OXFORD BULLETIN OF ECONOMICS AND STATISTICS, 76, 5 (2014) 0305–9049
doi: 10.1111/obes.12038

# Do Euro Area Countries Respond Asymmetrically to the Common Monetary Policy?*

Matteo Barigozzi†, Antonio M. Conti‡,¶ and Matteo Luciani§,¶

†*Department of Statistics, London School of Economics and Political Science, London, UK (e-mail: m.barigozzi@lse.ac.uk)*
‡*Economic Research Unit, Cagliari Branch, Bank of Italy, Cagliari, Italy (e-mail: antoniomaria.conti@bancaditalia.it)*
§*F.R.S.-FNRS, 5 Rue d'Egmont 1000, Brussels Belgium (e-mail: matteo.luciani@ulb.ac.be)*
¶*European Center for Advanced Research in Economics and Statistics (ECARES), Solvay Business School of Economics and Management, Université libre de Bruxelles, 50 av. F.D. Roosevelt, CP 114, B1050, Brussels, Belgium*

## Abstract

We investigate the possible existence of asymmetries among Euro Area countries reactions to the European Central Bank monetary policy. Our analysis is based on a Structural Dynamic Factor model estimated on a large panel of Euro Area quarterly variables. Although the introduction of the euro has changed the monetary transmission mechanism in the individual countries towards a more homogeneous response, we find that differences still remain between North and South Europe in terms of prices and unemployment. These results are the consequence of country-specific structures, rather than of European Central Bank policies.

## I. Introduction

Before the introduction of the common currency, every Euro Area (EA) member state's central bank had a different attitude towards the objectives of containing inflation and boosting economic growth (Clarida, Galí and Gertler, 1998; Mihov, 2001). After 1999, the European Central Bank (ECB) took over national central banks and imposed a common monetary policy. Nowadays, all EA countries are subject to this single policy, but are still

*A previous version of this article circulated under the title 'Measuring Euro Area Monetary Policy Transmission in a Structural Dynamic Factor Model'. We are grateful to two anonymous referees for excellent, helpful comments. This article has benefited also from discussions with Bjoern Dohëring, Mario Forni, Domenico Giannone, Stefano Neri and Ralph Setzer. This article was written while Antonio Conti was intern at the European Commission, Directorate General of Economic and Financial Affairs, and while Matteo Luciani was visiting at the Italy's Ministry of the Economy and Finance, Department of the Treasury, Economic and Financial Analysis and Planning Directorate. The hospitality of both institutions is gratefully acknowledged. Matteo Luciani acknowledges financial support from the Belgian National Bank and the IAP P6/07 contract, from the IAP programme (Belgian Scientific Policy), 'Economic policy and finance in the global economy'. Matteo Luciani is *chargé de recherches* F.R.S.-FNRS and gratefully acknowledges their financial support. The views here expressed are those of the authors and do not necessarily reflect those of the Bank of Italy or the Eurosystem.

JEL Classification numbers: C32, E32, E52.

characterized by different economic structures, legislations, fiscal policies and levels of public debt. Such a diversified environment makes the ECB decision process particularly challenging as member states' reaction to its policies might be different from country to country.

It is then natural to ask if there is any asymmetry in how single EA countries respond to the common monetary policy decided by the ECB. This is an important question both from the ECB and from member states' perspective. Indeed, while the ECB has to take into account possible asymmetries to avoid instabilities within the EA, member states have to consider their reaction to monetary policy before setting appropriate national policies.

The monetary transmission mechanism in the EA has been already investigated in the literature, both at the aggregate level (Monticelli and Tristani, 1999; Peersman and Smets, 2003; Cecioni and Neri, 2011) and among countries (Mojon and Peersman, 2003; Peersman, 2004) by means of Structural VAR (SVAR) models. In spite of some exceptions (Clements, Kontolemis and Levy, 2001; Ciccarelli and Rebucci, 2006; Rafiq and Mallick, 2008), a substantial consensus is reached by these studies on excluding asymmetric effects of monetary policy across member states.

In this article, we use a different approach. We study how single EA countries respond to ECB decisions by estimating a Structural Dynamic Factor model on a large panel of EA quarterly time series spanning the period from 1983 to 2007. We find that, although the introduction of the euro has changed the monetary transmission mechanism in the individual countries towards a more homogeneous response, EA countries react asymmetrically to the common monetary policy in terms of prices and unemployment, while no difference appears in terms of output. We conclude that these differences are the consequence of country-specific structures rather than of ECB policies, and hence they should be addressed by means of national fiscal policies, regulation and structural reforms.

Since the seminal contributions of Giannone, Reichlin and Sala (2005), Bernanke, Boivin and Eliasz (2005), Stock and Watson (2005) and Forni *et al.* (2009), a factor approach has been used as an alternative to SVAR for macroeconomic analysis. One major advantage of factor models is to allow for dealing with very large panels of data without suffering from the curse of dimensionality. Moreover, due to strong co-movements among macroeconomic time series, factor models often provide a realistic representation of the data by assuming the existence of few common shocks as the main source of business cycle fluctuations. In the present context, this implies the possibility of disentangling EA-wide from country-specific shocks: a desirable feature for analysing ECB monetary policy, which, by definition, is common to all member states.

Recent literature employed factor models to analyse EA economies although, in general, with a different focus with respect to this study. Sala (2003) studies the transmission of common monetary policy shocks across European countries but by estimating his model only on pre-euro data. Eickmeier and Breitung (2006) and Eickmeier (2009) conclude that heterogeneity across EA countries is mainly a result of idiosyncratic shocks. Favero, Marcellino and Neglia (2005) find homogeneous effects of monetary policy shocks on output gaps and inflation rates, while McCallum and Smets (2009) find heterogeneous responses in terms of real wage. Finally, Boivin, Giannoni and Mojon (2009) show that the common currency has contributed in shaping a greater homogeneity of the monetary policy transmission mechanism across countries.

Among these papers, the study most similar to ours is Boivin *et al.* (2009). However, while they are mainly interested in financial variables (bond yields, monetary aggregates and exchange rates), our main focus is on variables of economic activity (GDP and its components, prices and unemployment). In addition, the empirical procedure of our study differs from Boivin *et al*. (2009) in the choice of treatment of structural breaks and identification of monetary policy shocks. Hence, this article contributes to the procedure used to estimate a Structural Dynamic Factor model and to identify the impact of common monetary shocks on the economic activity of EA member states, and it also adds new evidence on possible cross-country asymmetries in the reaction to these shocks.

In the next Section, we outline the econometric methodology used in the empirical analysis. In Section III, we describe the data set and the data transformation used, highlighting some stylized facts related to the existence of co-movements in the EA data. In Section IV, we explain the identification strategy employed, while in Section V, we discuss country-specific impulse responses of prices, output, consumption, investment and unemployment to the common monetary policy shock. Finally, Section VI concludes. Additional results are in the Supplementary Appendix available online.

## II. Structural Dynamic Factor model

We consider here the Structural Dynamic Factor model firstly introduced by Giannone *et al*. (2005), Stock and Watson (2005) and Forni *et al*. (2009), which is a development of the model originally proposed by Stock and Watson (2002) and Bai (2003), and it is a particular case of the Generalized Dynamic Factor model by Forni *et al*. (2000) and Forni and Lippi (2001). Similar models were also proposed by Geweke (1977), Sargent and Sims (1977), Chamberlain and Rothschild (1983) and Bernanke *et al.* (2005).

We assume that there exist two kind of sources of business cycle fluctuations in the national EA economies: (i) few structural shocks common to all countries and affecting the whole EA (e.g. monetary policy or oil shocks) and (ii) many idiosyncratic shocks (capturing e.g. country/regional/sectoral-specific dynamics) having only marginal effects on the whole Area. Within this framework, we consider a shock as common if it has a non-negligible effect over all EA economies, while we consider a shock as idiosyncratic if it affects only some countries or some sectors. This representation is indeed very realistic: think, for example, of a national shock having only limited, although maybe non-null, effects outside the country where it originated, or to sectoral-specific effects as in constructions or manufacturing.

Each stationary time series $x_{it}$, $i = 1,\dots,n$ and $t = 1,\dots,T$, where $n$ is the number of variables and $T$ is the sample length, is written as the sum of two mutually orthogonal unobservable components which account for the two sources of fluctuations: (i) the common component $\chi_{it}$ and (ii) the idiosyncratic component $\xi_{it}$. The common component $\chi_{it}$ is a linear combination of $r \leqslant n$ common factors $f_{kt}$, $k = 1,\dots,r$, which are in turn driven by $q \leqslant r$ common shocks $u_{jt}, j = 1,\dots,q$.[1] Formally:

[1] The literature has often referred to $f_{kt}$ as the *static factors*, while to $u_{jt}$ as *dynamic factors*. For a formal treatment of the model presented in this Section, see Forni *et al*. (2009).

$$x_{it} = \chi_{it} + \xi_{it}, \tag{1}$$

$$\chi_{it} = \sum_{k=1}^{r} \lambda_{ik} f_{kt} = \boldsymbol{\lambda}_i' \mathbf{f}_t, \tag{2}$$

$$\mathbf{A}(L)\mathbf{f}_t = \mathbf{H}\mathbf{u}_t, \tag{3}$$

where $\boldsymbol{\lambda}_i$ is an $r$-dimensional vector of factor loadings, $\mathbf{A}(L)$ is an $r \times r$ matrix lag polynomial, $\mathbf{H}$ is an $r \times q$ matrix, $\mathbf{u}_t$, with $\mathbf{u}_t \sim iid(\mathbf{0}, \mathbf{I})$, is the $q$-dimensional vector containing the common shocks, which are orthogonal also to the idiosyncratic components at any lead and lag. The idiosyncratic components can be mildly cross-sectionally correlated, while, provided that stationarity is ensured, no assumption is made on their serial correlation properties.[2]

As proved by Forni *et al.* (2009), consistent estimation of (1–3), as both $n$ and $T$ go to infinity, and assuming both $q$ and $r$ are known, can be achieved in three steps. First, the factors $f_{kt}$ and the corresponding loadings in (2) are estimated by means of principal components. Second, $\mathbf{A}(L)$ in (3) is estimated by running a VAR on the estimated factors. Finally, given the residuals obtained from the VAR estimation, the common shocks are estimated as the $q$ largest principal components of the residuals, while $\mathbf{H}$ is estimated by projecting the residuals on the estimated shocks.[3]

From (1) to (3), we can write each observed macroeconomic variable as follows:

$$x_{it} = \sum_{j=1}^{q} b_{ij}(L) u_{jt} + \xi_{it} = \mathbf{b}_i(L)\mathbf{u}_t + \xi_{it},$$

where

$$\mathbf{b}_i(L) = \boldsymbol{\lambda}_i' \mathbf{A}^{-1}(L)\mathbf{H}, \tag{4}$$

are the impulse response functions of the common component of the $i$-th variable to the $q$ common shocks. In this article, we are just interested in the impulse response functions to the common shock representing the ECB monetary policy. Given its pervasive nature, the monetary policy shock is assumed to be one of the $q$ common shocks $u_{jt}$ and we denote it as $u_t^{mp}$. Without loss of generality, we assume the shock of interest to be the first one, so that the vector of common shocks is $\mathbf{u}_t = (u_t^{mp}, u_{2t}, \ldots, u_{qt})'$ and the impulse response functions (4) are written as follows:

$$\mathbf{b}_i(L) = b_i^{mp}(L) + \sum_{j=2}^{q} b_{ij}(L),$$

and we focus only on the first term on the right-hand side. However, it is well known that, unless additional restrictions are imposed, only the space spanned by the common factors is identified. As a consequence, impulse responses and common shocks in (4) are identified only up to multiplication by a $q \times q$ rotation matrix $\mathbf{R}$. In the present context, to

[2] The literature refers to this model as the *approximate factor model* to be distinguished from the *exact factor model* which is characterized by cross-sectionally-dynamically uncorrelated idiosyncratic component, that is, $\xi_{it} \sim iid(0, 1)$.

[3] Other estimation methods for model (1–3) have been proposed in Doz, Giannone and Reichlin (2011, 2012).

achieve identification, we impose economically meaningful restrictions as those in Forni *et al.* (2009), Forni and Gambetti (2010a) and Luciani (2013).[4]

Finally, to account for estimation uncertainty, we build confidence intervals using a bootstrap algorithm as in Bernanke *et al.* (2005) and Eickmeier (2009). At each iteration $d$, we bootstrap the estimated common shocks $\tilde{\mathbf{u}}_t{}^d$ and we generate new common factors as $\tilde{\mathbf{f}}_t{}^d = \hat{\mathbf{A}}^{*-1}(L)\hat{\mathbf{H}}\tilde{\mathbf{u}}_t^d$, where the * indicates that, as in Kilian (1998), we correct for the distortion induced by the VAR estimation on the common factors. We then estimate the parameters of equation (3) and identify the shocks as described in Section IV, thus obtaining new bootstrapped impulse response functions.[5]

Collecting together, all admissible impulse responses (the one on the sample and those on the bootstrap) gives a distribution of impulse responses from which we can get point estimates and confidence bands by computing the median and relevant percentiles.

### Structural Dynamic Factor models in the EA

Euro Area economic history is characterized by two different institutional frameworks separated by the fixing of exchange rates in January 1999. These two exchange rate regimes are likely to have determined a structural break in the data around 1999:Q1. For this reason, we assume the existence of a structural break in our data and we take it into account by proceeding as follows:

(i) we estimate the Structural Dynamic Factor model on a panel of 237 quarterly series from 1983:Q1 to 2007:Q4 to have consistent estimates of the space spanned by the common factors $\mathbf{f}_t$ and consequently of the space spanned by the common shocks $\mathbf{u}_t$;
(ii) we re-estimate the loadings $\boldsymbol{\lambda}_i$ for the pre-euro sample (1983:Q1–1998:Q4) and for the euro sample (1999:Q1–2007:Q4) separately;
(iii) we identify the monetary policy shock separately over the two subsamples.[6]

This procedure is justified on the basis of two results. On the one hand, Breitung and Eickmeier (2011) prove that in the presence of a structural breaks, factor loadings $\boldsymbol{\lambda}_i$ may be inconsistently estimated. On the other hand, Stock and Watson (2002) demonstrate that the space spanned by the common factors $\mathbf{f}_t$ can still be estimated consistently if there is *limited time variation* in the loadings. The latter is a reasonable assumption in the context of the EA, as the introduction of a single currency was indeed a gradual process which started in February 1992 with the Maastricht Treaty, and continued with the launching of the fixed exchange rate regime in January 1999 and the creation of the ECB, as established by the Treaty of Amsterdam effective since May 1999.

[4] Let $\mathbf{R}$ be a rotation matrix such that $\mathbf{RR}' = \mathbf{I}$, and let $\mathbf{c}_i(L) = \mathbf{b}_i(L)\mathbf{R}$, and $\boldsymbol{\epsilon}_t = \mathbf{R}'\mathbf{u}_t$, then the model $x_{it} = \mathbf{c}_i(L)\boldsymbol{\epsilon}_t + \xi_{it}$ is observationally equivalent to model (1–3). As in SVARs, structural shocks and impulse response functions are unique up to an orthogonal transformation (i.e. a rotation) and structural analysis in the present context becomes analogous to the standard structural analysis in VARs.

[5] As demonstrated by Bai and Ng (2006), when $N \gg T$ the estimated factors can be treated as if they are directly observed rather than estimated, and hence inference on impulse responses can be conducted by ignoring the idiosyncratic component.

[6] Since Weber, Gerke and Worms (2011) find that the transmission mechanism of monetary policy was similar before 1996 and after 1999 but different during the transition period 1996–1998, as a robustness check we perform our analysis on the subsample 1983–1996 rather than 1983–1998. Results are identical to the one obtained with our benchmark specification, and are available in the Supplementary Appendix to this article.

Given their relevance for our analysis, we need to formally test these assumptions on the behaviour of the common factors and their loadings. To do so, we first run a CUSUM Square test on the common factors (Brown, Durbin and Evans, 1975) and find no significant structural change (Figure 1). Then, we run the structural break test of Breitung and Eickmeier (2011) on the factor loadings, which indicates structural break on 1 January 1999 for all the series of interest (Table 1). These results are consistent with Breitung and Eickmeier (2011) and Canova, Ciccarelli and Ortega (2012).

### Testing for asymmetries

To evaluate the presence of significant differences across impulse responses, we should test the null-hypothesis of no differences. Olivei and Tenreyro (2010) propose a procedure to test for differences among impulse responses in VAR models. Their test consists in computing differences among observed impulse responses and then compare them with a distribution of distances obtained from data simulated from two different VARs. Unfortunately, this test is unfeasible in our case. Indeed, while Olivei and Tenreyro (2010) aim at comparing impulse responses of the same variable, but estimated from two different (VAR) models, we are interested in comparing responses of different variables, but estimated from the same (factor) model. Hence, in our case we should be able to simulate data from a factor model in which all impulse responses are equal. Building such a distribution would lead

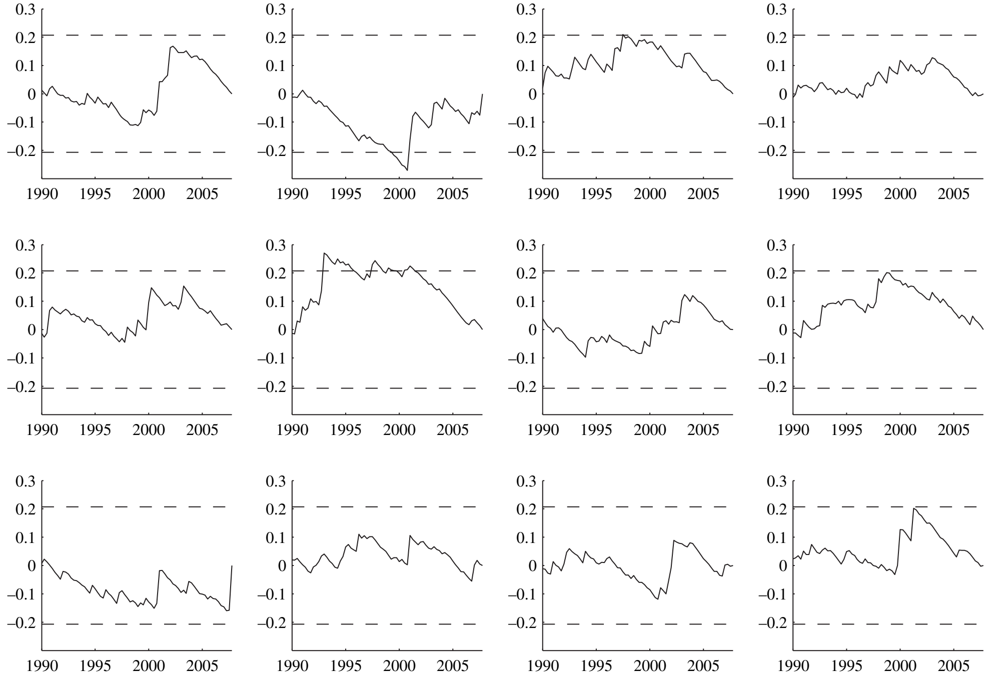

Figure 1. CUSUM square test on the static factors
*Notes*: Solid line is the CUSUM Square statistic of Brown *et al.* (1975), while the dashed lines are the 90% confidence bands computed using critical values as given in Durbin (1969) and Edgerton and Wells (1994).

TABLE 1

*Testing for structural break in the factor loadings: Breitung and Eickmeier test*

| | *BG* | *FR* | *GE* | *IT* | *NL* | *ES* | *FI* | *GR* | *IE* | *PT* | *EA* |
|---|---|---|---|---|---|---|---|---|---|---|---|
| Consumer Price Index | 63.44 | 71.35 | 72.50 | 55.46 | 60.86 | 63.32 | 41.56 | 37.75 | 48.56 | 57.69 | 80.68 |
| Gross Domestic Product | 81.41 | 74.14 | 76.29 | 54.57 | 58.66 | 79.83 | 70.87 | 30.82 | 39.30 | 45.10 | 88.59 |
| Consumption | 70.35 | 55.96 | 65.19 | 66.52 | 34.19 | 56.59 | | | | | |
| Investment | 49.91 | 66.55 | 52.65 | 51.76 | 48.81 | 68.91 | | | | | |
| Unemployment Rate | 66.52 | 55.61 | 34.15 | 47.61 | 54.45 | 48.21 | | | | | |

*Notes*: This Table shows the LM Statistic for the null of no structural break in the factor loadings on January the 1st 1999. This statistic is asymptotically distributed as a $\chi^2$ random variable with $r$ (number of factors degrees of freedoms). The 10%, 5% and 1% critical values are 18.5493, 21.0261 and 26.2170 respectively.

to a degenerate distribution in which all distances among impulse responses are zero, that is, a useless distribution for making inference.

Therefore, to have an approximate measure of asymmetries, we rely on a simple procedure with a clear and intuitive meaning. In particular, for each bootstrap, we compute the difference between the individual country response and the EA response. These gives us a distribution of differences between impulse responses. We consider the difference non-significant if zero is contained within the confidence bands.[7] A similar procedure is used also by Fielding and Shields (2011).

## III. Model set-up

### Data and data treatment

Data include EA aggregates, main macroeconomic variables for single EA member states, and key indicators for UK, USA and Japan. The database contains nine aggregate EA variables: GDP, CPI, short- and long-term rates, monetary aggregates (M1 and M3), unit labour cost, real effective exchange rate and the dollar/euro exchange rate. These aggregate variables are taken either from Eurostat, or from ECB, and, when necessary, they are backdated using data from the Area Wide Model Database (Fagan, Henry and Mestre, 2001).

We then have 35 variables for Germany, Italy and the Netherlands, 34 for France and Spain and 31 for Belgium. Variables included for these countries are interest rates, monetary aggregates, real effective exchange rate, an index of stock prices, GDP and its expenditure components, unemployment rate, unit labour costs, GDP deflator, producer price index, CPI together with its disaggregated categories, retail sales and number of cars sold. In addition, we also include CPI, GDP and interest rates for smaller EA countries (Finland, Greece, Ireland, Luxembourg and Portugal), and for UK, USA and Japan, as well as the spot oil price.

[7] We thank an anonymous referee for suggesting us this testing procedure.

TABLE 2

*The distribution of autocorrelations*

| *Percentile* | *Lag* | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| *Light* | 1 | 2 | 3 | 4 | 5 | 6 | 7 | 8 |
| 5 | 0.65 | 0.40 | 0.12 | 0.03 | 0.03 | 0.04 | 0.02 | 0.02 |
| 25 | 0.81 | 0.57 | 0.32 | 0.11 | 0.13 | 0.12 | 0.09 | 0.07 |
| 50 | 0.86 | 0.67 | 0.47 | 0.28 | 0.23 | 0.22 | 0.19 | 0.16 |
| 75 | 0.90 | 0.78 | 0.65 | 0.51 | 0.43 | 0.36 | 0.32 | 0.29 |
| 95 | 0.95 | 0.87 | 0.78 | 0.70 | 0.64 | 0.58 | 0.54 | 0.51 |
| *Heavy* | 1 | 2 | 3 | 4 | 5 | 6 | 7 | 8 |
| 5 | 0.08 | 0.02 | 0.01 | 0.01 | 0.01 | 0.01 | 0.01 | 0.01 |
| 25 | 0.24 | 0.09 | 0.06 | 0.07 | 0.06 | 0.04 | 0.04 | 0.04 |
| 50 | 0.36 | 0.18 | 0.15 | 0.15 | 0.14 | 0.11 | 0.10 | 0.10 |
| 75 | 0.54 | 0.32 | 0.27 | 0.25 | 0.23 | 0.18 | 0.18 | 0.16 |
| 95 | 0.90 | 0.74 | 0.59 | 0.50 | 0.39 | 0.35 | 0.30 | 0.33 |

*Notes*: Percentiles of the distribution of univariate autocorrelation functions when computing *light* transformations as in Boivin *et al.* (2009) or *heavy* transformations, that is, by replacing yearly with quarterly growth rates and taking first differences in interest rates and second differences in the log of prices and monetary aggregates.

Summing up, our data sets consists of 237 quarterly time series covering the period 1983Q1–2007Q4.[8] The complete list of variables, sources and the transformations used is available in the Supplementary Appendix.

A comment is necessary on the way we transform data to make them stationary. According to Uhlig (2009), the co-movements found by Boivin *et al.* (2009) in a similar data set, are actually the result of the autocorrelation induced in the transformed data. As a consequence, the existence of a factor structure, based on co-movements among series, would be just a by-product of data treatment. To cope with this critique, and compared with Boivin *et al.* (2009), we adopt a different set of transformations. As in Stock and Watson (2005) and Forni and Gambetti (2010c), we take second differences in the log of both prices and monetary aggregates, first differences in interest rates, and, when needed, growth rates are computed on a quarterly basis.[9] We label this kind of transformations as *heavy*, in contrast with *light* transformations used by Boivin *et al.* (2009). In the latter case, interest rates are kept in levels, the first difference of log of prices and monetary aggregates is taken, and, most importantly, growth rates are computed on a yearly basis.

With reference to Uhlig (2009) critique, our choice of *heavy* transformations is justified by Table 2, where we report selected percentiles of the distribution of the absolute value of univariate autocorrelations when considering *light vs. heavy* transformations.

[8]The sample starts in 1983 because of two main reasons. First, not all the series in the database, especially at the single European country level, are available before 1983. Second, although EA data at the aggregate level are available since 1970 at a quarterly frequency, by comparing alternative aggregation methods, Bosker (2006) shows that differences in EA artificial data are prominent before 1983 especially for inflation and interest rates, while vanishing thereafter.

[9]It is worth to note that, within the literature on money demand in the EA (Papademos and Stark, 2010, and reference therein), it is common practice to treat monetary aggregates as $I(2)$ variables. Furthermore, Beyer (2009) and Dreger and Wolters (2010), among others, show that inflation is an important determinant in describing a stable long run money demand equation for the EA, thus indicating that money growth and inflation are co-integrated, and therefore $I(1)$ variables.

TABLE 3

*Determining the number of common shocks: Onatski test*

| $q_0$ vs. $q_1$ | *1* | *2* | *3* | *4* | *5* | *6* | *7* | *8* |
|---|---|---|---|---|---|---|---|---|
| 0 | 0.029 | 0.050 | 0.069 | 0.088 | 0.104 | 0.121 | 0.135 | 0.151 |
| 1 | | 0.271 | 0.487 | 0.626 | 0.321 | 0.372 | 0.421 | 0.465 |
| 2 | | | 0.608 | 0.677 | 0.262 | 0.321 | 0.372 | 0.421 |
| 3 | | | | 0.390 | 0.195 | 0.262 | 0.321 | 0.372 |
| 4 | | | | | 0.108 | 0.195 | 0.262 | 0.321 |
| 5 | | | | | | 0.947 | 0.923 | 0.343 |
| 6 | | | | | | | 0.623 | 0.257 |
| 7 | | | | | | | | 0.142 |

*Notes*: This Table shows $p$-values of the null of $q_0$ common shocks against the alternative of $q_0 < q \leqslant q_1$ common shocks. The Discrete Fourier Transformation of the data is computed for $\omega_j = 2\pi s_j/T$ , with $s_j \in [2,\ldots,20]$, thus to includes waves between 1 and 12 years.

The median autocorrelation from lags 1 to 4 is between 0.36 and 0.15 in the *heavy* case, while it is between 0.86 and 0.28 in the *light* case. Similar results hold also for other percentiles.

### Number of common shocks and factors

After transforming data, we rely on specific tests and information criteria for determining the number of common factors $r$ and common shocks $q$. The latter is estimated by means of the test proposed by Onatski (2009), which suggests $q \in \{4, 5\}$ (Table 3) and the criterion by Hallin and Liška (2007) suggesting $q \in \{2, 3\}$. We choose as our baseline specification $q = 4$, that is, the average of these results.[10]

One possible way of fixing the number of common factors is to choose $r$ such that the variance explained by the factors is equal to the variance explained by the chosen $q$ shocks. This heuristic method suggests 13 factors (Table 4). An alternative is to resort to the criterion provided by Bai and Ng (2002), and its refinement by Alessi, Barigozzi and Capasso (2010), both suggesting either 9 or 14 factors. We choose as our baseline specification $r = 12$, that is, the average of what the mentioned criteria suggest.[11]

In Table 5, we show the share of variance accounted for by the estimated common component. When looking at the post-1999 sample, we find that 91% of aggregate GDP and 90% of aggregate CPI fluctuations are imputable to the common component. These values decrease if we look at country-specific GDP, CPI and unemployment rate, but are still considerably high in the majority of the cases, notwithstanding the heterogeneity and

[10] It is also worth noting that four common shocks is a parameterization considered plausible in the literature. In particular, in her discussion of Boivin *et al.* (2009), Reichlin (2009) rises some doubts about their choice of seven common shocks by arguing that a smaller number of common shocks would be much more plausible: 'when macroeconomists think of common shocks, they mention productivity, money, time preference or government, and it is difficult to think of many other candidates' (p. 130).

[11] Other criteria, not used in this article, to determine $q$ or $r$ are in Amengual and Watson (2007), Bai and Ng (2007), Kapetanios (2010) and Onatski (2010). Results for the criteria by Bai and Ng (2002), Hallin and Liška (2007) and Alessi *et al.* (2010), as well as robustness analysis for $q = \{3, 5\}$ and $r = \{9, 13\}$, are available in the Supplementary Appendix to this article.

TABLE 4

*Cumulated explained variance*

| *Number of factors* | 1 | 2 | 3 | 4 | 5 | 6 | 7 | 8 | 9 | 10 | 11 | 12 | 13 | 14 |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| $q$ | 0.21 | 0.34 | 0.43 | 0.51 | 0.58 | 0.64 | 0.68 | 0.73 | 0.76 | 0.79 | 0.82 | 0.84 | 0.86 | 0.88 |
| $r$ | 0.09 | 0.16 | 0.23 | 0.27 | 0.31 | 0.34 | 0.37 | 0.40 | 0.42 | 0.45 | 0.47 | 0.49 | 0.51 | 0.53 |

*Notes*: The Table shows the percentage of overall variance explained by the first $q$ common shocks estimated with the method of dynamic principal components as in Forni *et al.* (2000), and the first $r$ static factors estimated by static principal components. Variance is measured on a scale between 0 and 1.

TABLE 5

*Comovements in the Euro Area: explained variance*

| *Country* | *GDP* | | *CPI* | | *UR* | |
|---|---|---|---|---|---|---|
| | *1983–1998* | *1999–2007* | *1983–1998* | *1999–2007* | *1983–1998* | *1999–2007* |
| Euro Area | 0.85 | 0.91 | 0.79 | 0.90 | – | – |
| Germany | 0.71 | 0.76 | 0.69 | 0.80 | 0.48 | 0.69 |
| France | 0.74 | 0.78 | 0.49 | 0.82 | 0.55 | 0.60 |
| Netherlands | 0.31 | 0.73 | 0.73 | 0.75 | 0.54 | 0.58 |
| Belgium | 0.66 | 0.59 | 0.55 | 0.81 | 0.58 | 0.47 |
| Finland | 0.63 | 0.54 | 0.31 | 0.60 | – | – |
| Italy | 0.42 | 0.66 | 0.70 | 0.82 | 0.50 | 0.36 |
| Spain | 0.38 | 0.66 | 0.62 | 0.90 | 0.67 | 0.57 |
| Portugal | 0.67 | 0.45 | 0.51 | 0.70 | – | – |
| Ireland | 0.39 | 0.12 | 0.33 | 0.72 | – | – |
| Greece | 0.20 | 0.44 | 0.18 | 0.54 | – | – |

*Notes*: For each country, we report the variance explained by the common component of GDP, CPI and Unemployment Rate (UR). For each variable the first column refers to the 1983:Q1–1998:Q4 (pre-euro) sample, and the second column to the 1999:Q1–2007:Q4 (euro) sample. Values are given on a scale between 0 (no contribute of the common component) and 1.

large dimension of the data set at hand. Indeed, the variance of the common component is more than 60% of total GDP fluctuations for all countries but Belgium (59%), Finland (54%), Portugal (45%), Greece (44%) and Ireland (12%), while it is more than 70% of total CPI fluctuations for all countries but Portugal (70%), Finland (60%) and Greece (54%), and more than 50% of total unemployment fluctuations for all countries but Belgium (47%) and Italy (36%). When averaging common variances across all 237 considered variables, we have that the common component accounts for 51% of the total fluctuations.

The existence of cross-country heterogeneity in the co-movements both justifies our approach (co-movements imply a factor structure), and motivates our research question (heterogeneity suggests asymmetric reactions).

## IV. Identification of the monetary policy shock

We identify the monetary policy shock by means of sign restrictions (Faust, 1998; Canova and de Nicolò, 2002; Uhlig 2005), an identification strategy also used in the context of factor models by Eickmeier (2009) and Forni and Gambetti (2010b,c). Specifically, at each iteration we draw a vector of $q(q-1)/2$ angles $\boldsymbol{\omega}$ from a uniform distribution on $[0, 2\pi)$, which, by means of Givens transform, are used to construct an orthogonal matrix $\mathbf{R}(\boldsymbol{\omega})$ of dimension $q \times q$. We then compute the associated impulse responses and if they satisfy a prescribed set of sign restrictions (to be specified below) we accept the draw, otherwise we discard it. We stop this procedure once $K$ draws are accepted.

We rely on the following assumptions imposed only on EA variables for the first two lags: after a contractionary monetary policy shock the short-term interest rate, the real effective exchange rate and the dollar/euro exchange rate increase, while GDP, CPI and M1 decrease. These restrictions, which are theoretically consistent with a typical IS-LM model of an open economy, are commonly accepted in the literature (Peersman, 2005; Farrant and Peersman, 2006). Moreover, the choice of imposing restrictions only on the EA variables makes the identification scheme 'agnostic' on the responses of single countries (see also Eickmeier, 2009, for a similar identification scheme).

The choice of an 'agnostic' identification scheme leaves the room open to non-conventional reactions (Section V). Therefore, we also tried to impose country-specific restrictions, but we did not find rotation matrices able to satisfy all the restrictions. This result has an economic interpretation as discussed in Sections V and VI. In particular, we are able to satisfy some of, but not all, the country-specific restrictions. More specifically, restrictions on GDP are easily satisfied, but for Greece in the 1983–1998 sample, whereas restrictions for prices are satisfied only for Belgium, France, Germany and the Netherlands. There are very few rotations that satisfy the restrictions for prices in Spain and in Italy, but almost no rotations that satisfy the restrictions jointly for Italy and Spain. For Finland, Greece and Ireland we cannot find restrictions for the pre-euro sample, while for Portugal we cannot find them for the post-euro sample. Finally, it is worth noting that unconventional reactions of inflation or prices to the monetary policy shock are also found by Sala (2003) for Italy and Portugal, Eickmeier (2009) for Greece and Portugal, Boivin *et al.* (2009) for Germany, and Peersman (2004) for Austria and Italy.

To compute impulse responses and the related confidence intervals, we use the procedure described in Section II with 500 bootstrap draws. To keep computations feasible, for each of the $500+1$ samples we save $K = 10$ rotation matrices. Then, for each sample we select just one rotation matrix as suggested by Fry and Pagan (2011).[12]

An alternative strategy to identify monetary policy is to adopt a recursive identification scheme, that is, the Cholesky decomposition as in Boivin *et al.* (2009) and Forni and Gambetti (2010a). Although recently criticized (Canova and Pina, 2005; Carlstrom, Fuerst and Paustian, 2009; Uhlig, 2009; Castelnuovo, 2011, 2012a), this is the simplest, and perhaps, still, the most diffused identification scheme in SVAR literature (Christiano, Eichenbaum and Evans, 1999; Peersman and Smets, 2003; Weber *et al.*, 2011). The main problem of this identification scheme is that it relies on zero short-run restrictions, which are too binding and not necessarily based on economic theory. Differently, using sign restrictions, we are imposing restrictions often used implicitly in empirical analysis to validate the results, and consistent with macroeconomic models. However, if the shock

[12] Fry and Pagan (2011) point out that for each sample the distribution of the $\mathbf{R}(\omega)$ that satisfies the sign restrictions represents model uncertainty. However, when computing impulse responses with confidence bands what matters is sampling uncertainty, not model uncertainty. Hence, they suggest selecting for each sample just one rotation, namely the one which produces the impulse response closest to the median response.

of interest explains a marginal fraction of the forecast error of the variables of interest, the outcome of the exercise conducted with sign restrictions should also be taken cautiously (for a Monte Carlo experiment, see Paustian, 2007; Castelnuovo, 2012b). For these reasons, in the Supplementary Appendix we show also results obtained with Cholesky identification.

## V. Results

In this Section, we present impulse response functions to a monetary policy shock. The shock is normalized so that on impact it raises EA short-term rate of 50 basis points. In Figures 2–7, we show the impulse responses of CPI, GDP, together with consumption and investment, and unemployment rate both at the aggregate level, and at the country level. Each Figure contains the impulse responses, together with 68% confidence bands, estimated both on the pre-euro sample (grey solid line, and shaded area), and on the euro-sample (black solid and dashed lines).

Figure 2 shows impulse responses for the aggregate EA variables used in the identification of the monetary policy shock and it should be considered just as validation of our identification strategy. In both samples, output, prices, and the monetary aggregate M1

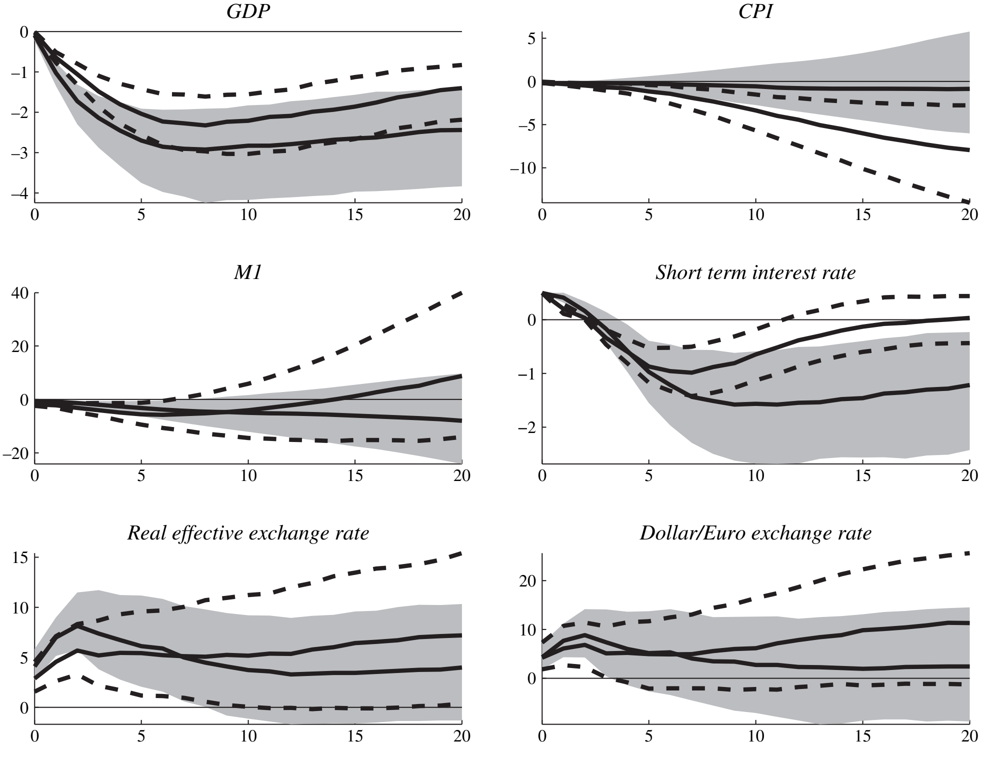


Figure 2. Impulse responses to a monetary policy shock: Euro Area aggregates
*Notes*: Solid line is the estimated impulse responses for the 1999:Q1–2007:Q4 (euro) subsample with 68% bootstrap confidence band (dashed). Shaded area is the 68% confidence band for the 1983:Q1–1998:Q4 (pre-euro) subsample.

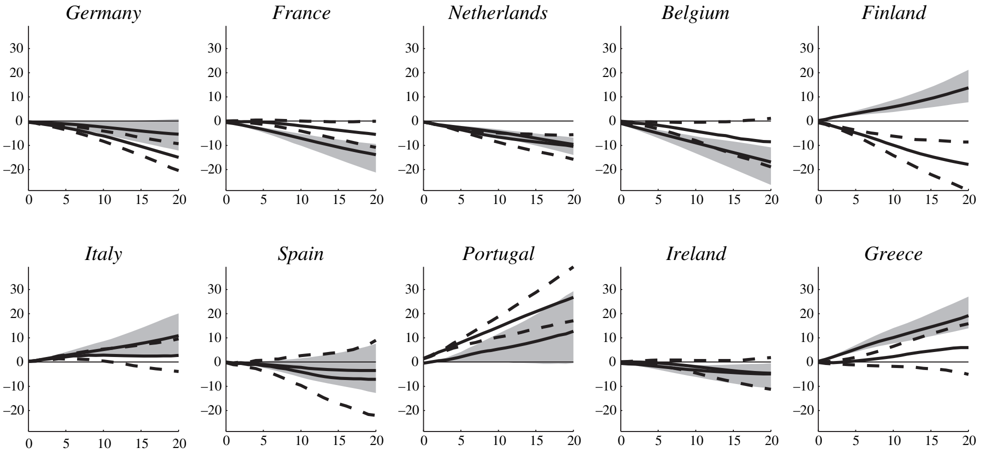


Figure 3. Impulse responses to a monetary policy shock: Consumer Price Index
*Notes*: Solid line is the estimated impulse responses for the 1999:Q1–2007:Q4 (euro) subsample with 68% bootstrap confidence band (dashed). Shaded area is the 68% confidence band for the 1983:Q1–1998:Q4 (pre-euro) subsample.

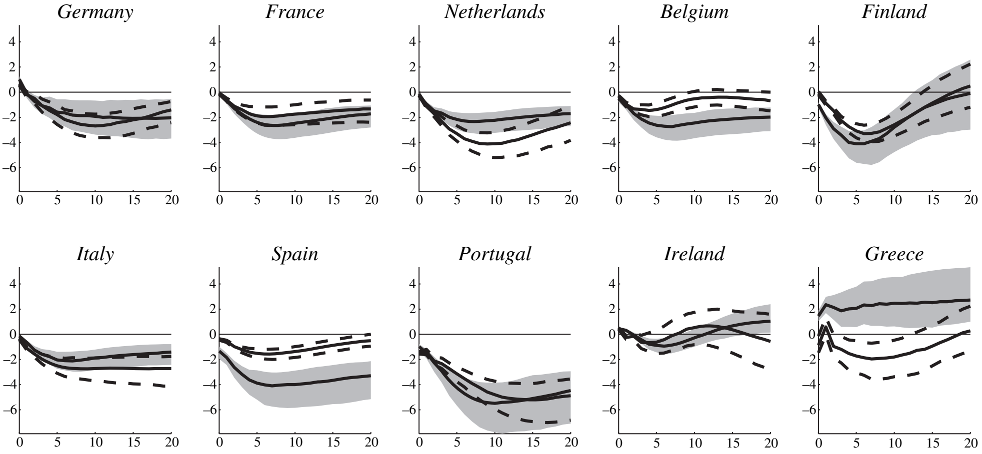


Figure 4. Impulse responses to a monetary policy shock: Gross Domestic Product
*Notes*: Solid line is the estimated impulse responses for the 1999:Q1–2007:Q4 (euro) subsample with 68% bootstrap confidence band (dashed). Shaded area is the 68% confidence band for the 1983:Q1–1998:Q4 (pre-euro) subsample.

respond negatively, while the short-term rate and exchange rates respond positively. Notice that we estimate a strong effect of monetary policy shocks on the economy (in particular on both GDP and CPI).[13] However, our estimates are not far from those usually found by the literature (Monticelli and Tristani, 1999; Sala, 2003; Eickmeier, 2009).[14]

[13]The reason for these large magnitudes in the response of prices is due to the heavy transformations choice. In particular, since we model CPI as an $I(2)$ process, we obtain explosive dynamics due to the need of cumulating twice the estimated IRF.

[14]Detailed information on the magnitude of the impulse responses estimated by the literature cited in this article can be found in the Supplementary Appendix.

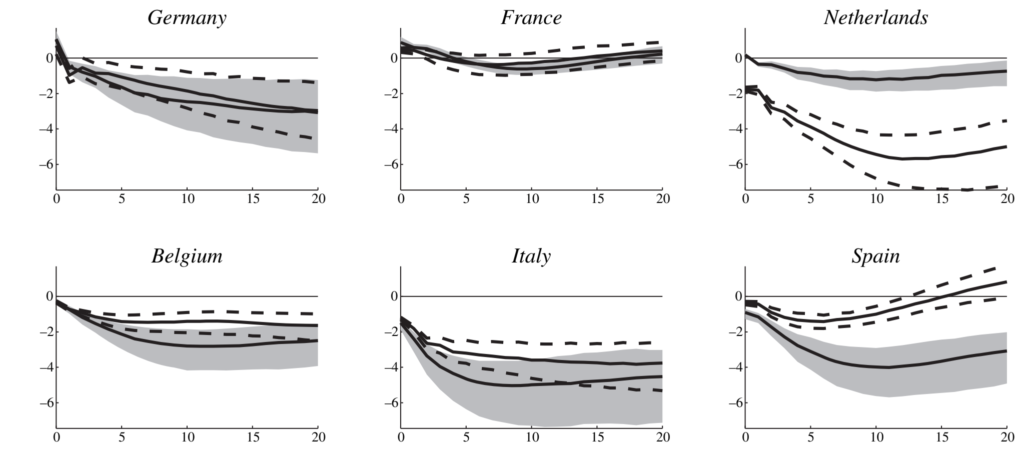


Figure 5. Impulse responses to a monetary policy shock: Consumption
*Notes*: Solid line is the estimated impulse responses for the 1999:Q1–2007:Q4 (euro) subsample with 68% bootstrap confidence band (dashed). Shaded area is the 68% confidence band for the 1983:Q1–1998:Q4 (pre-euro) subsample.

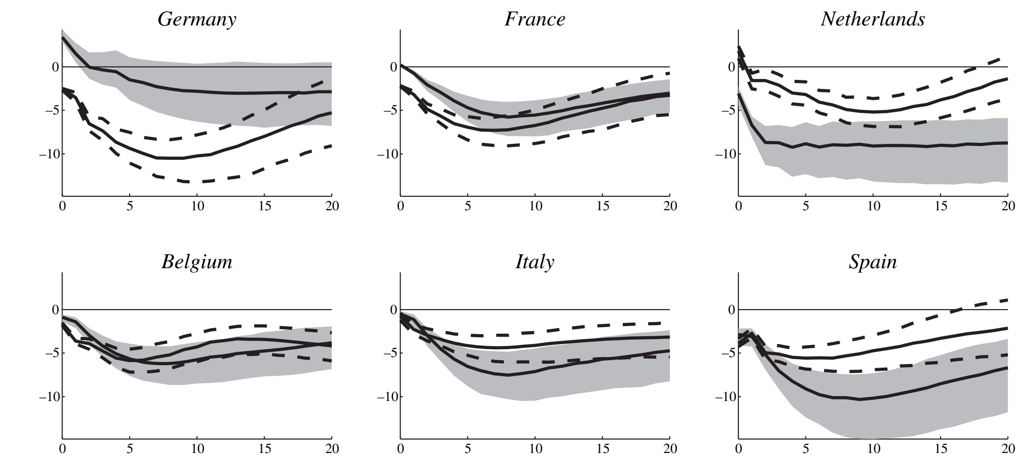


Figure 6. Impulse responses to a monetary policy shock: Investment
*Notes*: Solid line is the estimated impulse responses for the 1999:Q1–2007:Q4 (euro) subsample with 68% bootstrap confidence band (dashed). Shaded area is the 68% confidence band for the 1983:Q1–1998:Q4 (pre-euro) subsample.

When comparing pre-euro with euro sample impulse responses of aggregate CPI and GDP, we find that the introduction of the common currency amplified the response of CPI, while reducing the reaction of GDP. This result is consistent with an increase in prices' flexibility due to greater competition between EA industries.[15]

[15] It may be argued that, since before 1999 there was no common monetary policy, the relevant comparison would be between pre-1999 Bundesbank monetary policy and post-1999 ECB monetary policy (Sala, 2003). Hence, as a robustness check we estimated our model by imposing in the pre-1999 sample the identifying restrictions on the German short-term interest rate. Results are nearly identical to the one obtained with our benchmark specification, and are available in the Supplementary Appendix to this article.

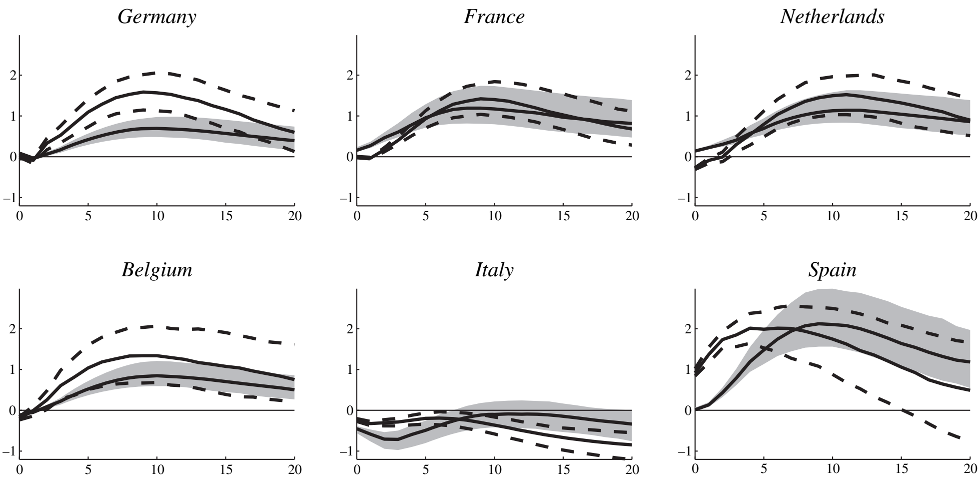


Figure 7. Impulse responses to a monetary policy shock: Unemployment Rate
*Notes*: Solid line is the estimated impulse responses for the 1999:Q1–2007:Q4 (euro) subsample with 68% bootstrap confidence band (dashed). Shaded area is the 68% confidence band for the 1983:Q1–1998:Q4 (pre-euro) subsample.

We then move to the analysis of country-specific variables. In Figure 8, we show results of the test on the asymmetries introduced in Section II. In each plot of Figure 8, the grey/black straight line is the median difference between the response of a given country and the response of a benchmark country, while the shaded area/dashed lines is/are the 68% confidence bands estimated on the pre-euro/euro sample. If at horizon $h$ the zero is contained within the confidence bands, it means that the impulse response of a given country and that of a benchmark country are not statistically different at horizon $h$. In panels (a) and (b), the benchmark country is the EA, while in panels (c–e) the benchmark country is Germany.

The goal is to understand whether there are asymmetries in the transmission mechanism of the common monetary policy to EA countries before and after 1999, and to understand which was the effect, if any, of the common monetary policy on the existing asymmetries. Notice also that, in terms of our research question what matters is the direction (i.e. signs) and the significance of the cross-country differences rather than the magnitudes, which turn out to be implausibly high for some of the countries.

**Cross-country differences before 1999**

*Prices*

Four countries out of ten (Finland, Italy, Portugal and Greece) exhibit a positive reaction (Figure 3). In addition to the four countries just mentioned, also the impulse responses of France, Belgium and the Netherlands are statistically different from those of the EA (Figure 8a) thus showing a high degree of pre-euro heterogeneity in prices.

*GDP*

All countries, but Greece, react as predicted by economic theory (Figure 4). The unconventional response of Greece seems to be related with the low percentage of variance of

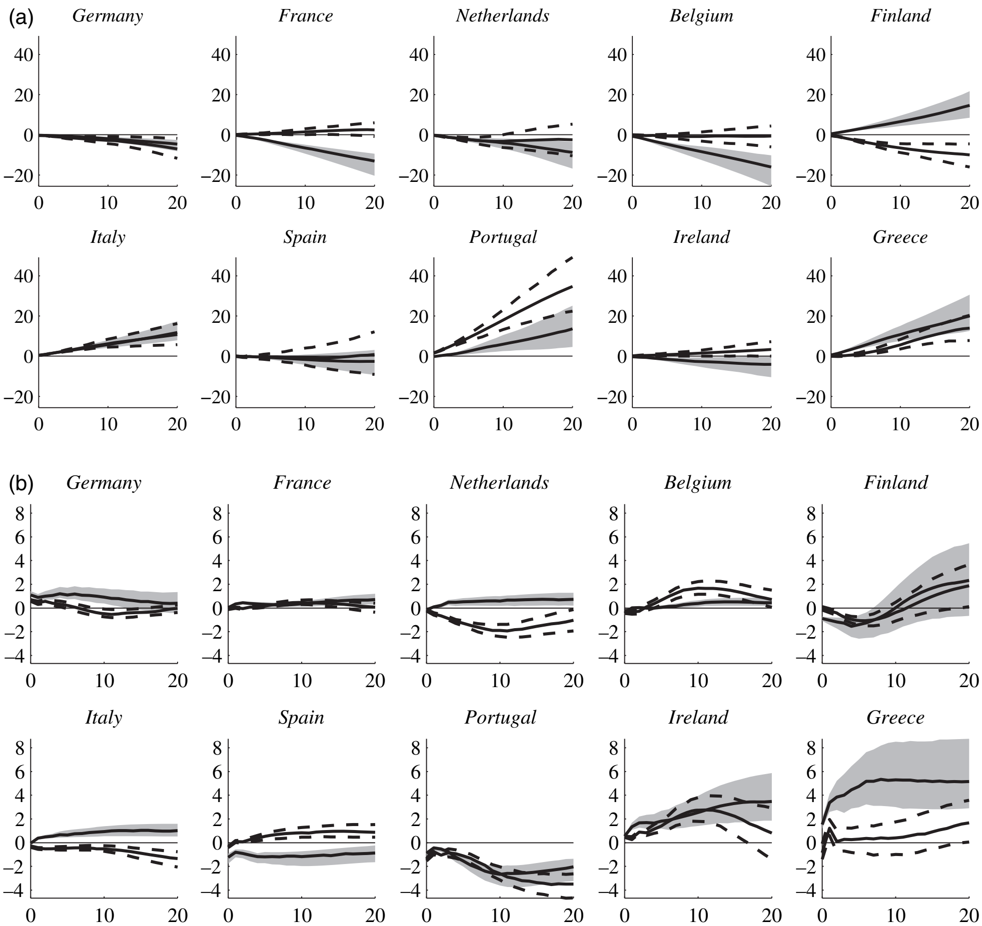


Figure 8. Quantifying asymmetries: distance from benchmark country. (a) Consumer Price Index, (b) Gross Domestic Product (c) Consumption, (d) Investments (e) Unemployment Rate

*Notes*: In each plot, the grey/black straight line is the median difference between the response of a given country and the response of a benchmark country, while the shaded area/dashed lines is/are the 68% confidence bands estimated on the pre-euro/euro sample. If at horizon *h* the zero is contained within the confidence bands, it means that the impulse response of a given country and that of a benchmark country are not statistically different at horizon *h*. In panels (a) and (b) the benchmark country is the Euro Area, while in panels (c–e) the benchmark country is Germany.

Greek GDP explained by the common component (Table 5). Indeed, it should be noted that Greece was not part of the European Monetary System, as it only joined it in stage II, that is, in 1999. When testing for asymmetries (Figure 8b) we find that also the reactions of Germany, the Netherlands, Italy, Spain, Ireland and Portugal are statistically different from those of the EA thus showing a high degree of heterogeneity pre-euro in output.

### *Consumption*

All countries display the expected negative path (Figure 5). However, when testing for asymmetries (Figure 8c) some differences emerge since both France and the Netherlands react significantly less than Germany, while Italy and Spain react significantly more.

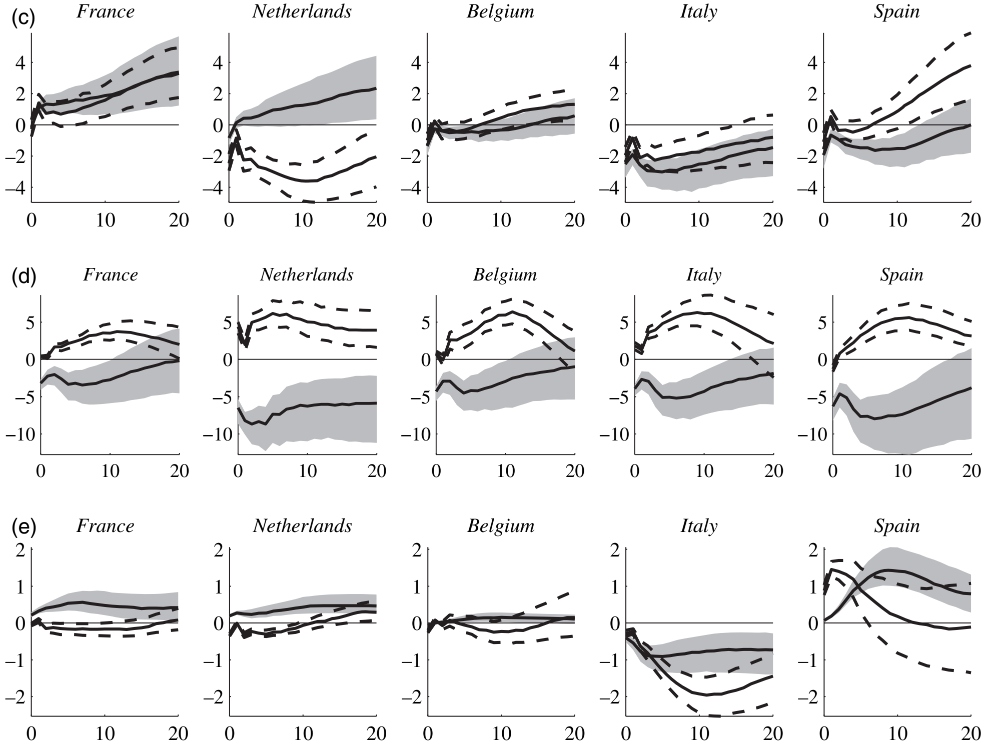


Figure 8. *(Continued)*

*Investment*

All countries react as predicted by economic theory (Figure 6) and all react significantly differently than Germany (Figure 8d).

*Unemployment Rate*

Unemployment in Germany, France, the Netherlands and Belgium follows a similar hump-shaped path, while impulse responses for Spain and Italy are different (Figure 7). However, when testing for asymmetries (Figure 8e) we find that also France and the Netherlands react in a significantly different way than Germany.

**Cross-country differences after 1999**

*Prices*

We can divide EA countries in three groups (Figure 3). In the first, we have countries for which we observe the expected negative response to the common monetary policy shock: Germany, France, the Netherlands and Finland. In the second group, we have countries for which we estimate a *mute* response (i.e. not significantly different from zero): Ireland, Belgium, Spain, Italy and Greece. Finally, the third group is composed of a single country for which we estimate a positive response, namely Portugal (Sala, 2003; Eickmeier, 2009). If we consider these results with reference to pre-euro impulse responses, the introduction of a common currency appears to have had a positive role in shaping homogeneity across countries CPIs. Table 5 provides an explanation of this switch in terms of explained variances of the common components: the higher this number, the more impulse responses are homogeneous. However, significant asymmetries persist between the EA and Finland, Italy, Portugal and Greece (Figure 8a), with Finland reacting more than the EA, and Italy, Portugal, and Greece reacting less. The response of the Mediterranean countries seems likely to be the consequence of price rigidities and of lack of competition.

*GDP*

Impulse responses are quite similar across countries (Figure 4). With respect to the pre-euro sample, the response of Greek GDP has now the expected negative sign. Overall, the introduction of the euro has helped in reducing asymmetries. However, small differences are still present between the EA and the Netherlands, Ireland, and Portugal (Figure 8b). While Ireland and Portugal GDP fluctuations are mainly idiosyncratic (Table 5), the strong reaction of Dutch GDP seems to be driven by consumption.

*Consumption*

Italy and the Netherlands display the deepest reaction in consumption with a minimum of roughly −5% and −3%, respectively (Figure 5), a result also found by Reichlin (2009) in the case of Italy. The response of Netherlands consumption is likely due to the particular dynamics of the series, which has nothing to do with monetary policy. Indeed, from 1999 to 2003 the year on year consumption growth trended downward as a consequence of firms' and households' balance sheets adjustments, weak profits and lower purchasing power of households. Germany, Belgium and Spain also show a significant contraction in consumption up to −1%, while the response is mute for France. The introduction of the euro has slightly reduced asymmetries for Italy and Spain (Figure 8c).

*Investment*

The reaction of investment is more homogeneous with the main exception of Germany for which a contraction up to −9% is observed (Figure 6). This result is likely due to the dynamics of the German construction sector, as the housing market was characterized by a post-reunification boom-bust cycle in residential investment (Knetsch, 2010). This anomalous response of Germany implies significant differences with respect to all other countries (Figure 8d), which, however, have all similar responses.

*Unemployment Rate*

As in the pre-euro sample, all countries but Italy and Spain show similar reactions (Figure 7). However, asymmetries are reduced between Germany and France and the Netherlands (Figure 8e). More specifically, on the one hand Spanish unemployment rate experiments a stronger boost than other countries, on the other hand, Italian unemployment seems not to respond to a common monetary policy shock. The first finding suggests large elasticity of Spanish labour market to monetary policy shocks likely due to the high share of fixed term contracts in the labour market (Güell and Petrongolo, 2007). In contrast, the mute response of Italian unemployment is the consequence of a rigid labour market which seems not to be related at all to the business cycle as confirmed from the low correlation (−0.07) between changes in unemployment rate and GDP growth.[16]

## VI. Discussion and conclusions

In this article, we ask the following question: is there any asymmetry in how single EA countries respond to the common monetary policy decided by the ECB?

To answer we estimate a Structural Dynamic Factor model on a large panel of EA quarterly time series spanning the period from 1983 to 2007. The data set incorporates data on the aggregate EA as well as country-specific key economic variables, such as gross domestic product, inflation, unemployment, consumption, investment and many others.

We find that, although the introduction of the euro has changed the monetary transmission mechanism in the individual countries towards a more homogeneous response, differences still remain between North and South Europe in terms of prices and unemployment. Due to their idiosyncratic nature, these differences can hardly be controlled by means of the common monetary policy; rather they should be addressed by means of national fiscal policies, regulation and structural reforms. Indeed, while before 1999 CPI responses were highly asymmetric, the introduction of the euro and of the single monetary policy, and the consequent increase in integration and competition within the EA, made prices more flexible thus responding more homogeneously to changes in interest rate. The remaining asymmetries are observed in the Mediterranean countries, which historically have less flexible prices and lack of market competition. Similarly, the asymmetries in labour markets seem to be the result of structural and socio-economic characteristics of single countries. This is the case for example with the rigid labour market structure in Italy, which makes Italian unemployment rate completely unresponsive to the single monetary policy.

In conclusion, EA countries react asymmetrically to the common monetary policy in terms of prices and unemployment, while no difference appears in terms of output. While the post-1999 reduction in asymmetries is consistent with the aims of the ECB (Boivin *et al.*, 2009), the remaining differences are beyond the scope of monetary policy, and they should be addressed by means of national reforms. As demonstrated by the recent/current public debt crisis, and by the skyrocketing of government bond spreads, these differences pose a threat to the region's stability: addressing them is fundamental for the future of Europe, and it should be a priority if economic cohesion is to be achieved.

*Final Manuscript Received: May 2013*

## References

Alessi, L., Barigozzi, M. and Capasso, M. (2010). 'Improved penalization for determining the number of factors in approximate static factor models', *Statistics & Probability Letters*, Vol. 80, pp. 1806–1813.

[16] Correlations for other countries are Belgium −0.27, France −0.36, Germany −0.29, the Netherlands −0.26 and Spain −0.34.

Amengual, D. and Watson, M. W. (2007). 'Consistent estimation of the number of dynamic factors in a large $N$ and $T$ panel', *Journal of Business & Economic Statistics*, Vol. 25, pp. 91–96.
Bai, J. (2003). 'Inferential theory for factor models of large dimensions', *Econometrica*, Vol. 71, pp. 135–171.
Bai, J. and Ng, S. (2002). 'Determining the number of factors in approximate factor models', *Econometrica*, Vol. 70, pp. 191–221.
Bai, J. and Ng, S. (2006). 'Confidence intervals for diffusion index forecasts and inference for factor-augmented regressions', *Econometrica*, Vol. 74, pp. 1133–1150.
Bai, J. and Ng, S. (2007). 'Determining the number of primitive shocks in factor models', *Journal of Business and Economic Statistics*, Vol. 25, pp. 52–60.
Bernanke, B. S., Boivin, J. and Eliasz, P. S. (2005). 'Measuring the effects of monetary policy: a Factor-Augmented Vector Autoregressive (FAVAR) approach', *The Quarterly Journal of Economics*, Vol. 120, pp. 387–422.
Beyer, A. (2009). *A Stable Model for Euro Area Money Demand: Revisiting the Role of Wealth*, Working Paper Series 1111, European Central Bank.
Boivin, J., Giannoni, M. and Mojon, B. (2009). 'How has the euro changed the monetary transmission mechanism?' in Acemoglu D., Rogoff K. and Woodford, M. (eds), *NBER Macroeconomics Annual 2008*, Chicago: University of Chicago Press, pp. 77–125.
Bosker, E. (2006). 'On the aggregation of eurozone data', *Economics Letters*, Vol. 90, pp. 260–265.
Breitung, J. and Eickmeier, S. (2011). 'Testing for structural breaks in dynamic factor models', *Journal of Econometrics*, Vol. 163, pp. 71–84.
Brown, R. L., Durbin, J. and Evans, J. M. (1975). 'Techniques for testing the constancy of regression relationships over time', *Journal of the Royal Statistical Society, Series B*, Vol. 37, pp. 149–192.
Canova, F. and de Nicolò, G. (2002). 'Monetary disturbances matter for business fluctuations in the G-7', *Journal of Monetary Economics*, Vol. 49, pp. 1131–1159.
Canova, F. and Pina, J. P. (2005). 'What VAR tell us about DSGE models', in Diebolt C. and Kyrtsou C. (eds), *New Trends in Macroeconomics*, New York: Springer Verlag, pp. 89–123.
Canova, F., Ciccarelli, M. and Ortega, E. (2012). 'Do institutional changes affect business cycles? Evidence from Europe', *Journal of Economic Dynamics and Control*, Vol. 36, pp. 1520–1533.
Carlstrom, C. T., Fuerst, T. S. and Paustian, M. (2009). 'Monetary policy shocks, Cholesky identification and DNK models', *Journal of Monetary Economics*, Vol. 56, pp. 1014–1021.
Castelnuovo, E. (2011). *Monetary Policy Shocks and Cholesky-VARs: An Assessment for the Euro Area*, mimeo, University of Padua.
Castelnuovo, E. (2012a). *In Cholesky–VARs We Trust? An Empirical Investigation for the U.S.*, mimeo, University of Padua.
Castelnuovo, E. (2012b). *Monetary Policy Neutrality? Sign Restrictions Go to Monte Carlo*, mimeo, University of Padua.
Cecioni, M. and Neri, S. (2011). *The Monetary Transmission Mechanism in the Euro Area: Has it Changed and Why?* Temi di discussione (Economic Working Papers) 808, Bank of Italy, Economic Research Department.
Chamberlain, G. and Rothschild, M. (1983). 'Arbitrage, factor structure, and mean-variance analysis on large asset markets', *Econometrica*, Vol. 51, pp. 1281–1304.
Christiano, L. J., Eichenbaum, M. and Evans, C. L. (1999). 'Monetary policy shocks: what have we learned and to what end?', in Taylor J. B. and Woodford M. (eds), *Handbook of Macroeconomics*, Vol. 1, Amsterdam: Elsevier, pp. 65–148.
Ciccarelli, M. and Rebucci, A. (2006). 'Has the transmission mechanism of European monetary policy changed in the run-up to EMU?' *European Economic Review*, Vol. 50, pp. 737–776.
Clarida, R., Galí, J. and Gertler, M. (1998). 'Monetary policy rules in practice: some international evidence', *European Economic Review*, Vol. 42, pp. 1033–1067.
Clements, B., Kontolemis, Z. G. and Levy, J. (2001). *Monetary Policy Under EMU: Differences in the Transmission Mechanism?* Working Paper 102, International Monetary Fund.
Doz, C., Giannone, D. and Reichlin, L. (2011). 'A two-step estimator for large approximate dynamic factor models based on Kalman filtering', *Journal of Econometrics*, Vol. 164, pp. 188–205.
Doz, C., Giannone, D. and Reichlin, L. (2012). 'A quasi maximum likelihood approach for large approximate dynamic factor models', *Review of Economics and Statistics*, Vol. 94, pp. 1014–1024.

Dreger, C. and Wolters, J. (2010). 'Investigating M3 money demand in the Euro Area', *Journal of International Money and Finance*, Vol. 29, pp. 111–122.
Durbin, J. (1969). 'Tests for serial correlation in regression analysis based on the periodogram of least squares residuals', *Biometrika*, Vol. 56, pp. 1–15.
Edgerton, D. and Wells, C. (1994). 'Critical values for the CUSUMSQ statistic in medium and large sized samples', *Oxford Bulletin of Economics and Statistics*, Vol. 54, pp. 355–365.
Eickmeier, S. (2009). 'Comovements and heterogeneity in the Euro Area analyzed in a non-stationary dynamic factor model', *Journal of Applied Econometrics*, Vol. 24, pp. 933–959.
Eickmeier, S. and Breitung, J. (2006). 'How synchronized are new EU member states with the Euro Area? Evidence from a structural factor model', *Journal of Comparative Economics*, Vol. 34, pp. 538–563.
Fagan, G., Henry, J. and Mestre, R. (2001). *An Area-wide Model (AWM) for the Euro Area*, Working Paper 42, European Central Bank.
Farrant, K. and Peersman, G. (2006). 'Is the exchange rate a shock absorber or a source of shocks? New empirical evidence', *Journal of Money, Credit and Banking*, Vol. 38, pp. 939–961.
Faust, J. (1998). 'The robustness of identified VAR conclusions about money', *Carnegie-Rochester Conference Series on Public Policy*, Vol. 49, pp. 207–244.
Favero, C. A., Marcellino, M. and Neglia, F. (2005). 'Principal components at work: the empirical analysis of monetary policy with large datasets', *Journal of Applied Econometrics*, Vol. 20, pp. 603–620.
Fielding, D. and Shields, K. (2011). 'Regional asymmetries in the impact of monetary policy shocks on prices: evidence from US cities', *Oxford Bulletin of Economics and Statistics*, Vol. 73, pp. 79–103.
Forni, M. and Gambetti, L. (2010a). 'The dynamic effects of monetary policy: a structural factor model approach', *Journal of Monetary Economics*, Vol. 57, pp. 203–216.
Forni, M. and Gambetti, L. (2010b). *Fiscal Foresight and the Effects of Government Spending*, Discussion Papers 7840, C.E.P.R.
Forni, M. and Gambetti, L. (2010c). *Macroeconomic Shocks and the Business Cycle: Evidence from a Structural Factor Model*, Discussion Papers 7692, C.E.P.R.
Forni, M. and Lippi, M. (2001). 'The generalized dynamic factor model: Representation theory', *Econometric Theory*, Vol. 17, pp. 1113–1141.
Forni, M., Hallin, M., Lippi, M. and Reichlin, L. (2000). 'The generalized dynamic factor model: identification and estimation', *The Review of Economics and Statistics*, Vol. 82, pp. 540–554.
Forni, M., Giannone, D., Lippi, M. and Reichlin, L. (2009). 'Opening the black box: structural factor models versus structural VARs', *Econometric Theory*, Vol. 25, pp. 1319–1347.
Fry, R. and Pagan, A. R. (2011). 'Sign restrictions in structural vector autoregressions: a critical review', *Journal of Economic Literature*, Vol. 49, pp. 938–960.
Geweke, J. (1977). 'The dynamic factor analysis of economic time series', in Aigner D. J. and Goldberger A. S. (eds), *Latent Variables in Socio-Economic Models*, Amsterdam: North Holland, pp. 365–383.
Giannone, D., Reichlin, L. and Sala, L. (2005). 'Monetary policy in real time', in Gertler M. and Rogoff K. (eds), *NBER Macroeconomics Annual 2004*, Cambridge, MA: MIT Press, pp. 161–200.
Güell, M. and Petrongolo, B. (2007). 'How binding are legal limits? Transitions from temporary to permanent work in Spain', *Labour Economics*, Vol. 14, pp. 153–183.
Hallin, M. and Liška, R. (2007). 'Determining the number of factors in the general dynamic factor model', *Journal of the American Statistical Association*, Vol. 102, pp. 603–617.
Kapetanios, G. (2010). 'A testing procedure for determining the number of factors in approximate factor models with large datasets', *Journal of Business and Economic Statistics*, Vol. 28, pp. 397–409.
Kilian, L. (1998). 'Small-sample confidence intervals for impulse response functions', *The Review of Economics and Statistics*, Vol. 80, pp. 218–230.
Knetsch, T. A. (2010). 'Trend and cycle features in German residential investment before and after reunification', in Bandt O., Knetsch T., Peñalosa J. and Zollino F. (eds), *Housing Markets in Europe*, Berlin, Heidelberg: Springer, pp. 187–211.
Luciani, M. (2013). 'Monetary policy and the housing market: a structural factor analysis', *Journal of Applied Econometrics*, forthcoming. Doi 10.1002/jae2318.
McCallum, A. and Smets, F. (2009). *'Real Wages and Monetary Policy Transmission in the Euro Area'*, mimeo, University of Michigan and ECB. Paper presented at ECB conference on Monetary policy transmission mechanism in the Euro Area in its first 10 years, Frankfurt, September 2009.

Mihov, I. (2001). 'Monetary policy implementation and transmission in the european monetary union', *Economic Policy*, Vol. 16, pp. 369–406.

Mojon, B. and Peersman, G. (2003). 'Monetary policy transmission in the Euro Area', in Angeloni I., Kashyiap A. K. and Mojon B. (eds), *Monetary Policy Transmission in the Euro Area*, Cambridge: Cambridge University Press, pp. 56–74.

Monticelli, C. and Tristani, O. (1999). *What Does the Single Monetary Policy do? A SVAR Benchmark for the European Central Bank*, Working Paper 2, European Central Bank.

Olivei, G. and Tenreyro, S. (2010). 'Wage-setting patterns and monetary policy: international evidence', *Journal of Monetary Economics*, Vol. 57, pp. 785–802.

Onatski, A. (2009). 'Testing hypotheses about the number of factors in large factor models', *Econometrica*, Vol. 77, pp. 1447–1479.

Onatski, A. (2010). 'Determining the number of factors from empirical distribution of eigenvalues', *The Review of Economics and Statistics*, Vol. 92, pp. 1004–1016.

Papademos, L. D. and Stark, J. (eds). (2010). *Enhancing Monetary Analysis*, European Central Bank Frankfurt am Main.

Paustian, M. (2007). 'Assessing sign restrictions', *The BE Journal of Macroeconomics*, Vol. 7, pp. 1–33.

Peersman, G. (2004). 'The transmission of monetary policy in the Euro Area: are the effects different across countries?' *Oxford Bulletin of Economics and Statistics*, Vol. 66, pp. 285–308.

Peersman, G. (2005). 'What caused the early millennium slowdown? Evidence based on vector autoregressions', *Journal of Applied Econometrics*, Vol. 20, pp. 185–207.

Peersman, G. and Smets, F. (2003). The monetary transmission mechanism in the Euro Area: more evidence from VAR analysis, in Angeloni I., Kashyiap A. K. and Mojon B. (eds), *Monetary Transmission Mechanism in the Euro Area*, Cambridge: Cambridge University Press, pp. 36–55.

Rafiq, M. S. and Mallick, S. K. (2008). 'The effects of monetary policy on output in EMU: a sign restrictions approach', *Journal of Macroeconomics*, Vol. 30, pp. 1756–1791.

Reichlin, L. (2009). Comment on how has the Euro changed the monetary transmission, in Acemoglu D., Rogoff K. and Woodford M. (eds), *NBER Macroeconomics Annual*, Chicago: University of Chicago Press, pp. 127–139.

Sala, L. (2003). *Monetary Transmission in the Euro Area: A Factor Model Approach*, mimeo, University Bocconi.

Sargent, T. J. and Sims, C. A. (1977). Business cycle modeling without pretending to have too much a-priori economic theory, in Sims C. A. (ed.), *New Methods in Business Cycle Research*, Minneapolis, MI: Federal Reserve Bank of Minneapolis, pp. 45–109.

Stock, J. H. and Watson, M. W. (2002). 'Forecasting using principal components from a large number of predictors', *Journal of the American Statistical Association*, Vol. 97, pp. 1167–1179.

Stock, J. H. and Watson, M. W. (2005). *Implications of Dynamic Factor Models for VAR Analysis*, Working Paper 11467, NBER.

Uhlig, H. (2005). 'What are the effects of monetary policy on output? Results from an agnostic identification procedure', *Journal of Monetary Economics*, Vol. 52, pp. 381–419.

Uhlig, H. (2009). Comment on how has the Euro changed the monetary transmission, in Acemoglu D., Rogoff K. and Woodford M. (eds), *NBER Macroeconomics Annual*, Chicago: University of Chicago Press, pp. 141–152.

Weber, A. A., Gerke, R. and Worms, A. (2011). 'Changes in Euro Area monetary transmission?' *Applied Financial Economics*, Vol. 21, pp. 131–145.