that OLS estimation of a VAR system provides consistent estimates both in the case of stationary and non-stationary variables, and in the latter case, both if the variables are cointegrated or not.

22. The number of lags was chosen so to obtain white noise residuals.

23. Standardization is necessary because otherwise the series with the higher variance will receive a big weight in the construction of the DE, biasing the results towards a smaller number of factors than the true one.

24. By averaging across frequencies, the first three DE explain more than 95% of the total variance.

25. In Appendix 1 the link between common trends and common factors is clarified.

26. We have computed the coherence (a sort of dynamic R^2) between the three components. High coherence means dynamic singularity. Results confirm that the vector Y_t^q is non-singular.

27. If we used a typical 4-variables country-specific VAR for each of the 8 countries, we would have had to impose $8(4(4-1)/2) = 48$ zero restrictions!
28. We identify the model with a simple Cholesky decomposition, while Bernanke-Mihov (1997) use a more "refined" identification, based on information on the institutional framework in which the Bundesbank conducts its monetary policy. The results are essentially the same.
29. See Appendix 2 for more details.
30. The identifying assumption is "light": we just ask that as a consequence of a monetary shock (an increase in the interest rate), the interest rate indeed increases on impact and then decreases (of course following the German response).
31. The weights are given by the share of each country in the European GDP.
32. The monetary restriction in the other countries is endogenous in the sense that according to our identification it was originated in Germany.
33. The IRFs represent the out-of-trend path of the variables in response to shocks. The right way to read the results is that a monetary shock causes prices to grow slower than their underlying drift.
34. This decomposition does not imply automatically a separation of the orthogonal shocks ν_t between transitory and permanent. Cointegration just tells us that some linear combinations are transitory and others are permanent. We need to impose more identifying restrictions in order to separate between permanent and transitory shocks among the ν's (see Blanchard-Quah (1989), King et al. (1991) or Warne (1993)).
35. This is true under the additional assumption that the idiosyncratic components of all the series are $I(0)$.
36. Recall that cointegration implies a rank reduction in the matrix of long run coefficients $C(1)$ of the Wold representation. The spectral density $f_X(\lambda) = \left|C(e^{-i\lambda})\right|^2 \frac{\sigma_\varepsilon^2}{2\pi}$ at frequency zero is: $f_X(0) = |C(1)|^2 \frac{\sigma_\varepsilon^2}{2\pi}$. The rank of $C(1)$ is the magnitude through which cointegration can be studied both in time and in frequency domain.
37. A vector $g \in \Re^m$ is called an *impulse vector* iff there is some matrix A, such that $AA' = \Sigma$, and such that g is a column of A.
38. Normalized to form an orthonormal basis of $\Re^m$.

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Table 1: Unit Root ADF Test

Country	ADF Test Statistic (# lags)
AUT	-3.76*** (4)
BEL	-3.63*** (4)
FRA	-3.70*** (10)
GER	-3.81*** (4)
ITA	-2.74* (12)
NED	-3.78*** (4)
POR	-2.96** (12)
SPA	-2.68* (12)

*: significance at 10%, **: 5%, ***: 1%

Table 2: Cointegration Test on the Aggregates

Cointegration Test on the 3-dim System $\{Y_t^{IP}, Y_t^{SHORT}, Y_t^{NOM}\}$

	Eigenvalue	Lik. Ratio
no coint	0.11	29.76**
1 coint	0.05	11.61
2 coint	0.02	3.16

Cointegration Test on the 2-dim System $\{Y_t^{IP}, Y_t^{SHORT}\}$

	Eigenvalue	Lik. Ratio
no coint	0.06	14.16
1 coint	0.02	3.84

Table 3: Optimal weights from the iterative procedure

	AUT	BEL	FRA	GER	ITA	NED	POR	SPA
IP	0.11	0.11	0.15	0.11	0.14	0.14	0.10	0.14
INT	0.12	0.09	0.16	0.12	0.18	0.12	0.12	0.09
NOM	0.34	0.02	0	0.32	0	0.32	0	0

Table 4: Variance of the Idiosyncratic Component of the Exchange Rates

	AUT	BEL	FRA	GER	ITA	NED	POR	SPA
$Var(\varepsilon_t^{NOM})$	1.7e^{-5}	0.002	0.006	1.9e^{-5}	0.63	3.7e^{-5}	0.69	0.41

Table 5: Proxies for the transmission channels

	AUT	BEL	FRA	GER	ITA	NED	POR	SPA
TOT	0.22	0.21	0.22	0.22	0.32	0.18	0.28	0.25
LOANS1	0.14	0.08	0.08	-	0.09	0.04	-	0.1
LOANS2	0.65	0.49	0.49	0.55	0.5	0.53	0.62	0.58
SMALL	-	0.55	0.56	0.57	0.78	0.61	0.78	0.74
THOM	2.38	2	2.28	1.97	2.57	2.1	2.3	1.79
CONC	0.17	0.44	0.33	0.24	0.28	0.60	-	0.34
EFFECT	2.67	1.33	2.33	2.33	2.67	1.67	2.33	2

Sources: see text

Table 6: Regression Analysis

	CdF	Mihov	Factor
TOT	**3.82 (1.89)**	**11.02 (5.08)**	**2.42 (2.11)**
LOANS	2.03 (0.49)	**8.88 (3.08)**	0.58 (0.26)
LOANS2	1.40 (0.85)	4.06 (0.91)	0.80 (0.82)
SMALL	0.25 (0.22)	0.05 (0.01)	-0.18 (-0.27)
THOM	**-0.71 (-2.42)**	-0.46 (-0.34)	**-0.49 (-3.57)**
CONC	-0.28 (-0.33)	**-2.55 (-1.87)**	-0.19 (-0.36)
EFFECT	-0.16 (-0.79)	0.72 (1.08)	-0.04 (-0.34)

t-ratios in brackets

Table 7: Comparison with Carlino-DeFina (1998b)

	AUT	BEL	FRA	GER	ITA	NED	POR	SPA
CDF	1.62 (5)	1.92 (2)	1.42 (8)	1.90 (3)	1.43 (6)	1.42 (7)	1.90 (4)	2.05 (1)
Factor	0.32 (4)	0.32 (5)	0.30 (7)	0.60 (2)	0.16 (8)	0.30 (6)	0.36 (3)	0.60 (1)

Ranking in brackets

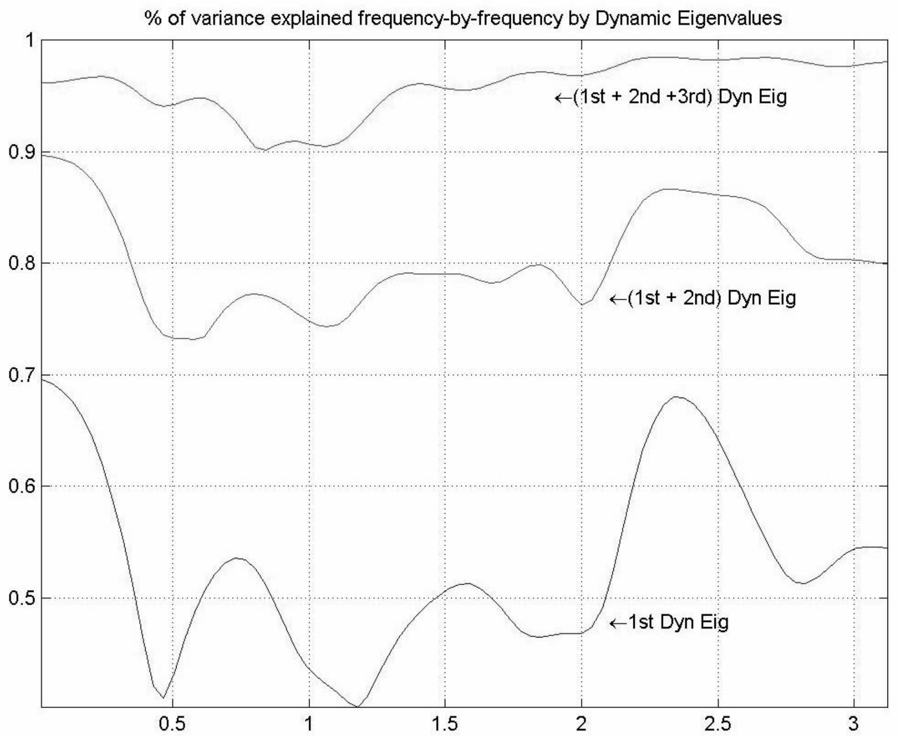

Figure 1: The Number of Factors

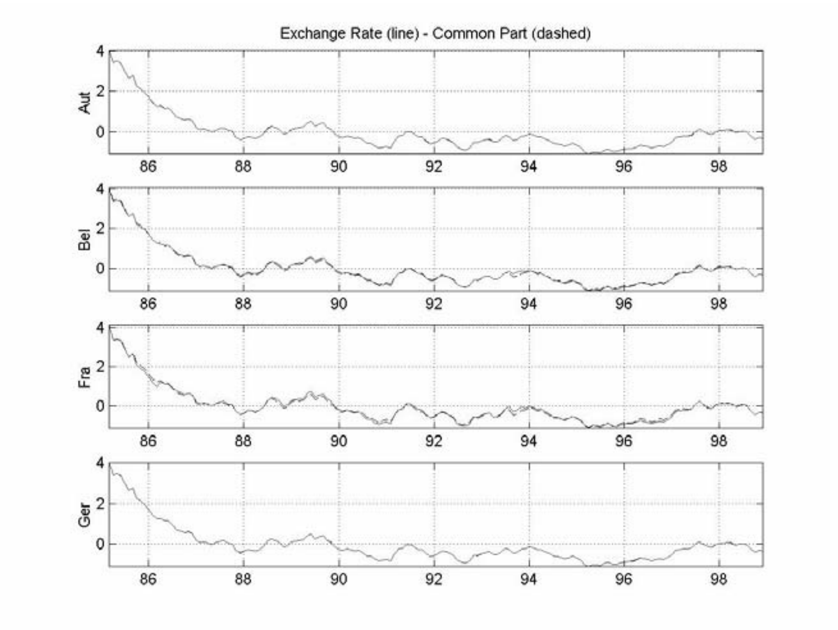

Figure 2: The Common Part of Exchange Rates

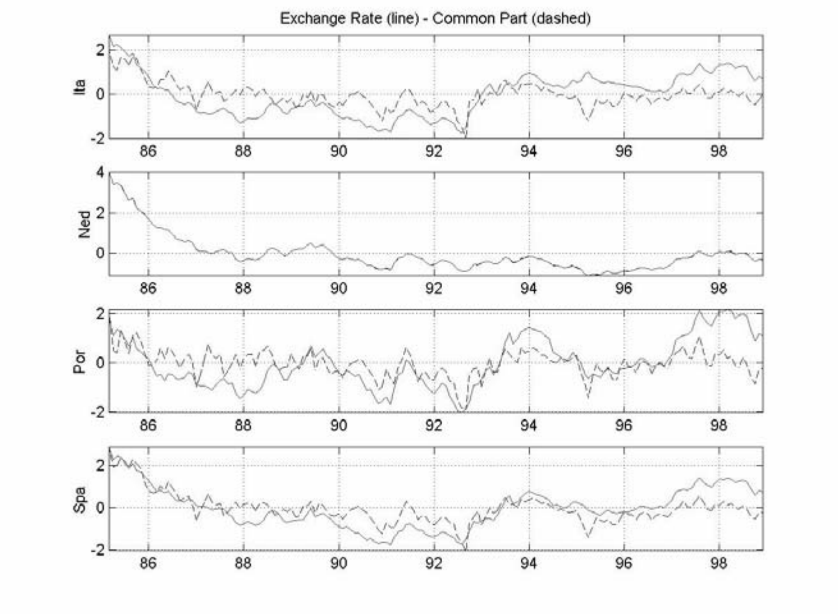

Figure 3: The Common Part of Exchange Rates (continued)

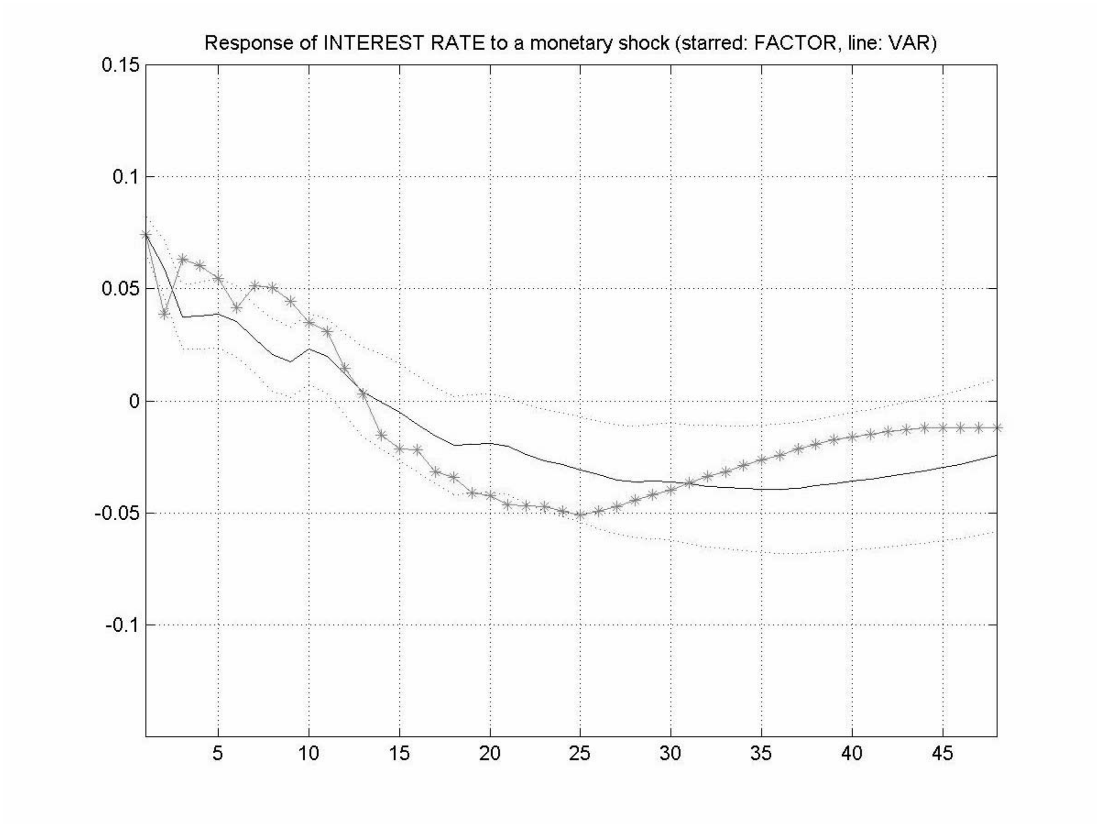

Figure 4: The Result of the Minimization Procedure

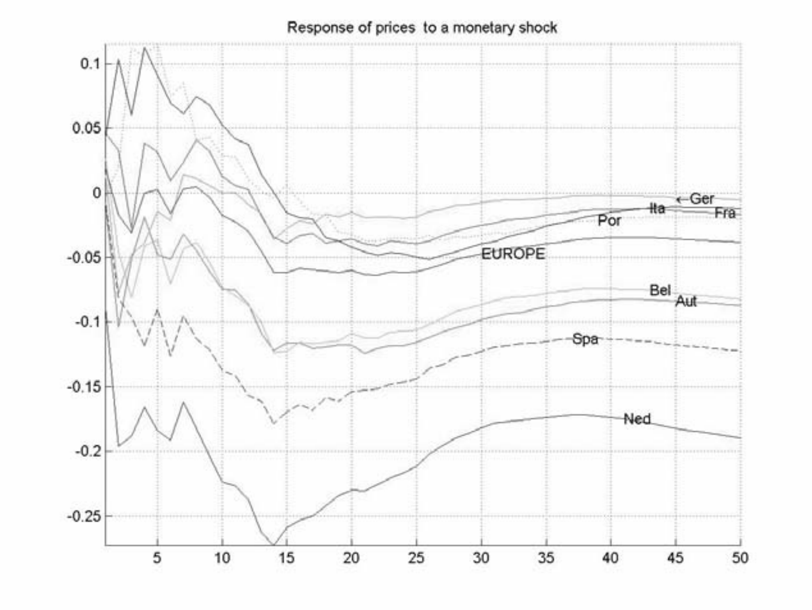

Figure 5: Response of Prices

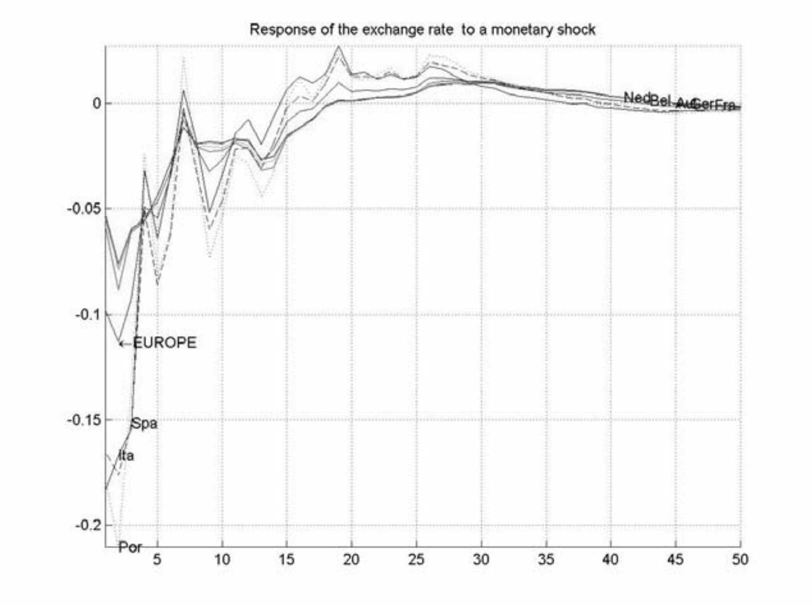

Figure 6: Response of the Exchange Rate

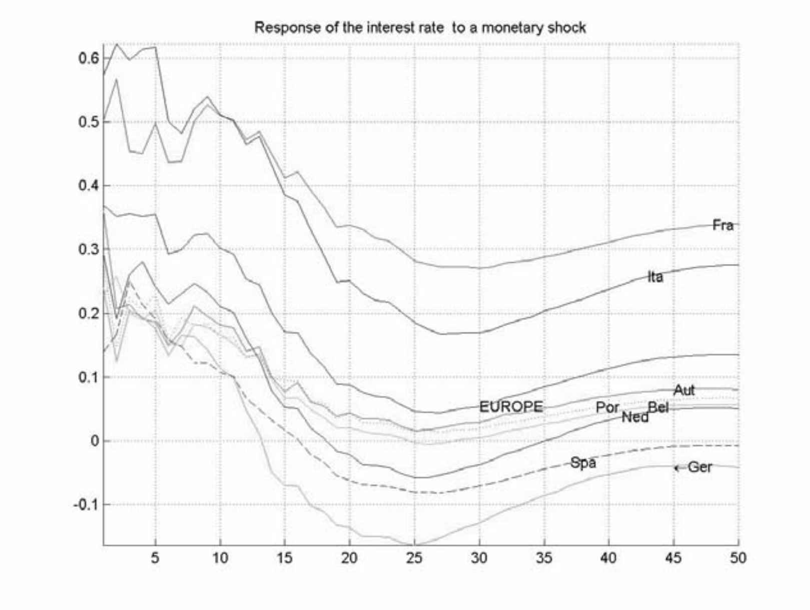

Figure 7: Response of the Interest Rates

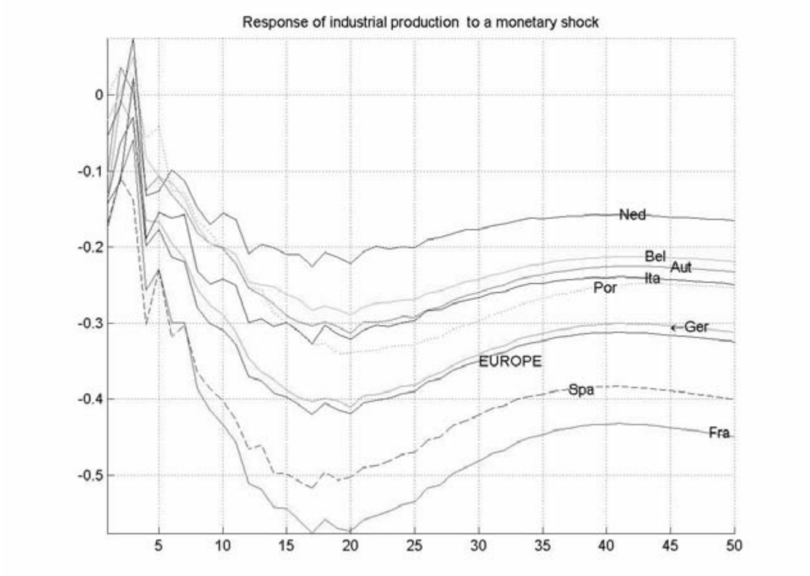

Figure 8: Response of Industrial Production

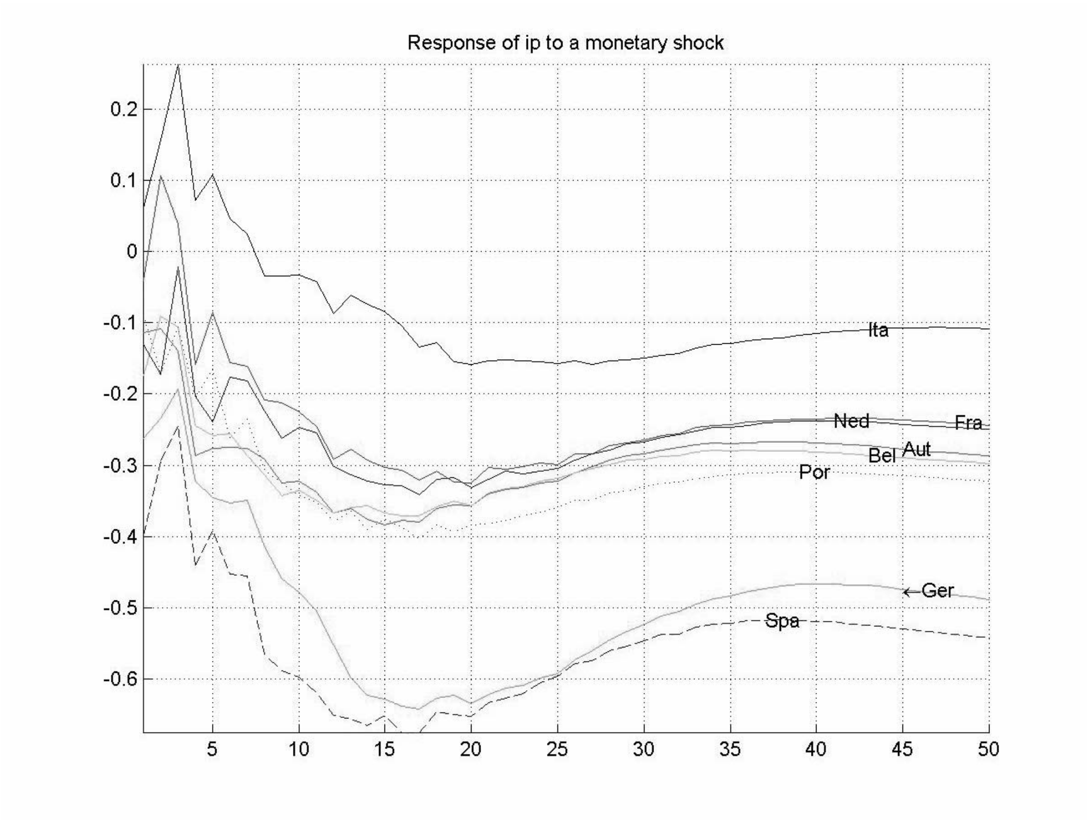

Figure 9: Response to a Common Interest Rate