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How Has the Euro Changed the Monetary Transmission Mechanism?

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Introduction

On January 1, 1999, the euro officially became the common currency for 11 countries of continental Europe, and a single monetary policy started under the authority of the European Central Bank (ECB).[1] The European Monetary Union[2] (EMU) followed decades of monetary policies set by national central banks to serve domestic interests, even though these national policies were constrained by monetary arrangements such as the European Monetary System (EMS), which was designed to limit exchange rate fluctuations. Approaching the tenth anniversary of the EMU, we begin to have sufficient data to potentially observe effects of the monetary union on business cycle dynamics.

This paper has three objectives. The first is to characterize the transmission mechanism of monetary policy in the euro area (EA) and across its constituent countries. The second is to document how this transmission might have changed since the creation of the euro. The third objective consists of providing a set of explanations, based on a structural open-economy model, for the observed differences over time and across countries in the responses of key macroeconomic variables.

Our first two objectives require an empirical model that captures empirically the EA-wide macroeconomic dynamics, while allowing us to estimate the potentially heterogeneous transmission of EA shocks within individual countries. The factor-augmented vector autogression (FAVAR) model proposed by Bernanke, Boivin, and Eliasz (2005) is a natural frame-

ariables. It also allows us to measure the spillovers between individual ountries and the EA.

Many papers have attempted to characterize the dynamics of European conomies. One common strategy has been to model the EA economy sing only EA aggregates. Examples include evidence based on VARs Peersman and Smets 2003), more structural models (the ECB area-wide odel [AWM]; Fagan, Henry, and Mestre 2005), and optimization-based acroeconomic models (Smets and Wouters 2003; Christiano, Motto, nd Rostagno 2007; Coenen, McAdam, and Straub, forthcoming [the ew AWM]). Alternatively, authors have estimated models using country-vel data either to analyze the effects of various macroeconomic shocks r for forecasting, using models of national central banks (Fagan and Iorgan 2006) or VARs (e.g., Mihov 2001; Mojon and Peersman 2003).

An important feature of the FAVAR is that it allows us to model intly the dynamics of EA-wide variables and country-level variables ithin a single consistent empirical framework. In that respect, we see ur empirical strategy as an improvement over the numerous papers hat have compared impulse responses to shocks on the basis of models stimated separately for each country (e.g., Angeloni, Kashyap, and Iojon 2003, chaps. 3, 5). The estimated model suggests that a significant raction of country-level variables such as the components of output and rices, employment, productivity, and asset prices can be explained by A-wide common factors.

In order to characterize the monetary transmission mechanism, we lentify unexpected monetary policy shocks and estimate their dynamic ffects on the national macroeconomic variables. We are particularly in-erested in documenting differences over time and across countries in he sensitivity of national economies to such shocks. (In the appendix o the working paper version of this paper [Boivin, Giannoni, and Mojon 008], we also document the effects of identified oil price shocks.) It is mportant to note that it is not because we believe that monetary policy hocks constitute an important source of business cycle fluctuations that ve are interested in documenting the effects of such shocks. In fact, uch of the empirical literature finds that monetary shocks contribute elatively little to business cycle fluctuations (e.g., Sims and Zha 2006). nstead, monetary policy affects importantly the economy through its

change and assuming that policy is conducted subsequently according to that particular policy rule.

The estimated monetary transmission mechanism is largely consistent with conventional wisdom. A monetary policy tightening in the EA as a whole or in Germany triggers an appreciation of the exchange rate and a downward adjustment of demand and eventually of prices. For the period preceding the EMU, we find considerable heterogeneity in the transmission of these shocks across countries. In particular, we find larger responses of long-term interest rates in Italy and in Spain, which contribute to larger contractions of consumption in these two countries. Also, restrictive monetary policy in the EA tended to trigger a depreciation of the lira and the peseta and a smaller decline of exports of these countries than in the rest of the EA.

The creation of the euro has contributed to a widespread reduction in the effect of monetary shocks. In particular, long-term interest rates, as well as consumption, investment, output, and employment, respond less to short-term interest rate shocks in the new monetary policy regime, whereas trade and the effective real exchange rate respond more strongly. While the monetary transmission mechanism has become more homogeneous along the yield curve, some striking asymmetries persist, for instance, in the response of national monetary aggregates to common interest rate shocks, suggesting pervasive differences in national savings practices.

We use a structural open-economy model to explore some potential explanations for this evolution of the transmission mechanism of monetary policy. More precisely, we extend the model of Ferrero, Gertler, and Svensson (forthcoming) with, among other things, a risk premium on intra-area exchange rates for the period prior to the EMU. This deviation from the uncovered interest rate parity is necessary to replicate a larger response of Italian and Spanish interest rates to German monetary shocks. Using a calibrated version of this model, we show that the combination of two ingredients can replicate the evolution of the estimated transmission mechanism since the start of the EMU: the elimination of the exchange rate premium that plagued some of the European countries by fixing the intra-area exchange rates and a shift in monetary policy, mainly toward a more aggressive response to inflation and output.

he FAVAR and its relation to the existing literature. In Section III, we iscuss the empirical implementation, describing the data used in our stimation, our preferred specification of the FAVAR, and its basic em-irical properties. Section IV studies the effects of monetary shocks in he EA and in individual countries and discusses their changes since the reation of the EMU in 1999. Section V attempts to explain the cross-ountry differences as well as the changes over time in the monetary ransmission mechanism. Section VI presents conclusions.

. Econometric Framework

We are interested in modeling empirically the EA-wide macroeconomic ynamics, while allowing heterogeneity in the transmission of EA hocks within individual countries. A natural framework to achieve his goal is the FAVAR model described in Bernanke et al. (2005). The odel is estimated using indicators from individual European econo-ies as well as from the EA. The general idea behind our implementa-on is to decompose the fluctuations in individual series into a omponent driven by common European fluctuations and a component hat is specific to the particular series considered. EA-wide common hocks can then be identified from the multidimensional common com-onents. The FAVAR also allows us to characterize the response of all ata series to macroeconomic disturbances, such as monetary policy hocks or oil price shocks. Importantly, by modeling jointly EA and ountry-level dynamics, this framework allows each country's sensitiv-y to EA shocks to be different.

. *Description of the FAVAR Model*

We provide here only a general description of our implementation of he empirical framework and refer the interested reader to Bernanke t al. (2005) for additional details. We assume that the economy is af-ected by a vector C_t of common EA-wide components to all variables ntering the data set. Since we will be interested in characterizing the ffects of monetary policy, this vector of common components includes short-term interest rate, R_t, to measure the stance of monetary policy.

actors F_t, where K is relatively small. These unobserved factors may reflect general economic conditions such as "economic activity," the general level of prices," and the level of "productivity," which may not easily be captured by a few time series, but rather by a wide range of economic variables.[3] We assume that the joint dynamics of π_t^{oil}, F_t, and R_t are given by

$$C_t = \Phi(L)C_{t-1} + v_t, \tag{1}$$

where

$$C_t = \begin{bmatrix} \pi_t^{\text{oil}} \\ F_t \\ R_t \end{bmatrix},$$

and $\Phi(L)$ is a conformable lag polynomial of finite order that may contain a priori restrictions, as in standard structural VARs. The error term v_t is independently and identically distributed with mean zero and covariance matrix Q.

The system (1) is a VAR in C_t. The additional difficulty, with respect to standard VARs, however, is that the factors F_t are unobservable. We assume that the factors summarize the information contained in a large number of economic variables. We denote by X_t this $N \times 1$ vector of "informational" variables, where N is assumed to be "large," that is, $N > K + 2$. We assume furthermore that the large set of observable "informational" series X_t is related to the common factors according to

$$X_t = \Lambda C_t + e_t, \tag{2}$$

where Λ is an $N \times (K + 2)$ matrix of factor loadings, and the $N \times 1$ vector e_t contains (mean zero) series-specific components that are uncorrelated with the common components C_t. These series-specific components are allowed to be serially correlated and weakly correlated across indicators. Equation (2) reflects the fact that the elements of C_t, which in general are correlated, represent pervasive forces that drive the common dynamics of X_t. Conditional on the observed short-term interest rate R_t, the variables in X_t are thus noisy measures of the underlying unobserved factors F_t. Note that it is in principle not restrictive to assume that X_t depends

ynamics, such as real activity and inflation) and by series-specific components unrelated to the general state of the economies, e_t. For instance, 2) specifies that indicators of country-level economic activity or inflation are driven by a European interest rate, EA latent factors F_t, and a omponent that is specific to each individual series (representing, e.g., easurement error or other idiosyncrasies of each series). The dynamics f the EA common components are in turn specified by (1).

As in Bernanke et al. (2005), we estimate our empirical model using a ariant of a two-step principal component approach. In the first step, e extract principal components from the large data set X_t to obtain onsistent estimates of the common factors.[5] Stock and Watson (2002) nd Bai and Ng (2006) show that the principal components consistently ecover the space spanned by the factors when N is large and the number of principal components used is at least as large as the true number f factors. In the second step, we add the oil price inflation and the hort-term interest rate to the estimated factors and estimate the structural VAR (1). Our implementation differs slightly from that of Bernanke al. since we impose the constraint that the observed factors (π_t^{oil} and $_t$) are among the factors in the first-step estimation.[6] This guarantees hat the estimated latent factors recover dimensions of the common dynamics not captured by the observed factors.[7]

This procedure has the advantages of being computationally simple nd easy to implement. As discussed by Stock and Watson (2002), it also mposes few distributional assumptions and allows for some degree of ross-correlation in the idiosyncratic error term e_t. Boivin and Ng (2005) ocument the good forecasting performance of this estimation approach ompared to some alternatives.[8]

Interpreting the FAVAR Structure

arious approaches have been used in the literature to model macroconomic dynamics in the EA. As we illustrate in this subsection, these pproaches can be interpreted as special cases of the FAVAR framework. Our approach thus merges some of the strengths of these existing pproaches and allows us to answer a broader set of questions.

As in Bernanke et al. (2005) and in Boivin and Giannoni (2006a), we

other variables in two regions. In terms of the notation in our empirical framework, all these variables would either be included in C_t or be linear combinations of the elements of C_t. The dynamic evolution of these variables can be approximated by a VAR of the form (1).[9]

The existing approaches that model the dynamics of EA variables can be interpreted as special cases of the FAVAR model, in the case in which the elements of C_t are perfectly observed, so that the system (1)–(2) boils down to a VAR. Interpreted in this way, the various existing empirical models differ about the assumptions they make about the variables included in C_t, the indicators used to measure C_t, and the restrictions imposed on the coefficients of (1)–(2).

One approach is to assume that the elements of C_t are observed and correspond to EA aggregates.[10] Such a model can be estimated directly using a VAR on EA aggregates only (e.g., Peersman and Smets 2003) or a constrained version of a VAR corresponding, for example, to the ECB AWM (Fagan et al. 2005) or even optimization-based macroeconomic models (Smets and Wouters 2003; Christiano et al. 2007; Coenen et al., forthcoming [the new AWM]). Models estimated only on EA aggregates are silent about the regional effects of a shock.

A second approach is to assume that the elements of C_t are observed and correspond to variables of different regions. In that case, the FAVAR boils down to multicountry VARs and could be estimated directly, as in, for example, Eichenbaum and Evans (1995) and Scholl and Uhlig (2008).

A third approach is to assume that elements of C_t are observed and correspond to variables of a specific country. A large literature has in fact analyzed the cross-country differences in the response of monetary policy using country-level models that are estimated separately (see Guiso et al. [1999], Mojon and Peersman [2003], Ciccarelli and Rebucci [2006], and references therein). By construction these models focus on country-specific shocks and do not explicitly identify the effects of EA-wide shocks such as changes in the stance of monetary policy that would affect all countries simultaneously. The transmission of such shocks could potentially be amplified through trade and expectation spillovers.[11]

Importantly, in all these cases, since the variables necessary to capture the EA dynamics are observed, there is no need to use the large set of in-

monetary models cannot be measured directly. By using a large data set, one is able to extract empirically the components that are most important in explaining fluctuations in the entire data set. While each common component does not need to represent any single economic concept, the common components C_t should constitute a linear combination of all the relevant latent variables driving the set of noisy indicators X_t to the extent that we extract the correct number of common components from the data set.

An advantage of this empirical framework is that it provides summary measures of the state of these economies at each date, in the form of factors that may summarize many features of the economy. We thus do not restrict ourselves to summarizing the state of the economies with particular measures of inflation and of output. Another advantage, as Bernanke et al. argue, is that this framework should lead to a better identification of the monetary policy shock than standard VARs, because it explicitly recognizes the large information set that the central bank and financial market participants exploit in practice and also because, as just argued, it does not require one to take a stand on the appropriate measures of prices and real activity that can simply be treated as latent common components. Moreover, for a set of identifying assumptions, a natural by-product of the estimation is to provide impulse response functions for any variable included in the data set. This is particularly useful in our case since we want to understand the effects of macroeconomic shocks on a wide range of economic variables across EA countries.

Other papers have in fact followed a similar route. Sala (2001) estimates the effects of German and EA composite interest rate shocks using a factor model. He stresses large asymmetries in the response of either output or prices to this shock. Favero, Marcellino, and Neglia (2005) compare the effects of monetary policy shocks on output and inflation in Germany, France, Italy, and Spain for alternative specifications of factor models. They find largely homogeneous effects on output gaps and inflation rates across countries. Eickmeier (2006) and Eickmeier and Breitung (2006) characterize the effects of common shocks on GDP and inflation in 12 countries of the EA and in new European Union members that will adopt the euro in the future. They conclude

cross countries of the EA. In that regard, we believe that countries of he EA, and their move toward a common currency, provide a unique xperiment for monetary economists. For this reason our focus is not rictly on the response of countries' GDPs and inflation rates, but on any relevant dimensions of the economy. We thus seek to take full dvantage of the FAVAR structure to document the effect of various hocks on various measures of real activity, such as GDP and its com-onents, employment and unemployment, various inflation measures, nd financial variables. Although our scope is broader, our approach is imilar to that used by McCallum and Smets (2007), who use a similar AVAR to study how the responses of wages and employment to mone-ary shocks in the EA depend on national and sectoral labor market haracteristics.

II. Empirical Implementation

A. Data

he data set used in the estimation of our FAVAR is a balanced panel f 245 quarterly series, for the period running from 1980:1 to 2007:3. We limited the sample to the six largest economies of the EA, that is, ermany, France, Italy, Spain, the Netherlands, and Belgium, for which e could gather a balanced panel of 33 economic quarterly time series hat are available back to 1980. Given that these countries account for 0% of the EA population and output, we deem it unlikely that the in-usion of other EA countries would alter our estimates of EA business ycle characteristics.

The 33 economic variables that we gathered for each country and the A include two interest rates, M1, M3, the effective exchange rate, an ndex of stock prices, GDP, and its decomposition by expenditure, the ssociated deflators, producer price index and consumer price index CPI), the unemployment rate, employment, hourly earnings, unit labor ost measures, capacity utilization, retail sales, and number of cars sold. n addition to these 231 country-level and EA-level variables, we also nclude an interest rate and real GDP for the United Kingdom, the nited States, and Japan; the euro/dollar exchange rate; an index of

iven in the appendix. The graphs of the data are available in the appendix of the worker paper version of this paper.

We take year-on-year (yoy) growth rates of all time series except for nterest rates, unemployment rates, and capacity utilization rates. The oy transformation is preferred to limit risks of noise due to improper r lack of seasonal adjustment in the data.

Sample Period

he choice of the sample period is delicate. On the one hand, our interst lies in characterizing the monetary transmission in the period since he start of the monetary union in January 1999. We therefore have bout nine years of data that correspond to the strict monetary union. lowever, the objective of stabilizing exchange rates within what would ecome the EA started much earlier. In fact, already in the 1970s, European governments set up mechanisms that aimed at limiting exchange ate fluctuations within Europe.[13] The march to the monetary union has een gradual, and each country has progressed at its own speed. The egs of Austria, Belgium, and the Netherlands to the deutsche mark ere not realigned after the early 1980s. The last realignment of the rench franc to core EMS currencies (the deutsche mark, the Belgian anc, and the Dutch guilder) took place in January 1987. Ex post, we now that the parity between the French franc, the Belgian franc, the utch guilder, and the deutsche mark hardly changed at all since nuary 1987. However, a significant risk premium for fear of realignent plagued the French currency until 1995. Finally, countries such as aly and Spain—as well as Greece, Portugal, Ireland, and Finland, hich are not in our sample—saw their currencies fluctuate vis-à-vis eir future partners in the monetary union well into the 1990s. lthough interest rates remained much higher in Italy and Spain than Germany up until the mid-1990s because of risk premia, changes in he interest rates set by the Bundesbank would be echoed in domestic onetary conditions because of the official peg to the deutsche mark.

Another key aspect of the process of monetary integration is the dere of nominal convergence. Inflation rates were much further apart in he 1970s and early 1980s than ever since.

Preferred Specification of the FAVAR

For the model selection, the short sample size severely constrains the class of specifications we can consider, especially the number of lags in (1), as well as the number of latent factors. We were thus forced to consider models with no more than eight factors and three lags. Among those, our approach has been to search for the most parsimonious model for which the key conclusions that we emphasize below are robust to the inclusion of additional factors and lags. On the basis of this, our preferred specification is one with a vector of common components C_t containing five latent factors in addition to the short-term interest rate and oil price inflation and a VAR equation (1) with one lag. As we show below, these common factors explain a meaningful fraction of the variance of country-level variables.

European Factors and EA Countries' Dynamics

To assess whether our FAVAR model provides a reasonable characterization of the individual series, we now determine the importance of area-wide fluctuations for individual countries. Note that from equation (2), each of the variables X_{it} of our panel can be decomposed into a component $\lambda_i' C_t$ that characterizes the effects of EA-wide fluctuations and a component e_{it} that is specific to the series considered:

$$X_{it} = \lambda_i' C_t + e_{it}. \tag{3}$$

It is important to note that each variable may be affected very differently by the multidimensional vector C_t summarizing EA-wide fluctuations, since the estimated vectors of loadings λ_i may take arbitrary values. We first start by determining the extent to which key European variables are correlated with EA factors over three samples. We then discuss how the importance of these factors has changed over time. In the next section, we document how monetary shocks get transmitted to the EA and across the different countries.

Several studies have recently attempted to determine the degree of comovement of a few macroeconomic series across countries.[14] Forni

re somewhat less correlated with EA-wide common factors than their A counterparts. However, Agresti and Mojon (2003) show that the omovement of either consumption or investment across EA countries smaller than the comovement of GDP. Hence there is a possibility that he tightness of economic variables with the EA business cycle may be neven across countries and across variables of different kinds. This is hy we consider a large number of economic variables rather than a ouple of macroeconomic indicators in our analysis.

Comovements between European Variables and EA Factors

able 1 reports the fraction of the volatility in the series listed in the ble that is explained by the seven EA-wide factors C_t (i.e., five latent actors, the log change of the oil price, and the EA short-term interest te). This corresponds to the R^2 statistics obtained by the regressions of hese variables on the appropriate set of factors.

Columns 1–3 report the R^2 statistics obtained by regressing the re-pective EA-wide series on the common factors for our entire sample, subsample representing the period preceding the monetary union, nd the sample starting in 1999 representing the period in which the MU is in place. These numbers indicate that most of the variables sted are strongly correlated with the common factors, both before nd after the monetary union.[15] While the short-term interest rate is a ommon factor by assumption, other key variables such as EA real DP growth, CPI inflation, bond yields, and the unemployment rate l have R^2 statistics above 0.9. The common factors therefore summa-ze quite well the information contained in these EA series. Not all se-es, however, are as strongly correlated with the common factors. For stance, the growth rate of the monetary aggregate M1 and public con-umption for the EA, with R^2 statistics of only 0.43 and 0.54, display uch less comovement with the common factors.

Instead of estimating latent factors from our large data set, we could lternatively impose key EA macroeconomic variables such as GDP, onsumption, inflation, exchange rate, bond yield, and unemployment s observed factors. Our proposed approach dominates, however, since he latent factors explain a substantially larger share of the variance of

Table 1
R^2 for Regressions of Selected Series on Common Factors

	Euro Area			Average R^2 over Countries		
	1987:1–2007:3 (1)	1987:1–1998:4 (2)	1999:1–2007:3 (3)	1987:1–2007:3 (4)	1987:1–1998:4 (5)	1999:1–2007:3 (6)
Short-term interest rate	1.00	1.00	1.00	.97	.97	1.00
Bond yield	.96	.96	.95	.94	.94	.95
Stock price	.65	.71	.91	.61	.77	.88
Real exchange rate	.78	.82	.93	.73	.79	.93
M1	.43	.65	.73	.42	.65	.53
M3	.70	.92	.74	.50	.71	.69
Deflator GDP	.88	.89	.88	.73	.81	.84
Deflator personal consumption expenditure	.88	.90	.72	.77	.90	.83
Deflator investment	.89	.93	.88	.63	.71	.75
Deflator exports	.86	.80	.97	.72	.71	.94
Deflator imports	.93	.89	.99	.82	.78	.97
CPI	.94	.97	.90	.78	.91	.83
Real GDP	.94	.97	.96	.79	.84	.90
Consumption	.88	.92	.90	.71	.75	.81
Public consumption	.54	.71	.54	.42	.59	.63
Investment	.92	.93	.94	.65	.76	.78
Exports	.70	.71	.93	.67	.68	.88
Imports	.84	.95	.93	.74	.81	.89
Employment	.85	.90	.97	.78	.85	.85
Unemployment rate	.92	.97	.96	.86	.93	.96
Hourly earnings	.94	.97	.69	.79	.92	.74
Unit labor costs	.89	.96	.88	.81	.92	.89
Capacity utilization ratio	.86	.92	.95	.67	.80	.77
Retail	.73	.80	.60	.53	.67	.60

than those for the entire EA area, as expected, to the extent that each country has country-specific features that are not summarized by the common factors C_t and that tend to average out when considering the EA as a whole. Nonetheless, the table shows that, on average, over the six European countries, most of the variables are also strongly correlated with the common factors. Again, for the entire sample, country-level measures of GDP growth, short and long interest rates, inflation, employment, and unemployment all show, on average, high degrees of comovement with the common factors, whereas growth rates of M1, M2, and public consumption show much lower degrees of comovement.

ide range of variables in our panel, then the R^2 for the monetary aggregate series should be high even if no such series was used for the estimation of the latent factors. In this case, the estimated latent factors should e capturing the common movements in the data that are generated by uctuations in the monetary aggregates. So in theory, provided that we low for a sufficiently large number of latent factors, the composition of he panel should not matter for the estimation of latent factors.

Looking across countries reveals that the correlation with the common factors is broadly similar across countries in each of the subsamples. Table 2 reports the average R^2 statistic for each country across the ariables listed in table 1. It shows that country-level R^2's vary between 64 and 0.77 for the entire sample, between 0.74 and 0.84 in the first ubsample, and between 0.78 and 0.87 in the post-EMU sample.

Table 2 also shows that in the case of Germany, the Netherlands, and elgium, the R^2's are sensibly lower for the entire sample than for each f the subsamples considered. This suggests that the relationship between the variables in those countries and the common factors must ave changed between the pre-1999 and post-1999 periods. Finally, e observe that Italian and Spanish variables have become somewhat ss tied to EA-wide developments over time. This comes essentially om the growth rates of real variables. This is particularly clear for pain, since its GDP has grown at a faster pace than that of the rest of he EA since 1995, but is less clear-cut for Italy, which has grown slightly ss rapidly than the rest of the EA.

V. Monetary Policy Regimes and the Monetary Transmission Mechanism

n the last section, we have documented that the variables of each individual country have been on average fairly highly correlated with the

able 2
verage R^2 for Regressions of Selected Series on EA Factors

	1987:1–2007:3 (1)	1987:1–1998:4 (2)	1999:1–2007:3 (3)
uro area	.83	.88	.87
ermany	.69	.77	.82

EA-wide common factors. Nonetheless, aggregate shocks affecting the entire EA may have different implications for each individual country. To assess this, we use our estimated FAVAR to characterize the effects of monetary policy shocks, which we measure here as an unanticipated increase in the EA short-term interest rate of 100 basis points on the national economies considered. Our empirical model is well suited for this since it allows us to determine simultaneously the effects of such shocks on all country-level variables.

As mentioned above, the data reveal changes over time in the degree of comovement of key European variables with EA-wide common factors. A natural implication of such changes is that the transmission of monetary policy may have evolved over time. We thus report the effects of monetary policy shocks both for our benchmark sample and for the post-EMU period. The description of the effects of this shock is a natural starting point in a context in which several countries have chosen to adopt a common currency and therefore to submit their economy to a single monetary policy.

Identification

To identify monetary policy shocks, we proceed similarly to Bernanke et al. (2005) by assuming in the spirit of VAR analyses that the latent factors F_t and the oil price inflation π_t^{oil} cannot respond contemporaneously to a surprise interest rate change, whereas the short-term rate r_t can respond to any innovation in the factors F_t or in oil prices. Of course, we do not restrict in any way the response of factors F_t and π_t^{oil} in the periods following the monetary shock. This constitutes a minimal set of restrictions needed to identify monetary policy shocks. We also impose that all prices and quantity series respond to monetary policy only through its lagged effect on F_t (and potentially π_t^{oil}). This guarantees that none of these variables responds contemporaneously to unexpected monetary shocks, as is often assumed. These restrictions do not, however, prevent any of the financial variables such as bond yields, stock prices, and exchange rates from responding contemporaneously to the short-term interest rate.

In the next section, we present a theoretical model that is designed to

Our assumption that the monetary policy instrument is the short-erm EA interest rate is certainly appropriate for the post-EMU period uring which the ECB has set the short-term EA interest rate. It may be ss appropriate, however, for the pre-EMU period, during which each ational central bank could in principle choose its own interest rate. As Peersman and Smets (2003), Smets and Wouters (2003), and many thers, during the pre-EMU period, our monetary policy shock is a ficti-ous shock that we estimate would have been generated by the ECB ad it existed.

In the pre-EMU period, the German central bank (i.e., the Bundesbank) ssumed a central role in setting the level of interest rates for all countries articipating in the EMS. Given the Exchange Rate Mechanism (ERM) in lace, which limited fluctuations in nominal exchange rates, most of the ther national central banks had to respond to changes in interest rates y the Bundesbank. For this reason, we verified the robustness of our re-ults for the pre-EMU period by identifying a monetary policy shock as a urprise increase in the German short-term interest rate. The results ob-ained are briefly described in Section IV.E, which discusses the robust-ess of our results, and are reported in the appendix of the working paper ersion of this paper.

Effects of Monetary Policy Shocks in the Euro Area in the 1988–2007 Period

igures 1*a*–1*c* report the estimated impulse responses to an unexpected 00 basis point increase in the EA short-term interest rate. While the ark solid lines plot the responses of the variables in each country for he full sample of 1988–2007 along with the 90% confidence intervals dotted lines), the dashed lines plot the responses for the post-EMU period arting in 1999. The figures plot in a column the responses of a particular ariable. The first five plots in each column show the impulse responses in he EA, Germany, France, Italy, and Spain. The bottom two plots combine he responses for all countries in the two different samples. They reveal the ifferences across regions in each sample.

We first start by describing the response of the EA economy in the 988–2007 period by focusing on the plots in the first row. These plots

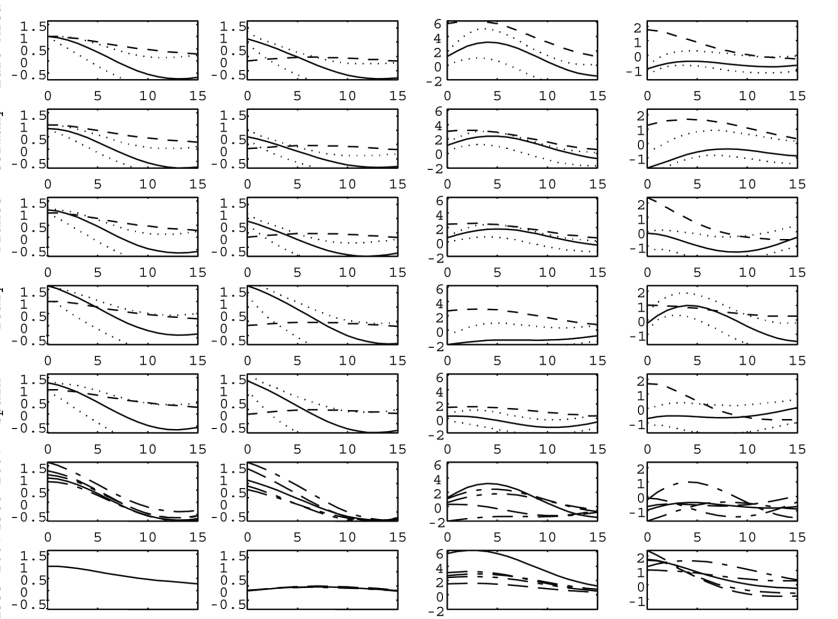
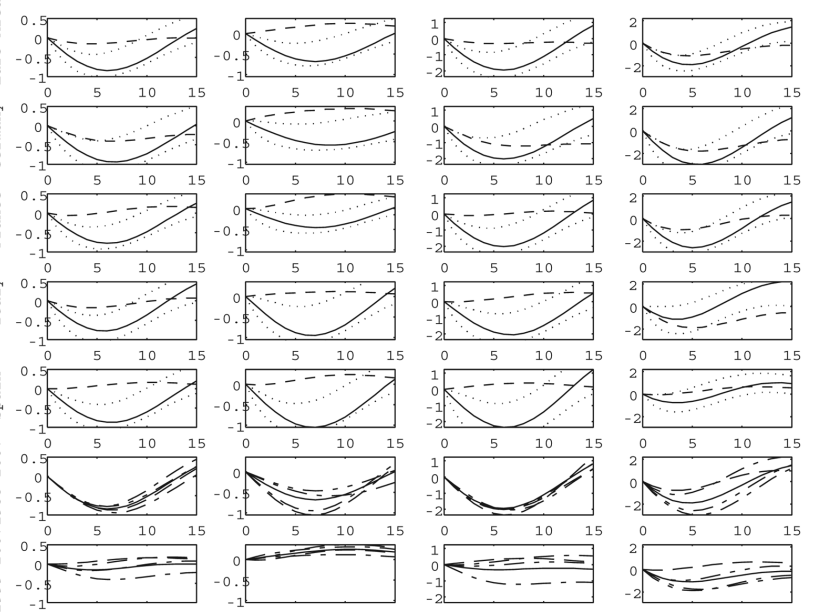


Fig. 1*a*. Impulse response functions to a monetary tightening in EA (shock equals 100 basis point increase in short-term rate; responses are expressed in year-over-year growth rates except for interest rates).

The real GDP yoy growth rate falls by about 1% after a year and a half and does not revert to a positive value before 3 years. Our point estimate of the impact of monetary policy on output tends to be larger than in Smets and Wouters (2003) and various estimates reported in Angeloni et al. (2003). The large drop in output reflects a broad-based decline in aggregate consumption, investment, and exports.[16] The decline in overall economic activity is furthermore clearly reflected in a fall in employment reaching about 0.7% after 6 quarters and a subsequent increase in the unemployment rate. It is followed by a reduction in hourly earnings and in CPI inflation.

Cross-Country Differences in the 1988–2007 Period



g. 1*b*. Impulse response functions to a monetary tightening in EA (shock equals 100 asis point increase in short-term rate; responses are expressed in year-over-year growth tes).

ountries.[17] This heterogeneity gets amplified when looking at long-erm yields. In fact, the Italian and Spanish bond yields rise almost vice as much as the yields of some other countries such as Germany, rance, or the Netherlands.

Consistent with the larger rise in bond yields in Italy and Spain over he whole sample and with the interest rate parity condition, the Italian nd Spanish currencies depreciate with respect to the other countries' urrencies in the pre-EMU period. The Italian and Spanish real effective xchange rates depreciate on impact and in subsequent quarters, whereas the price levels remain unchanged in the period of the shock igs. 1*a* and 1*c*).[18] Instead, all the other countries see their real exchange ates appreciate on impact and for several quarters after the shock, in esponse to the monetary tightening.

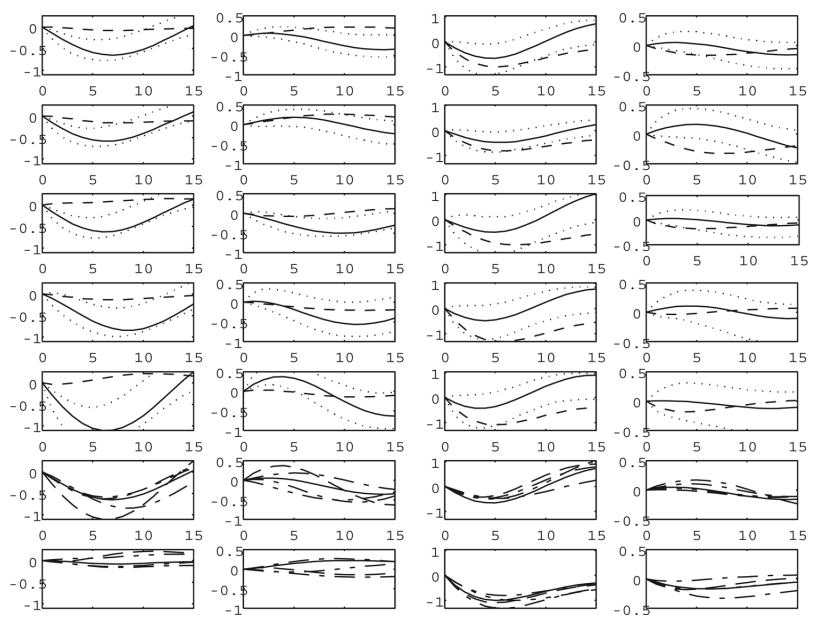


Fig. 1*c*. Impulse response functions to a monetary tightening in EA (shock equals 100 basis point increase in short-term rate; responses are expressed in year-over-year growth rates).

and investment also falls more. The depreciation of the Italian and Spanish real exchange rates, however, mitigates the fall in exports, thus contributing to a more homogeneous output response. These figures thus clearly reveal how diverse responses of bond yields and exchange rates affect differently the various European economies when we consider economic adjustments in the pre-EMU period.

We note that the responses of CPI inflation reveal a temporary "price puzzle" in Germany and Italy following a tightening of the artificial EA interest rate. While the price increases may be explained in Italy by the exchange rate depreciation—a feature that the model we present below is able to replicate—the price increase in Germany is more difficult to rationalize. One possibility is that the artificial EA interest rate may not properly capture surprise monetary shocks for Germany. In fact, when

ery similar to the ones reported in our benchmark specification in gures 1*a*–1*c*.

Finally, it should be stressed that the effects of interest rate shocks on I3 (as well as on M1) are quite different across countries. We have seen Section III.E that the monetary aggregates are markedly more loosely elated to the common factors than most other variables under consid- ration. This may reflect the pervasive differences in the national habits nd in the availability of savings instruments across countries of the A. The ECB (2007) report on financial integration points to, among ther things, the large differences in financial assets of household sec- rs across countries (from four times annual consumption in Belgium nd Italy to only twice in France and Germany), large differences in the omposition of financial wealth, and different pass-through of the mar- et interest rate to deposit interest rates (see Kok Sørensen and Werner 2006] and references therein).

As we noted, the responses that we have documented reveal much rger increases in interest rates and sharper drops in consumption in aly and Spain than in the other EA countries. Italy, for instance, was ubject to considerable speculative attacks in the early 1990s. That rced the Bank of Italy to increase short-term rates considerably more han, for example, in Germany, in order to defend its currency—thereby ading to a more important contraction of economic activity—until it ad to abandon the ERM in September 1992. One might thus wonder hether the effects that we uncovered are due to this unusual event hat was the crisis of the ERM. To investigate this question, we reesti- nated the impulse response functions for the entire sample, except that e excluded the observations from the third quarter of 1992 to the sec- nd quarter of 1993. We find that the responses of short- and long-term nterest rates are almost identical to the one reported in figure 1*a*. The nly notable difference is that the response of consumption is slightly maller in all countries, but we still observe a much larger contraction f consumption in Italy and Spain than in the other EA countries. So the acts that we have documented do not appear to be simply an artifact of few observations around the ERM crisis.

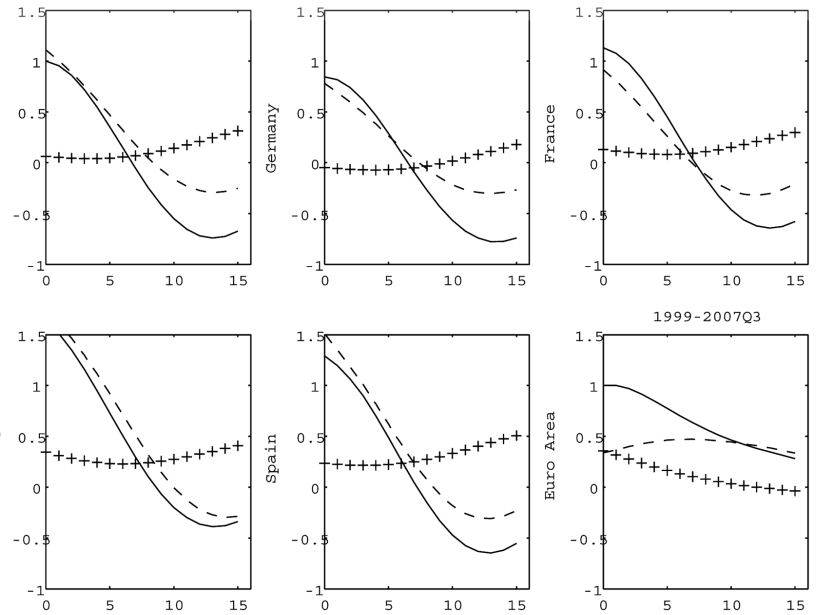
. *Has the Transmission Changed with the EMU?*

mplies that we should be extremely cautious in interpreting the results. We nevertheless trust that the estimates provide an indication on the irection of evolution of the effects of monetary policy with respect o the full-sample estimates.

Several results are worth emphasizing for the post-1999 period, again n the face of a 100 basis point increase in the short-term interest rate. irst, the short-term interest rate responses are indistinguishable for all ountries, given that they refer to the same currency. Second, the rise in ond yields in the EMU period is almost half of the one estimated for he entire sample, and the large differences across countries that were bservable prior to the EMU vanish entirely. The EA effective exchange ate appreciates considerably more than it did over the full sample. One eason for this is that real exchange rates uniformly appreciate in EA ountries, including Italy and Spain.[19]

Given the relatively small change in bond yields, measures of economic activity such as real GDP, consumption, and investment fall much ss, if at all, in the EMU period. As a result, employment falls much less, nd the unemployment rate's increase is sensibly smaller.

Altogether, it appears that a major characteristic of the new monetary olicy regime is the lack of response of long-term interest rates to surise increases in the short-term interest rate.[20] We illustrate this evolution by comparing in figure 2 the response of the long-term interest rate dashed lines) to the response of an artificial long-term interest rate excluding a term premium (crosses). The latter is obtained by appealing o the expectations hypothesis and is computed as the average response f the short-term interest rate over the subsequent 28 quarters, that is, a heoretical bond of 7-year maturity. A striking difference between the ull sample and the post-1999 regime is that, since the launch of the uro, the response of long-term interest rates displays a smaller term remium (i.e., a smaller difference between the market long-term rate nd the artificial rate). The responses of these interest rates are represented n the lower right plot of figure 2 for the EA, but they are almost identical or all individual countries in the post-1999 period. Moreover, over the enre sample, the term premium gap is the largest in Italy and in Spain, hich suggests that prior to the launch of the euro, the premium for the sk of devaluation or depreciation of the peseta and the lira increased



g. 2. Responses of short- and long-term interest rates to monetary shock in EA. First ve subplots represent responses of interest rates in 1998–2007 sample. Bottom-right plot ntains responses for EMU period. The solid line is the short-term interest rate, the shed line is the bond yield, and the crossed line is the theoretical bond yield without rm premium, based on the expectations hypothesis.

onetary policy regime characterized by more stable long-term interest tes and a sharper response of the EA-wide real exchange rate to mone- ry policy shocks.

Finally, the responses of several variables (some not reported) remain eterogeneous across countries in the EMU period. To name a few, the sponses of M1 are twice as negative in Spain and Belgium as in France, ermany, and Italy. M3 increases in all countries, though to a different xtent. Relatively larger responses of German exports and investment arry through to a larger GDP response than in other EA countries. Public onsumption responses range from positive in Belgium and Italy—the vo countries with the largest stock of government debt—to sharply neg- tive in the Netherlands. We also note some differences in labor market ynamics, aspects analyzed in depth in McCallum and Smets (2007).

hecks with respect to the econometric specification of the FAVAR. In par-
cular, we estimated the above impulse response functions with models
hat admit additional lags, additional latent factors, and quarter-on-quarter
rowth rates, and we consider shocks to the German interest rate instead
f the EA average interest rate.

Most of the results described above are robust. In particular, the larger
esponse of the Italian and Spanish interest rates and of their consump-
on are common outcomes of all these alternative specifications when
stimated over the full sample. Interestingly, Italy and Spain also stand
ut in response to an unexpected oil price increase, with Italian and
panish bond yields increasing more than in the other countries of the
A and consumption falling more (see the appendix of the working pa-
er version). This provides further evidence that bond markets and
redibility issues may contribute to the different responses of European
conomies to various shocks prior to the EMU.

In all specifications considered, we observe a smaller response of con-
umption after 1999 than in the full-sample estimates, following a
onetary tightening. However, the specification with quarter-on-quarter
rowth rates and several lags shows that, because of a large response of
xports, GDP declines as much in the post-1999 period as in the full sam-
le. These impulse response functions, however, are much less precisely
stimated than in our benchmark specification.

In the case in which the monetary policy shock is defined in terms of
he German short-term interest rate, nearly all the results reported in
gure 1 carry through. As mentioned above, however, the price puzzle
or German CPI is very much attenuated. This reflects that the identi-
cation of area-wide monetary shocks in the period prior to the euro is
ifficult. However, except for the response of German prices, nearly all
ther impulse responses are strikingly similar for a German or an area-
ide monetary policy shock.

V. Explaining the Evolution of the Transmission Mechanism: The Role of Monetary Regimes and Interest Rate Parity

s discussed in the previous section, the empirical characterization of
he transmission of monetary policy in the EA displays a rich picture. In

ountries. Such reductions in activity are offset by a relatively strong nprovement in net exports, thereby resulting in a moderate contrac- on of real GDP. In the EMU period, however, a similar increase in he EA interest rate results in a much more homogeneous response of ndividual EA countries and a quantitatively smaller reduction in eco- omic activity measures.

While the European economy has changed in many dimensions since he monetary union, we now attempt to determine to what extent the onetary regime in place can explain the differences in the transmission f monetary policy both across countries and over time. To do so, we use n open-economy DSGE monetary model along the lines of Clarida, alí, and Gertler (2002), Obstfeld and Rogoff (2002, 2005), Altissimo, enigno, and Rodriguez-Palenzuela (2004), Corsetti and Pesenti (2005), enigno and Benigno (2006), Ferrero et al. (forthcoming), and others.[21] he specific variant considered here builds on the work of Ferrero t al. This framework, while stylized, is sufficiently rich to generate a ontrivial effect of monetary policy variables such as output, consump- on, net exports, and inflation measures. It also allows for different onsumption responses across regions and a switching of expendi- ures in consumption and net exports in response to real exchange rate ovements.

We proceed by presenting the model. The model is explained in etail in Ferrero et al. (forthcoming), so we merely summarize it here, mphasizing the changes relative to their model. We next discuss he calibration of the model parameters, including those charac- erizing monetary policy. Finally, we analyze the model's implica- ons, attempting to provide an explanation for the stylized facts just escribed.

. *A Stylized Two-Country Model*

he model involves two large countries, Home (H) and Foreign (F), of qual size. Each country is populated by a representative household hat consumes tradable and nontradable goods and contains a contin- uum of workers who supply labor to intermediate-goods firms. Each f these firms hires one worker and produces either tradable or non-

ampa and Goldberg 2006), we assume as in Monacelli (2005) that nonopolistically competitive importers of foreign tradable goods resell hem to residents at prices set in domestic currency in a staggered fash-on.[22] In order to account for different consumption behavior across ountries, we assume incomplete financial markets across countries even though the household provides perfect insurance within each ountry) by assuming that a single bond is traded internationally. s in Ferrero et al. (forthcoming), one simplification is that we treat s nondurable consumption all domestic interest rate sensitive ex-enditures, including what is commonly labeled as investment. How-ver, as mentioned in Woodford (2003, chap. 5), to the extent that we re not interested in distinguishing consumption and investment, this hould not affect importantly the model's predictions for the other ariables.[23]

We will consider two monetary regimes. The pre-EMU regime is haracterized by distinct central banks in each country, each setting hort-term interest rates according to a generalized Taylor rule that ay include responses to exchange rate fluctuations. Area-wide variables re obtained by aggregating the relevant variables across the two coun-ies. In the post-EMU regime, instead, a supranational authority—the CB—is assumed to set an EA-wide interest rate according to a general-ed Taylor rule involving area-wide variables.

In order for the model to be consistent with the identifying assump-ons made in our empirical FAVAR to identify the monetary policy hocks, we assume in contrast to Ferrero et al. (forthcoming) but simi-rly to Rotemberg and Woodford (1997) and Christiano et al. (2005) hat the households' aggregate consumption decisions and all firms' ricing decisions are made prior to the realization of exogenous shocks, o that prices and consumption do not respond contemporaneously to he monetary shock. In addition, we allow households to form habits in onsumption and the firms that do not reoptimize their prices to index hem to past inflation. Such deviations from Ferrero et al.'s model al-ow the model to generate responses of consumption and inflation to hocks that are more in line with the FAVAR estimates.

As a last departure from Ferrero et al., we allow for a wedge in the ncovered interest rate parity (UIP) condition. This wedge, assumed to

e.g., Froot and Thaler 1990; Bekaert and Hodrick 1993; Engel 1996; Mark nd Wu 1998; Rossi 2007) or conditionally on monetary policy shocks Eichenbaum and Evans 1995; Scholl and Uhlig 2008). While Bekaert, Wei, and Xing (2007) find smaller departures from the UIP than reported reviously, when adjusting for small-sample bias, they find evidence f a time-varying risk premium displaying a highly persistent compo- ent in expected exchange rate changes. As discussed below, such wedge will prove to be important in explaining the differential re- ponses of consumption and investment across countries in the pre- MU period.

We now describe the environment, following closely the model of errero et al.

Households

We assume that in each country, the representative household maximizes lifetime expected utility of the form

$$E_{t-1}\left\{\sum_{s=0}^{\infty}\theta_{t+s-1}\left\{\frac{(C_{t+s}-\omega C_{t+s-1})^{1-\sigma}}{1-\sigma}-\left[\int_0^{\gamma}\frac{L_{Ht+s}(f)^{1+\varphi}}{1+\varphi}df\right.\right.\right.$$

$$\left.\left.\left.+\int_{\gamma}^{1}\frac{L_{Nt+s}(f)^{1+\varphi}}{1+\varphi}df\right]\right\}\right\}, \tag{4}$$

where E_{t-1} is the expectation operator, conditional on the information up o the end of period $t-1$; C_t denotes aggregate consumption; $\omega \in (0,1]$ is he degree of internal habit persistence; $\sigma^{-1} > 0$ would correspond to the lasticity of intertemporal substitution in the absence of habit forma- on; φ is the inverse of the Frisch elasticity of labor supply; and $L_{kt}(f)$ epresents hours worked by worker $f \in [0,1]$ in an intermediate-goods rm, in sector k, that is, either the home tradable sector H (with measure γ) r the domestic nontradable sector N (with measure $1-\gamma$). As in Ferrero t al. (forthcoming), the discount factor θ_t evolves according to $\theta_t = \beta_t\theta_{t-1}$ nd $\beta_t \equiv e^{\xi_t}/[1+\psi(\log \bar{C}_t - \bar{\vartheta})]$, where $\bar{C}_t$ corresponds to the household's onsumption level but is treated by the household as exogenous, and ξ_t is preference shock.[24]

The consumption index C_t is an aggregate of tradable C_{Tt} and non-

with $\gamma \in [0,1]$ representing the share of tradable goods. The consumption of tradable goods combines in turn home-produced goods C_{Ht} and foreign-produced goods C_{Ft} as follows:

$$C_{Tt} \equiv [\alpha^{1/\eta}(C_{Ht})^{(\eta-1)/\eta} + (1-\alpha)^{1/\eta}(C_{Ft})^{(\eta-1)/\eta}]^{\eta/(\eta-1)}.$$

The coefficient $\alpha \in (0.5,1]$ denotes home bias in tradables, and η is the elasticity of substitution among domestically produced and imported tradables. The home CPI, which minimizes the cost of consumer expenditures, is given by

$$P_t = P_{Tt}^{\gamma} P_{Nt}^{1-\gamma},$$

where the price of tradables is given by $P_{Tt} = [\alpha P_{Ht}^{1-\eta} + (1-\alpha)P_{Ft}^{1-\eta}]^{1/(1-\eta)}$. In the foreign country, we assume symmetric preferences, consumption aggregates, and price indices, which we denote by starred (*) variables and coefficients.[25]

Optimal behavior on the part of each household requires first an optimal allocation of consumption spending across differentiated goods. While we assume that households choose their level of total consumption on the basis of information available at date $t-1$, we let them choose the allocation of their consumption basket after the contemporaneous shocks have been realized. The optimal allocation of (domestically and foreign-produced) tradable goods as well as nontradable goods then takes the usual form:

$$C_{Tt} = \gamma\left(\frac{P_{Tt}}{P_t}\right)^{-1} C_t, \quad C_{Nt} = (1-\gamma)\left(\frac{P_{Nt}}{P_t}\right)^{-1} C_t, \tag{5}$$

$$C_{Ht} = \alpha\left(\frac{P_{Ht}}{P_{Tt}}\right)^{-\eta} C_{Tt}, \quad C_{Ft} = (1-\alpha)\left(\frac{P_{Ft}}{P_{Tt}}\right)^{-\eta} C_{Tt}. \tag{6}$$

As in Ferrero et al. (forthcoming), we assume that there is a single internationally traded one-period bond. We denote by B_t the nominal holdings at the beginning of period $t+1$, denominated in units of the home currency. The household's budget constraint in the home country is then given by

$$P_t C_t + B_t = I_{t-1} B_{t-1} + \int_0^{\gamma} W_{Ht}(f) L_{Ht}(f) df$$

$$\int_{\gamma}^{1} W_{Ft}(f) L_{Ft}(f) df + \Phi_t \tag{7}$$

ector k, and Υ_t combines aggregate dividends, lump-sum taxes, and ansfers. Maximizing the utility function (4) subject to (7) yields the fol- owing optimal choice of expenditures:

$$E_{t-1}\{\Lambda_t P_t\} = E_{t-1}\{(C_t - \omega C_{t-1})^{-\sigma} - \omega\beta_t(C_{t+1} - \omega C_t)^{-\sigma}\}, \quad (8)$$

here Λ_t is the household's marginal utility of additional nominal in- ome at date t. This expression makes clear that the plan for aggregate onsumption at date t is made on the basis of information available at ate $t-1$. The marginal utilities of income must in turn satisfy the Euler quation

$$1 = E_t\left\{I_t \frac{\beta_t \Lambda_{t+1}}{\Lambda_t}\right\}. \quad (9)$$

urthermore, the optimal choice of labor supply equalizes the real vage with the marginal rate of substitution between consumption nd leisure.

The representative household in the foreign country is very similar. ne difference, however, between the two countries is that the foreign ond is not traded internationally. The foreign household's budget con- raint, expressed in units of the foreign currency, is then

$$P_t^* C_t^* + D_t^* + \frac{B_t^*}{\mathcal{E}_t} = I_{t-1}^* D_{t-1}^* + \frac{I_{t-1} B_{t-1}^*}{\mathcal{E}_t e^{\mu_{t-1}}} + \int_0^{\gamma} W_{Ft}^*(f) L_{Ft}^*(f) df + \int_{\gamma}^{1} W_{Nt}^*(f) L_{Nt}^*(f) df + \Upsilon_t^*, \quad (10)$$

here the labor income indicates that foreign workers and firms operate in ther the foreign tradable sector or the nontradable sector; D_t^* represents e foreign household's holdings of the foreign debt; B_t^* denotes the for- gn household's holdings of the domestic bond, issued in the home cur- ency; and $\mathcal{E}_t$ is the nominal exchange rate, that is, the amount of home urrency needed in exchange for a unit of foreign currency. In contrast Ferrero et al. (forthcoming) but as in McCallum and Nelson (2000) or ustiniano and Preston (2006), we introduce an exogenous term $e^{\mu_{t-1}}$ that an be interpreted as a risk premium shock or a bias in the foreign house- old's expectation of the period t revenue from holding home bonds.

hat foreign citizens may hold bonds of both countries, they must be indifferent between holding home and foreign bonds. This results in he following UIP condition:

$$E_t\left\{I_t \frac{\mathcal{E}_t}{\mathcal{E}_{t+1}e^{\mu_t}} \frac{\beta_t^* \Lambda_{t+1}^*}{\Lambda_t^*}\right\} = E_t\left\{I_t^* \frac{\beta_t^* \Lambda_{t+1}^*}{\Lambda_t^*}\right\}. \tag{11}$$

Firms

We have three types of firms: final-goods firms, intermediate-goods rms, and importing retailers.

Final-goods firms. In each sector H and N, final-goods firms, which are cting on a competitive market, combine intermediate goods to produce utput

$$Y_{Ht} \equiv \left[\gamma^{-(1/\theta)} \int_0^{\gamma} Y_{Ht}(f)^{(\theta-1)/\theta} df\right]^{\theta/(\theta-1)},$$

$$Y_{Nt} \equiv \left[(1-\gamma)^{-(1/\theta)} \int_{\gamma}^{1} Y_{Nt}(f)^{(\theta-1)/\theta} df\right]^{\theta/(\theta-1)},$$

where $\theta > 1$ is the elasticity of substitution among intermediate goods. ost minimization for the final-goods firms implies the following de- and functions for intermediate-goods firms:

$$Y_{Ht}(f) = \gamma^{-1}\left[\frac{P_{Ht}(f)}{P_{Ht}}\right]^{-\theta} Y_{Ht},$$

$$Y_{Nt}(f) = (1-\gamma)^{-1}\left[\frac{P_{Nt}(f)}{P_{Nt}}\right]^{-\theta} Y_{Nt}, \tag{12}$$

where the price indices P_{Ht} and P_{Nt} aggregate underlying prices $P_{kt}(f)$. Each intermediate firm f in sector $k = H, N$ produces output $Y_{kt}(f)$ by iring labor $L_{kt}(f)$ and using the production function

$$Y_{kt}(f) = A_t L_{kt}(f),$$

Intermediate firms. Intermediate firms are assumed to set prices in a staggered manner. A fraction $1-\xi$ of firms (chosen independently of the history of price changes) can choose a new price in each period. Our informational assumptions imply that the firms that get to reset their prices must do so using information available at period $t-1$. In addition, we assume that if a price is not reoptimized, it is indexed to lagged inflation in sector $k = H, N$ according to the rule

$$P_{kt}(f) = P_{k,t-1}(f)\left(\frac{P_{k,t-1}}{P_{k,t-2}}\right)^{\delta} \tag{13}$$

for some $\delta \in [0,1]$. Given that the problem is the same for all firms of sector k that reset their price at date t, they all choose an optimal price $P^o_{k,t}$ that maximizes

$$E_{t-1}\left\{\sum_{s=0}^{\infty}\xi^s\Lambda_{t,t+s}\left[P^o_{kt}\left(\frac{P_{k,t+s-1}}{P_{k,t-1}}\right)^{\delta} - MC_{k,t+s}(f)\right]Y_{k,t+s}(f)\right\}$$

subject to the demand for their good (12). In the previous expression, $\Lambda_{t,t+s} = \beta_{t,t+s}\Lambda_{t+s}/\Lambda_t$ is the stochastic discount factor between periods t and $t+s$, $\beta_{t,t+s} = \Pi_{j=0}^{s-1}\beta_{t+j}$ for $s \geq 1$, and $\beta_{t,t} = 1$.

The price index then satisfies

$$P_{kt} = \left\{(1-\xi)(P^o_{kt})^{1-\theta} + \xi\left[P_{k,t-1}\left(\frac{P_{k,t-1}}{P_{k,t-2}}\right)^{\delta}\right]^{1-\theta}\right\}^{1/(1-\theta)}.$$

Importing retailers. To model the imperfect pass-through found in the data, we assume that monopolistically competitive retailers import foreign tradable goods and sell them to domestic consumers, as in Monacelli (2005). These retailers also set their prices in a staggered fashion so that the law of one price does not hold at the consumer level. As for the intermediate firms, a fraction $1-\tilde{\xi}$ of retailers choose a new price in each period on the basis of information available at period $t-1$. Again, if a price is not reoptimized, it is indexed to lagged inflation in that sector, according to the rule (13). Since the problem is identical for retailers that reset their price at date t, they all choose an optimal price $P^o_{F,t}$ in domestic cur-

ubject to the demand for the imported good (6). In the above expres-on, $P^*_{F,t}$ denotes the price of foreign tradable goods in a foreign cur-ency. The price index of imported goods in the domestic currency atisfies

$$P_{Ft} = (1 - \tilde{\xi})(P^o_{Ft}) + \tilde{\xi} P_{F,t-1} \left(\frac{P_{F,t-1}}{P_{F,t-2}} \right)^{\tilde{\delta}}.$$

Monetary Policy

We consider two distinct monetary regimes, one referring to the pre-MU period, in which each national central bank sets its own interest rate ccording to a generalized forward-looking Taylor rule, and one refer-ng to the monetary union, in which a supranational central bank sets ommon short-term interest rates.

More specifically, in the pre-EMU regime, we assume that the home ational central bank sets its short-term riskless interest rate according o

$$i_t = \rho i_{t-1} + (1 - \rho)(\phi_\pi E_t \bar{\pi}_{t+h} + \phi_y y_t + \phi_i i^*_t + \phi_e \Delta e_t) + \varepsilon_t, \quad (14)$$

where $i_t \equiv \log(I_t/I)$ corresponds to the deviations of the interest rate om its steady-state value, $\bar{\pi}_t \equiv \log(P_t/P_{t-4})$ denotes deviations of yoy PI inflation around the steady state (assumed to be zero), y_t represents ercent deviations of output from trend, $\Delta e_t = \log(\mathcal{E}_t/\mathcal{E}_{t-1})$ denotes ercent nominal depreciation of the home currency, and the indepen-ently and identically distributed shock ε_t measures unexpected inter-st rate disturbances. The foreign central bank follows a similar rule:

$$i^*_t = \rho^* i^*_{t-1} + (1 - \rho^*)(\phi^*_\pi E_t \bar{\pi}^*_{t+h^*} + \phi^*_y y^*_t + \phi^*_i i_t + \phi^*_e \Delta e_t) + \varepsilon^*_t, \quad (15)$$

where, again, the asterisks refer to foreign variables or coefficients. ote that we allow for cross-country interactions since the national entral banks may respond to fluctuations in the exchange rate or to he other country's interest rate. Clarida et al. (1998) and Angeloni and edola (1999) argue that such rules provide a good characterization of

eriods. We assume that the common central bank—corresponding to he ECB—sets interest rates according to the interest rate rule

$$i_t^{ea} = \rho^{ea} i_{t-1}^{ea} + (1-\rho^{ea})(\phi_\pi^{ea} E_t \bar{\pi}_{t+h}^{ea} + \phi_y^{ea} y_t^{ea}) + \varepsilon_t^{ea}, \tag{16}$$

where area-wide inflation and output are defined as $\bar{\pi}_t^{ea} = (\bar{\pi}_t + \bar{\pi}_t^*)/2$ nd $y_t^{ea} = (y_t + y_t^*)/2$.

Equilibrium Characterization

o close the model, we use equilibrium conditions stating that the supply of tradable and nontradable goods must be equal to the respective emands in each country and that international financial markets clear. o characterize the response of various variables to monetary shocks, we olve a log-linear approximation to the model's equilibrium conditions round a deterministic state, using standard techniques. We thus implicitly assume that the shocks are small enough for the approximation to be alid. In the steady state, both economies are symmetric; the trade balance nd foreign debt are equal to zero; output in each sector grows at the constant trend productivity growth rate g; the relative prices of all goods, including the real exchange rate $Q_t \equiv \mathcal{E}_t P_t^*/P_t$, are equal to one; inflation is qual to zero; and the real interest rate is equal to $(1+g)/\beta$, where β is the eady-state value of β_t.

The log-linearized equilibrium conditions are described in the appendix of the working paper version.

Model Calibration

We calibrate the model's parameters in order to provide its quantitative redictions and to determine whether we can replicate at least some of he stylized facts mentioned above. In particular, we focus our attention n changes in responses of key macroeconomic variables between the re-EMU and EMU periods. We also focus on the difference in responses cross countries in the pre-EMU period, especially the differences between Italy and Spain on the one hand and Germany along with other A countries on the other hand. We assume that Home (H) stands for aly or Spain and Foreign (F) stands for Germany along with the other

acrifices somewhat the model's ability to replicate the empirical re-
ponses, we did check that the model's predictions are not too sensitive
o the chosen parameter values. However, as we will see below, coeffi-
ents of the policy rules do play an important role in the shape of the
esponses to various shocks.

Structural Parameters

s mentioned, most structural parameters are taken from Ferrero et al.
orthcoming) and are roughly in line with values chosen in other stud-
s (e.g., Obstfeld and Rogoff 2005) and with some microeconomic data.
e set the same values for both countries. The steady-state growth rate
f the economy g is set to 0.5%, so that annual growth is 2%. The steady-
ate discount factor β is set to 0.99. The parameters describing the evolu-
on of the discount factor $\vartheta = -1,000$ and $\psi = 7.2361 \cdot 10^{-6}$ are chosen
o that fluctuations in β_t have no noticeable implications on the econ-
my dynamics.[26] The Frisch elasticity of labor supply is $\varphi^{-1} = 0.5$. The
asticity of substitution among intermediate goods $\theta = 11$ results in a
eady-state markup of 10% in the tradable and nontradable sectors. We
et the probability that intermediate-goods firms and importing retailers
o not reoptimize their price to $\xi = \tilde{\xi} = 0.66$, corresponding to a mean
uration between price reoptimizations of 3 quarters. Smets and Wouters
2002) find evidence that import prices display a degree of price sticki-
ess similar to that of domestic prices on the basis of estimated responses
o monetary shocks in the EA. For the parameters that determine the
penness of the economies, we set the share of tradables in the con-
umption basket γ to 0.25, the preference share for home tradables $\alpha =$
7 (it would be 0.5 in the absence of home bias), and the elasticity of
ubstitution between home and foreign tradables is $\eta = 2$, as in Ferrero
al.

Ferrero et al. assume a log utility function of consumption and no
abit persistence or inflation indexing. However, this yields sharp re-
ponses in inflation and consumption to monetary shocks, in contrast
o the empirical evidence. To generate more realistic hump-shaped re-
ponses of consumption expenditures and output of the model econ-
my, we assume some degree of habit persistence ω.[27] We calibrate this

Policy Rule Coefficients

We calibrate the policy rule coefficients for the home and foreign countries in the pre-EMU period using estimates of Angeloni and Dedola (1999, table 9b). These authors estimate interest rate rules of the form (14)–(15) jointly for Italy and Germany, for the period 1988–97, which covers nearly entirely our pre-EMU sample. Their preferred specification involves horizons on inflation expectations of $h = h^* = 0$, so that the central banks set interest rates in response to inflation that has occurred over the past year. As the estimates are obtained using monthly data, we convert them for application to quarterly data.[28] We thus have $\rho = 0.79$, $\phi_\pi = 1.22$, $\phi_y = 0.30$, and $\phi_i = 0.41$ for Italy[29] and $\rho^* = 0.82$, $\phi^*_\pi = 1.41$, $\phi^*_y = 0.30$, and $\phi^*_i = 0$ for Germany. Angeloni and Dedola do not include a bilateral deutsche mark/lira exchange rate in their policy rules, but they include the dollar/deutsche mark exchange rate. Since we abstract from the world outside of the EA in the model, we assume that German monetary policy does not respond to the exchange rate ($\phi^*_e = 0$), whereas the annualized Italian interest rate responds with a short-run coefficient of 0.4 to the exchange rate depreciation. This is meant to capture the fact that the Italian central bank was required to maintain its exchange rate within narrow bands, as long as it took part in the exchange rate mechanism. This results in a long-run coefficient $\phi_e = 5$.

For the post-EMU period, we estimate an interest rate rule of the form (16) on EA data, using generalized method of moments, similarly to Clarida et al. (1998). We use as instruments the current value of inflation and detrended output as well as three latent factors extracted from the EA indicators. Our preferred horizon is $h = 2$. As the estimated coefficient on the lagged interest rate is relatively high, $\rho^{ea} = 0.93$, the implied long-run responses to expected inflation and output fluctuations are also quite strong: $\phi^{ea}_\pi = 13.03$ and $\phi^{ea}_y = 8.01$.[30] Nonetheless, we verify that our conclusions remain robust to smaller values of these coefficients.

Wedge in Uncovered Interest Rate Parity

We assume that μ_t follows an AR(1) process that is allowed to respond to monetary shocks

$$\mu_t = \rho_\mu \mu_{t-1} + \nu \varepsilon_t^* + \varepsilon_{\mu t},$$

where ε_t^* are foreign monetary policy shocks and $\varepsilon_{\mu t}$ denotes other possible shocks to that wedge. By allowing μ_t to respond to monetary shocks, we hope to capture in an arguably reduced form the effect of monetary shocks on the risk premium emphasized by Scholl and Uhlig (2008). We assume that this wedge is very persistent, setting $\rho_\mu = 0.98$, and will consider different values of the parameter ν.

4. *The Model's Quantitative Predictions: Explaining the Changes in the Monetary Transmission Mechanism*

Having calibrated the model, we can now determine whether it can replicate the stylized facts mentioned above, namely, the cross-country differences in responses as well as their changes with the introduction of the euro that we report in figures 1*a*–1*c*.

Pre-EMU Cross-Country Differences

Figures 3*a*–3*d* indicate the responses of key variables to an unexpected interest rate increase of 100 basis points in the foreign economy—which stands for Germany—in the case in which both economies set their interest rates according to the estimated policy rules (14) and (15). This is meant to replicate the effects of a monetary policy tightening in the pre-EMU period, reported in figures 1*a*–1*c*.

Figure 3*a* shows the responses of the home economy (i.e., Italy or Spain, solid lines) and the foreign economy (i.e., Germany, dashed lines) in the absence of a wedge in the UIP condition ($\nu = 0$, so $\mu_t = 0$). The unexpected increase in the foreign short-term rate is associated with a rise in the long-term rate and a drop in output, consumption, and inflation. As the domestic currency depreciates more than prices adjust, the domestic real exchange rate (q_t) also depreciates, and home terms of trade (TOT, measuring foreign prices relative to domestic prices, in domestic currency) increase. This stimulates an increase in net exports of home goods.

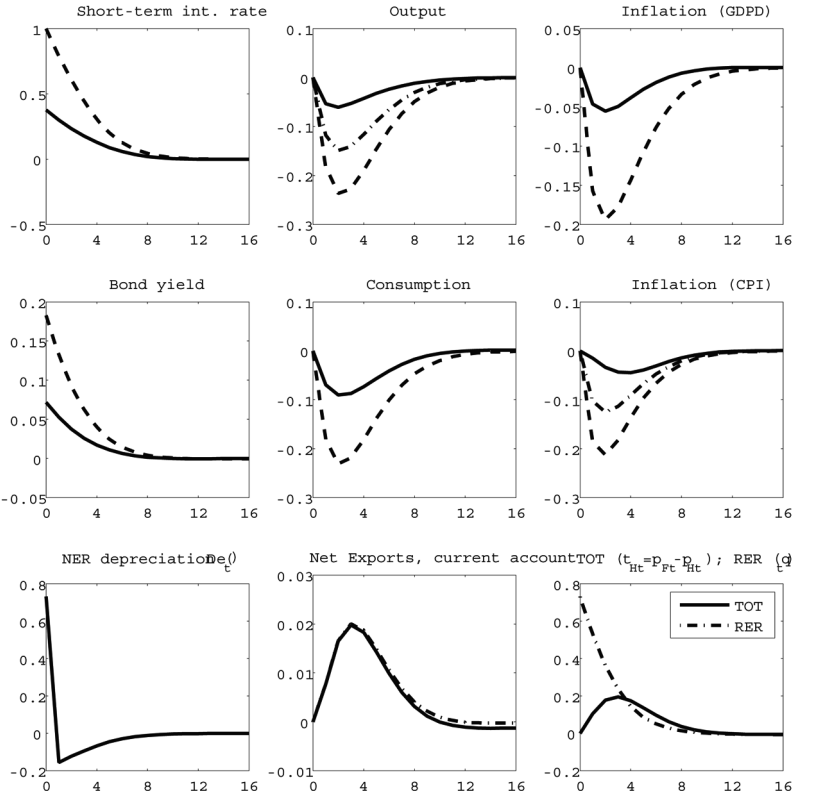


Fig. 3*a*. Model-based responses to a 100 basis point monetary tightening in a foreign country (pre-EMU, $\nu = 0$).

The response of home interest rates just described, however, is at odds with the interest rate responses that we had documented for countries such as Spain and Italy in figure 1*a*. In fact, in pre-EMU data, these short- and long-term rates increased significantly more than those estimated for Germany and other countries. They were also associated with sharp contractions in consumption and employment in those countries. Instead, the model-based responses display a milder response of the home variables. One might think that by letting the home country's central bank respond more to exchange rate fluctuations (i.e., a larger ϕ_e), we

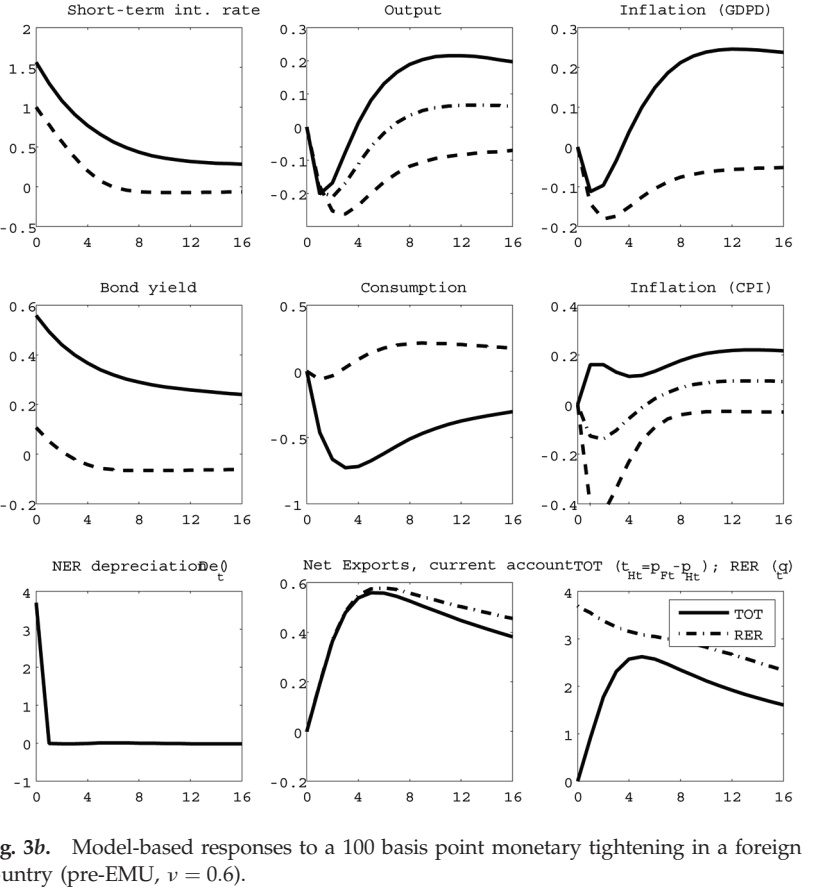


Fig. 3*b*. Model-based responses to a 100 basis point monetary tightening in a foreign country (pre-EMU, $\nu = 0.6$).

impulse responses, which are identical in both countries, thus correspond to those of a single closed economy.[31] In addition, changes in structural parameters do not generically modify the picture presented. The basic version of the model cannot replicate the transmission of monetary policy observed in low-credibility regimes since long-term rates are tightly tied to expected future riskless short-term rates. One key parameter, however, that allows us to deviate from the standard case and seems to explain the stylized facts reported in figures 1*a*–1*c* is ν. Figure 3*b* reports the model-based responses of the same variables in

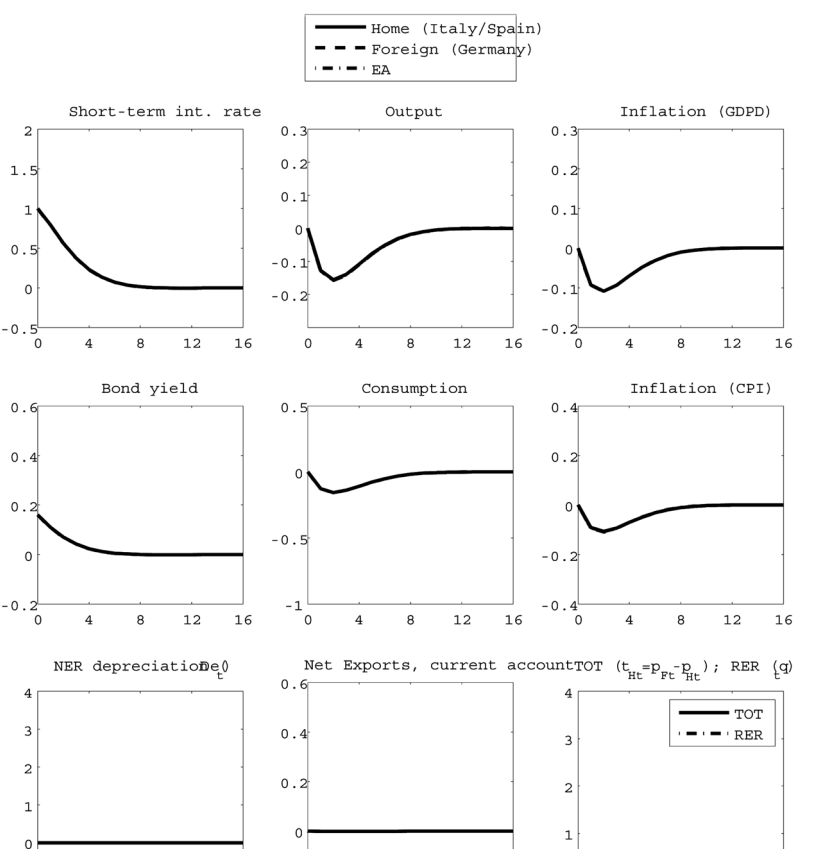


Fig. 3*c*. Model-based responses to a 100 basis point monetary tightening in a common area (monetary union case, $\nu = 0.6$).

require a higher return on domestic (internationally traded) bonds than they do on foreign securities, even after accounting for the rational expectation of nominal exchange rate changes.

Such an exchange rate risk premium appears important to explain the stylized facts reported above. In fact, in figure 3*b*, not only do short- and long-term rates respond more strongly at home than in the foreign country, but these interest rate responses also generate a larger drop in

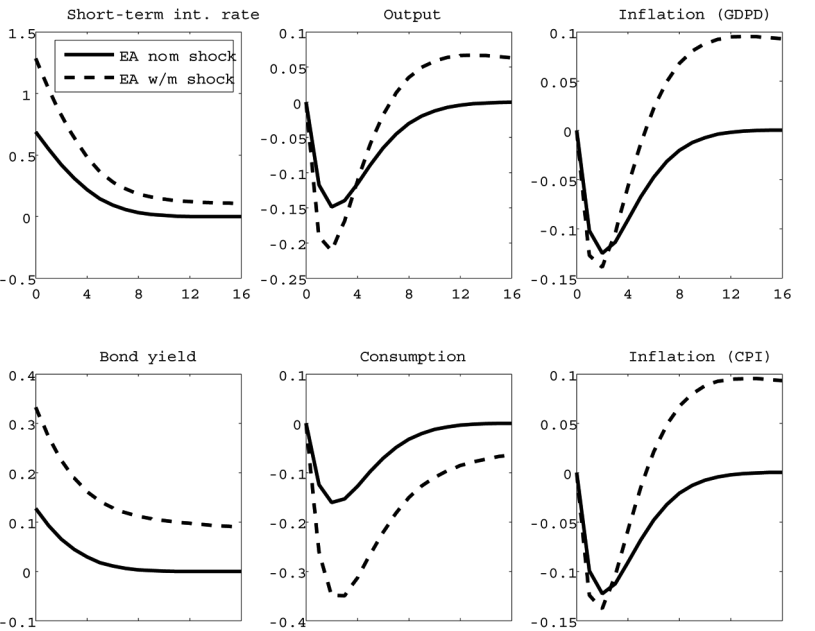
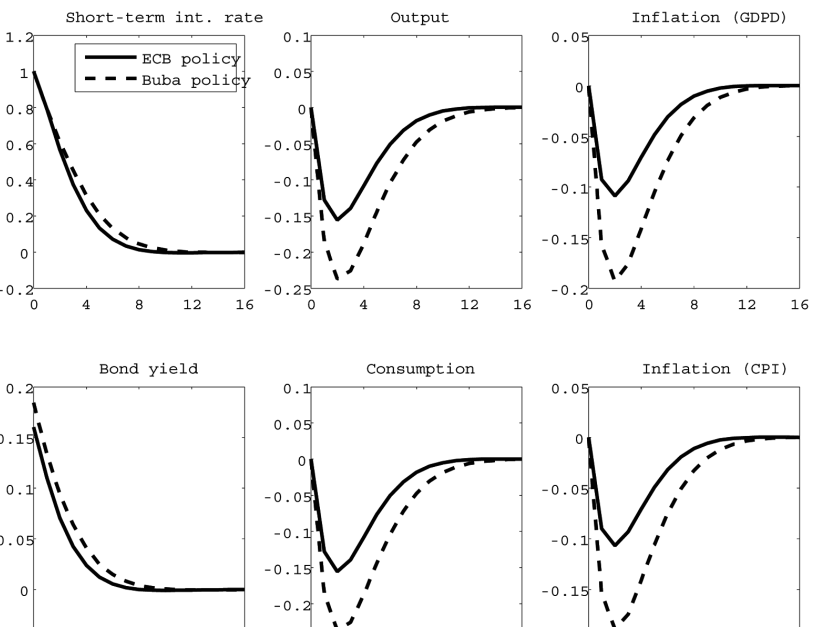


Fig. 3*d*. Model-based responses of aggregate EA variables to a 100 basis point monetary tightening in a foreign economy (solid lines: $\nu = 0$; dashed lines: $\nu = 0.6$).

Interestingly, prices aggregated for both regions (dashed-dotted line) can also increase following the monetary tightening, to the extent that inflation in the depreciating country more than offsets the inflation reduction in the other region. This can explain an apparent "price puzzle" in the home country or in the area as a whole in response to monetary tightening.

The exercise just performed thus suggests that conditional on EA-wide (or German) monetary shocks, changes in the risk premium on Italian and Spanish securities may provide an important explanation for the large observed responses in bond yields and the fact that consumption and investment used to fall considerably more in those countries than in the rest of the EA.

Monetary Union and Changes in the Monetary Transmission Mechanism



g. 3*e*. Model-based responses of EA variables to a 100 basis point monetary tightening the EA under alternative policy rules (solid lines: ECB policy rule; dashed lines: undesbank policy rule).

onetary union, when monetary policy is conducted according to the stimated rule (16). Since both countries are symmetric in the calibration, ey both respond identically to the EA interest rate shock.

Comparing figures 3*a*–3*b* on one hand with figure 3*c* on the other and reveals important differences in the responses to an unexpected crease in the interest rate set by the foreign central bank in the pre-MU regime and the common central bank in the EMU regime. The odel predicts that the home economy benefits in many respects from articipating to the monetary union in response to such a shock. In par-cular, when we remove exchange rate risks in the EMU regime, home short- and long-term) interest rates increase by less, as we observed in he empirical responses. As a result, home consumption falls by less nd output remains more stable.

shock in the foreign country. The figure reveals that the responses of short- and long-term interest rates, output, consumption, and inflation are all more muted when we remove the risk premium shock μ_t.

The EMU has not only contributed to smaller responses by removing exchange rate risks. The model predicts that a monetary policy that has more consistently aimed at stabilizing inflation and output in the EA, since the start of the EMU, should result in a smaller observed response of aggregate economic activity and inflation to monetary shocks, as observed in the data. To illustrate this, figure 3*e* shows the model-based responses of EA variables to a monetary tightening (of 100 basis points) assuming two different policy rules: the one estimated for the ECB (solid lines) and the one estimated previously (by Angeloni and Dedola [1999]) for the Bundesbank (dashed lines). The figure indicates that stronger responses by the ECB to inflation and output fluctuations, essentially stemming from a very inertial rule, have resulted in smaller responses of economic activity and inflation. It is important to stress, however, that the smaller response of output and inflation is not due to the fact that the economy is less sensitive to monetary policy. All elasticities describing the behavior of the private sector, such as the intertemporal elasticity of substitution and so on, are maintained constant in this experiment. It is only the stronger commitment to inflation and output stabilization that results in such an outcome.[32]

VI. Conclusion

In this paper, we have provided an empirical characterization of the monetary transmission mechanism in key European economies, exploiting the richness of the cross-country differences and the fact that a major change in monetary regime has occurred in 1999 with the adoption of the euro by 11 European countries. The combination of the cross-country heterogeneity and the changes over time provides a unique laboratory for the analysis of numerous macroeconomic indicators.

Focusing on six major European economies, we have argued that a large fraction of the fluctuations in these economic variables can be captured by a low-dimensional vector of common components. This finding is useful to the extent that it allows us to characterize the effects of

esponses can be rationalized by a two-country model, provided that
e allow for a disturbance in the uncovered interest rate parity condi-
on, which may be interpreted as a risk premium shock. In addition,
espite the short samples, we have detected preliminary evidence of
mportant changes in the transmission of monetary policy since the
art of the EMU.

We have argued that some of the changes since 1999 can be explained
y the change in the monetary regime. In particular, our model predicts
hat when an exchange rate risk has been removed through the mone-
ry union and there is a central bank that is more decisively focused on
flation and output stabilization, the impact of monetary disturbances
n measures of economic activity has been reduced, as observed in the
ata. While private consumption and investment in Italy and Spain ap-
ear to have been especially hard hit by German monetary policy dis-
urbances in the pre-EMU period, the new monetary regime has
ontributed to stabilizing them more effectively, in part because long-
erm interest rates have become much more effectively anchored in
uch countries since the start of the monetary union.

We have also found that the exchange rate channel has become rel-
tively more powerful in the monetary union period than in the pre-
ious decade and that national monetary aggregates appear much
ss driven by EA common shocks and show more heterogeneous re-
ponses to monetary policy shocks than most other macroeconomic
ariables.

ppendix

ata Description

he data were extracted from Haver Analytics, and their source is
ther the OECD Main Economic Indicators or OECD Quarterly Na-
onal Accounts databases. The sources for monetary aggregates are
he national central banks for their respective countries and the ECB
or the EA. The national accounts published in Haver are available start-
g at different dates: 1978 in France, 1981 in Italy, 1988 in the Nether-
nds, 1991 in Germany, 1995 in Spain and Belgium, and 1995 for some
eflators in the EA. Missing data were backdated using yoy growth

Endnotes

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1. At that date, the conversion rates of the national currencies of the Eurozone were fixed irrevocably, and a 3-year transition period started until the introduction of the euro banknotes and coins in January 2002. Since then other countries such as Greece, Slovenia, Malta, and Cyprus have adopted the euro.

2. We refer to the EMU as stage III of the European Monetary Union, which involves the launch of the euro in January 1999.

3. As long as a sufficient number of unobserved factors are included, the inclusion of oil price inflation as an observable factor should not affect our results. It does, however, allow us to identify oil price shocks and document their effects. We report such results in the appendix of the working paper version of this paper.

4. In fact, Stock and Watson (1999) refer to (2) as a dynamic factor model.

5. While alternative strategies to the estimation of factor models with a large set of indicators exist (see, among others, Forni et al. 2000; Kose, Otrok, and Whiteman 2003; Bernanke et al. 2005; Boivin and Giannoni 2006a; Doz, Giannone, and Reichlin 2007), the evidence suggests that they perform similarly in practice.

6. In contrast to the approach adopted here, Bernanke et al. do not impose the constraint that the observed factors are among the common components in the first step. They instead remove these observed factors from the space covered by the principal components by peforming a transformation of the principal components exploiting the different behavior of what they call "slow-moving" and "fast-moving" variables in the second step. Our approach here follows Boivin and Giannoni (forthcoming) and Boivin, Giannoni, and Mihov (forthcoming).

7. More specifically, we adopt the following procedure in the first step of the estimation. Starting from an initial estimate of F_t, denoted by $F_t^{(0)}$ and obtained as the first K principal components of X_t, we iterate through the following steps: (1) we regress X_t on $F_t^{(0)}$ and the observed factors $Y_t = [\pi_t^{oil}, R_t]'$ to obtain $\hat{\lambda}_Y^{(0)}$; (2) we compute $\tilde{X}_t^{(0)} = X_t - \hat{\lambda}_Y^{\prime(0)} Y_t$; (3) we estimate $F_t^{(1)}$ as the first K principal components of $\tilde{X}_t^{(0)}$; (4) we repeat steps (1)–(3) multiple times.

8. Note that this two-step approach implies the presence of "generated regressors" in the second step. According to the results of Bai (2003), the uncertainty in the factor estimates should be negligible when N is large relative to T. Still, the confidence intervals on the impulse response functions used below are based on a bootstrap procedure that accounts for the uncertainty in the factor estimation. As in Bernanke et al. (2005), the bootstrap procedure is such that (1) the factors can be resampled on the basis of the observation equation, and (2) conditional on the estimated factors, the VAR coefficients in the transition equation are bootstrapped as in Kilian (1998).

9. For a formal description of the link between the solution of a dynamic stochastic general equilibrium (DSGE) model in state space form and a VAR, see, e.g., Sims (2000) and Fernández-Villaverde et al. (2007). Boivin and Giannoni (2006a) establish formally the link between DSGE models and the FAVAR representation (1)–(2) in the context

11. Van Els et al. (2003) show that spillovers across countries tend to reinforce the effects monetary policy on output and on prices. See also Fagan and Morgan (2006).
12. Boivin and Giannoni (2006a) argue, e.g., that inflation is imperfectly measured by any ngle indicator and that it is important to use multiple indicators of it for proper inference.
13. Major steps in this process include the start of the EMS in 1979, the entrance of Spain d Portugal into the EMS in 1986, the post-reunification exchange rate crisis of 1992–93, and e announcement of the parities between national currencies and the euro in May 1997.
14. For instance, Kose et al. (2003) and Stock and Watson (2005) study the comovement output, consumption, and investment for a large panel of countries and for Group of countries, respectively. Giannone and Reichlin (2006) analyze the comovement of output ross EA countries. In addition, the ECB is carefully monitoring real and nominal hetero-eneity across countries (Benalal et al. 2006).
15. Camacho, Péréz-Quirós, and Saiz (2006) argue, however, that the EA business cycle rgely reflects the world business cycle.
16. Among the reponses not reported in the figures, the growth rate of M1 also falls, d the stock market drops by 10% on impact. Public consumption, however, remains nchanged for about a year and starts falling only after that.
17. The responses of the variables in Belgium and the Netherlands (not reported) are ery similar to those of the EA and countries such as Germany and France but different om the responses in Italy and Spain. The responses of Belgium and the Netherlands are vailable from the authors on request.
18. Recall that the variables in the FAVAR are expressed in yoy growth rates. The im-ulse response functions of yoy growth rates and (log) levels are identical for the first four uarters following the shock.
19. The real exchange rate response is larger for the EA than for each of the individual ountries since much of the trade of the individual countries is with other European onomies, whereas the EA real exchange rate measures appreciations and depreciations olely relative to countries outside of the EA.
20. This result is consistent with those of Ehrmann et al. (2007), who use daily interest tes to compare the responses of French, German, Italian, and Spanish long-term yields news in France, Germany, Italy, and Spain before and after 1999.
21. For a larger-scale model, see, e.g., Faruquee et al. (2007).
22. Corsetti and Dedola (2005) propose an alternative model of limited pass-through in hich distributing imported goods requires nontradables.
23. In fact, macroeconomic models that successfully explain the behavior of investment ten assume adjustment costs in investment (e.g., Basu and Kimball 2003; Christiano, chenbaum, and Evans 2005). As shown in Woodford (2003), such adjustment costs yield log-linearized Euler equation for investment that is very similar to the one for consump-on in the presence of internal habit formation. It follows that the intertemporal alloca-on of aggregate expenditures can be approximated by a similar Euler equation in which e degree of habit formation also serves as a proxy for investment adjustment costs. onetheless, in treating investment similarly to nondurable expenditures, we do abstract om the effects of investment on future production capacities.
24. This formulation of the discount factor incorporates—in the case in which the rep-sentative household stands for a continuum of households—the stimulative effect on dividual consumption of an increase in average consumption, as in Uzawa (1968). owever, as emphasized in Ferrero et al. (forthcoming), the parameter ψ is calibrated such a small value that this effect is negligible. It merely serves as a technical device guarantee a unique steady state in the case of incomplete financial markets across untries. One can alternatively obtain such a unique steady state by assuming a constant scount factor β but introducing a debt-elastic interest rate premium in the budget con-raints (7) and (10) below, as in Benigno (2001), Kollmann (2002), Schmitt-Grohe and ribe (2003), and Justiniano and Preston (2006).

26. As mentioned above, the assumption of a variable discount factor is merely a technical device yielding a unique steady state.
27. The degree of habit persistence also proxies for investment adjustment costs in the case in which consumption expenditures include also investment expenditures.
28. For the conversion, we assume that monthly values of (annualized) short-term interest rates are constant in a given quarter and equal to the corresponding (annualized) quarterly rate. In that case, the coefficient on the quarterly lagged interest rate is $\rho = \rho_m/(3 - 2\rho_m)$, where ρ_m is the policy coefficient on the monthly lagged interest rate. The long-run coefficients on inflation, output, and the foreign interest rate remain unchanged at the quarterly frequency.
29. Angeloni and Dedola's (1999) estimated policy rule for Spain is similar to that estimated for Italy.
30. While this representation of the policy rule appears very aggressive, it is important to realize that this is due to the large coefficient on the lagged interest rate. The policy rule may equivalently be written in terms of changes in the interest rate: $\Delta i_t^{ea} = 0.91E_t\bar{\pi}_{t+2}^{ea} + 0.56y_t^{ea} - 0.07i_{t-1}^{ea} + \varepsilon_t^{ea}$.
31. Recall that our calibration is such that apart from the policy rules, the home and foreign economies are perfectly symmetric.
32. Boivin and Giannoni (2006b) argue that a stronger commitment to inflation stabilization in U.S. monetary policy since the early 1980s can similarly explain the observed reduction in estimated responses of inflation and output in the U.S. economy in the post-1980 period.

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