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CARLOS CUERPO
PILAR PONCELA

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ECONOMIC FORECASTING WITH MULTIVARIATE MODELS ALONG THE BUSINESS CYCLE

Carlos Cuerpo*

Pilar Poncela*

Abstract

The absence of an agreed model to forecast the main economic aggregates at different time horizons remains an important challenge for econometric analysis, especially in light of the recent subprime crisis. We conduct an out-of-sample forecasting performance exercise with structural (Dynamic Stochastic General Equilibrium) and non-structural (Dynamic Factor) models. We implement it over different stages of the business cycle to Spanish, euro area as well as US data over 1980Q1-2010Q4. Our results suggest accuracy gains from non-structural models in the short-run. Structural models perform better in the medium-run and their benefits increase at all time horizons when considering disruptive times.

Keywords: DSGE models; dynamic factor models; forecasting.

JEL classification: C32, C53, E32, E37.

Corresponding author: Carlos Cuerpo, Department of Economic Analysis: Quantitative Economics/ Fundamentos del Análisis Económico: Economía Cuantitativa. Universidad Autónoma de Madrid (UAM). 28049 Cantoblanco Madrid, despacho E-III 301A. E-mail: carlos.cuerpo@estudiante.uam.es

* Department of Economic Analysis: Quantitative Economics/ Fundamentos del Análisis Económico: Economía Cuantitativa. Universidad Autónoma de Madrid (UAM).

1. Introduction

Understanding the future of macroeconomic forecasting requires understanding the interplay between measurement and theory, and the corresponding evolution of the non-structural and structural approaches to forecasting. (Diebold, 1998)

The ability of economists to forecast the main aggregate macroeconomic variables has undergone a complete revolution in the last 40 years. The stagflation period that affected the developed economies during the late 70s, along with theoretical dissatisfactions with Keynesian principals shaped a new way of understanding forecasting and led to the pioneering contributions of Sargent and Sims (1977) and Sims (1980): non-structural models that minimized their theoretical roots. Non-structural models were not subject to changes in the dominant theoretical paradigm and could also escape the Lucas (1976) critique, as their forecasts were not tied to a specific path of the policy variables. Pivoting on Box and Jenkins (1970) contributions and Sims (1980) multivariate extension, Vector Autoregressive (VAR) models became the workhorse in macroeconomic forecasting. Indeed, Unrestricted VAR (UVAR) models do not impose excessive identification restrictions, leading to a better in-sample fit.

However, good in-sample fit did not grant a good out-of-sample forecasting performance, as indicated in Stock and Watson (1996) work. In order to avoid omitted variables biases and allow for the identification of structural shocks through the model innovations, researchers could have a tendency towards increasing the number of variables in the analysis. This strategy could generate inaccurate estimates and bad predictive results due to over-fitting, as the number of parameters to estimate increases with the square of the number of variables included in the model. VAR modeling limitations led thus to the parallel development of two major lines of research:

- The restriction of the parameter space by imposing Bayesian constraints through a priori information. The first Bayesian VAR (BVAR) models were based on purely statistical beliefs. For example, the famous ”Minnesota prior”, developed by Doan, Litterman and Sims (1984) and Litterman (1986), restricts the higher lags coefficients to near to zero values. The development of Monte Carlo type algorithms, such as the Gibbs-Sampler, developed by Geman and Geman (1984) or the Metropolis *et al*. (1953) algorithm, generalized in Hastings (1970), as well as their application to the field of economics have positioned BVAR models as benchmarks in forecasting major macroeconomic aggregates, as indicated in Zarnowitz and Braun (1994). For example, see Kinal and Ratner (1986) BVAR application to forecast New York data, extended by Sims (1992) for the U.S. economy or Amisano and Serati (2002) for the euro area.

- Sargent and Sims (1977) findings that a few shocks can explain most of the dynamics of macroeconomic aggregates set the scenario for Dynamic Factor Models (DFM). Following this evidence, Geweke (1977) assumes that all time series can be expressed as the sum of two orthogonal components: a common one, which captures the dynamic shared by all series, and an idiosyncratic component, understood as a residual. Information technologies developments and the availability of real-time data have facilitated the exploitation of DFM’s potential, as stated in Forni *et al*. (2005a and 2005b), Forni *et al*. (2009), and Stock and Watson (2002), among others, where the orthogonality assumption among the idiosyncratic components is relaxed and the number of series considered in the models can be very large.

Economic theory would not lag behind for a long time. Dissatisfaction with the theoretical basis of non-structural models and the need to generate forecasts conditional on economic policy as a guide to policymakers stimulated the birth of structural models, explicitly grounded in the optimizing behavior of economic agents. The so-called new classical macroeconomics school came to acknowledge the need for micro-founded, dynamic, stochastic macroeconomic models that could escape from the Lucas critique in policy guiding.

The contributions by Hansen and Sargent (1980) and Kydland and Prescott

(1982) started the literature of Dynamic Stochastic General Equilibrium (DSGE) models. They provided a connection between theory and data thanks to the state space representation of the decision rules obtained from the solution of the models. Since then, the parallel evolution of macroeconomic theory towards a widespread paradigm, the New Neoclassical Synthesis (see Goodfriend (2002) for an introduction or Gali and Gertler (2007) for a historical overview) and econometric techniques that ease the estimation of large-scale DSGE models, especially Bayesian econometrics techniques (see Greenberg (2007) for an introduction and Geweke (2005) and Canova (2007) for a deeper discussion) have provided useful tools and DSGE models have been adopted by many Central Banks as they serve a great variety of purposes such as policy guiding, historical analysis and counterfactual experiments.

However, as discussed in Sims (2002a), forecasting exercises have traditionally been backed mainly by large-scale macro-econometric models and expert judgment analysis, without an explicit theoretical structure or a consistent treatment of expectations. Interestingly, since the contribution of Smets and Wouters (2004), economists have begun to look seriously at structural DSGE models as effective tools in forecasting. These authors capture the statistical features of the main aggregate variables by applying Bayesian estimation techniques to large scale models. Ever since, the forecasting performance of DSGEs has been tested against traditional UVAR, BVAR benchmarks (see Smets and Wouters, 2005 for a closed economy application and Adolfson *et al.*, 2007 or Christofel *et al.*, 2010) for an open economy case), more sophisticated set-ups such as dynamic factor models à la Stock and Watson (2002) (see Wang, 2009) and even against experts judgment with real-time data (for example in Adolfson *et al.*, 2007, Edge, Kiley and LaForte, 2009, Kolasa *et al.*, 2009 and Rubaszek and Skrzypczynski, 2008).

Following this literature, DSGE forecasting ability is comparable to that of the competing models, especially in the medium to long-term horizon, as suggested in Monti (2008) and Wang (2009). These results have inspired new estimation and forecasting practices in an attempt to combine the best of the various methodologies, and minimize the issues resulting from poorly specified models.

Generically, the solution of a DSGE model can be cast into a VAR representation, which is used to assess and validate the DSGE empirically. Ingram and Witheman (1994) goes one step further and builds a BVAR model whose priors consist of a real business cycle model, away from the traditional statistical Minnesota prior, obtaining accuracy gains in out-of-sample forecasts of the main US aggregates. In this line, Del Negro and Schorfheide (2003, 2004) and Del negro *et al*. (2004) use a DSGE model to build the prior density function for a VAR (DSGE-VAR) and develop an estimation procedure to obtain the posterior distribution of the DSGE structural parameters by minimizing the divergence between the UVAR estimates and those from the VAR representation of the DSGE (Kullback-Leibler divergence). An application of this methodology can be found in Hodge *et al.* (2008) for the Australian economy. These authors have obtained competitive results in the prediction of GDP and inflation, both against non-structural models as well as purely theoretical models. Waggoner and Zha (2010) further extend the methodology through a combination between a BVAR and a DSGE model, assigning state-dependent probabilities to the two models (to their likelihood functions).

Giannone *et al.* (2006) criticize, however, the VAR representation of the DSGE model, as the presence of measurement errors in observed variables would contaminate the inference results, especially in the short term. Moreover, they find that observed variables usually follow a factor-type structure. They therefore decompose the spectral density of the model into two components, a common and an idiosyncratic one. The DSGE representation through a factor structure leads to higher inference accuracy and to the ability of handling high-frequency data in real-time. In this line, in a pioneering article, Boivin and Giannoni (2006) relax the assumption that DSGE theoretical concepts are measured adequately by a single series and estimate a structural model with a wide range of data, relating the factors of the resulting system with the endogenous variables of the DSGE model. Following this approach, Bäurle (2008) finds significant forecasting accuracy improvements at all time horizons. Schorfheide *et al.* (2010), simplify the Boivin and Giannoni (2006) approach by introducing a two steps estimation process which creates a link between the DSGE model variables and some other non-modeled variables that can therefore be forecast indirectly through auxiliary regressions.

The bulk of the literature restricts its attention to economies that are linearized around a steady state or long-run equilibrium, yielding approximate decision rules and likelihood functions, from which inferences and forecasts are constructed. The work of Fernández-Villaverde and Rubio-Ramírez (2005 and 2007) noted the importance of quadratic terms in the approximation to decision functions, since second-order errors may have first order effects on the likelihood function. The consequences of non-linearities in terms of forecasting have been evaluated by Pichler (2007), which assesses the trade-off between the sampling error introduced by the non-linear filters and the error due to the linear approximation of the model, which turns out to be more impeding in terms of out-of-sample forecasting.

Moreover, another related field of research seeks to incorporate the know-how or experts views into the DSGE framework. Monti (2008) and Giannone *et al.* (2009) successfully integrate the soft information or expert judgments of high frequency information in real time in a DSGE model, obtaining GDP forecasting gains, as monthly data is incorporated into the model.

Despite the growing literature on the predictive capabilities of DSGE models and their variants, there are a number of weaknesses, unexplored issues or inconclusive results. First, so far, the vast majority of forecast comparison exercises have focused on fairly stable business cycle periods (except for Waggoner and Zha, 2010). There is a general consensus regarding the misbehavior of model forecasts during the latest financial crisis, which has revealed the need to rethink the academic agenda. It is obvious that any economic disruption with respect to the past implies a great challenge in out-of-sample forecasting exercises. However, we might wonder to what extent these special circumstances will affect the different models. Second, the literature covering DSGE models is generally biased towards theoretical improvements and it is worth checking if forecasting gains follow these refinements. Enriched models allow a finer description of the existing transmission channels but might not prove to improve misspecified models' forecast accuracy.

Sticking to a linearized environment, we will conduct a comparative analysis of the forecasting performance of structural versus non-structural models, both in a smooth-growth period (2003Q1-2007Q3) and during a crisis (2006Q3-2010Q4) in order to overcome the first two issues. We will check the robustness of our results by applying the exercise to Spanish, euro area 16 as well as to US data. Our sample covers the period 1980Q1-2010Q4. The forecasting performance will be assessed with a recursive procedure, checking the out-of-sample RMSE from 1 to 8 periods.

Section 2 covers the main methodological issues while section 3 specifies the data collection and treatment as well as the model estimations results[1]. Section 4 presents the results and main findings and finally, section 5 concludes.

2. Methodology

2.1. Non-Structural Models

2.1.1. Reference models

In order to introduce the notation and establish the framework for comparison of the DSGE and DFM models, we first define UVAR and BVAR models.

Vector Autoregressive Models

We consider that all time series are generated from the linear combination of three elements: their own lagged values (dynamics), lagged values of the other variables (cross-dynamics) and innovations or specific shocks. A simplified system of order 1 could be specified in the following form:

$$\begin{bmatrix} z_{1,t} \\ z_{2,t} \\ \vdots \\ z_{k,t} \end{bmatrix} = \begin{bmatrix} \varphi_{1,1} & \varphi_{1,2} & \cdots & \varphi_{1,k} \\ \varphi_{2,1} & \varphi_{2,2} & \cdots & \varphi_{2,k} \\ \vdots & \vdots & \cdots & \vdots \\ \varphi_{k,1} & \varphi_{k,2} & \cdots & \varphi_{2,1} \end{bmatrix} \begin{bmatrix} z_{1,t-1} \\ z_{2,t-1} \\ \vdots \\ z_{k,t-1} \end{bmatrix} + \begin{bmatrix} u_{1,t} \\ u_{2,t} \\ \vdots \\ u_{k,t} \end{bmatrix}, \quad (1)$$

where $u_t = [u_{1,t} \cdots \quad u_{k,t}]' \sim N(0, \Sigma_u)$. Furthermore, to implement a univariate analysis we impose a diagonal structure on the matrix of parameters as well as on the variance covariance (VCV) matrix.

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[1] Detailed results on the model estimations and tests undertaken are available upon request

To allow for the estimation of the $z_t = [z_{1,t} \cdots \quad z_{k,t}]'$system via Kalman filtering techniques, we impose a "State Space" structure on it. By further generalizing for *p* lags, we obtain the transition equation,

$$\begin{bmatrix} z_t \\ z_{t-1} \\ \cdots \\ z_{t-p+1} \end{bmatrix} = \begin{bmatrix} \varphi_1 & \varphi_2 & \cdots & \varphi_{p-1} & \varphi_p \\ I & 0 & \cdots & 0 & 0 \\ \cdots & \cdots & \cdots & \cdots & \cdots \\ 0 & 0 & \cdots & I & 0 \end{bmatrix} \begin{bmatrix} z_{t-1} \\ z_{t-2} \\ \cdots \\ z_{t-p} \end{bmatrix} + \begin{bmatrix} u_t \\ 0 \\ \cdots \\ 0 \end{bmatrix}, \tag{2}$$

where now and φ_i is the kxk matrix of coefficients that relates z_t to its i-th lag, being *i* = 1, ..., *p*. In matrix notation, this transition equation describes the dynamics of the state vector Z_t,

$$Z_t = \Gamma Z_{t-1} + U_t, \tag{3}$$

where $Z_t = [z'_t \quad z'_{t-1} \dots \quad z'_{t-p}]'$, $U_t = [u'_t \quad 0' \dots \quad 0']'$and Γ is defined in an obvious way from (2).

By adding the corresponding measurement equation,

$$Y_t = \mathrm{H} Z_t + e_t, \tag{4}$$

with Y_t being the observables vector and assuming zero mean Gaussian errors and both noises U_t and e_t uncorrelated for all lags, the state space representation is complete and ready for estimation via the Kalman filter, which will evaluate its Likelihood Function (LF) and can therefore be used to estimate the unknown parameters $\zeta = (P \quad H \quad \Sigma_u \quad \Sigma_e)$ *via* maximum likelihood algorithms. Notice that, in order to estimate a pure VAR(p) model, $e_t = 0_{np,1}$ and $H = [I \quad 0 \quad 0]$, while adding measurement noise in equation (4) imposes a VARMA structure for the observed Y_t series.

As seen in the introduction, UVAR models can lead to overfitting problems that directly influence the forecasting results. In order to minimize this issue, we apply Bayesian estimation methods to VAR models.

Bayesian VAR models

The Bayesian approach allows us to combine beliefs with observed historical data to reduce UVAR dimensionality problems. By using a priori information (either statistical or economic), we provide the parameters with an initial value while specifying our trust on it at the same time, and observed data is then used as a device to review the prior beliefs.

We will take the ”Minnesota prior” developed by Doan, Litterman and Sims (1984) and Litterman (1986), as a reference point for the specification of the mean vector of the parameters. That is, all the variables of the system will follow a random walk with drift. However, as the original Minnesota prior was designed for non-stationary data, the prior was adapted following Lütkepohl (2005) whenever stationary series were used. The vector Z_t is made up of k variables and their corresponding p lags. The β coefficients associated to the mean can be cast as $\beta = vec(\varphi)$, with $\varphi = (\mu_t \quad \varphi_1 \quad \cdots \quad \varphi_p)$. The ”Litterman prior” will restrict β mean and variance covariance matrix $\beta \sim N(\beta^*, \mathrm{V}_\beta)$.

The prior density of the first moment of the coefficients will be equal to one for the first lag of all the coefficients and zero otherwise. The prior density of the second moment of the constant is considered as diffuse, leaving its estimation to the data. As for the rest of the coefficients of the variance covariance matrix, their `a priori will depend on a vector of hyperparameters $\Pi = (\pi_1 \quad \pi_2 \quad \pi_3)$, which will set three dimensions; the general dynamics, g that can follow a harmonic or a geometric decay process ruled by π_3, the first order own dynamics, with π_1 representing the trust on the prior density over the mean and finally the cross dynamics, where the out-of-sample specification of the dynamic interaction between the series will depend on π_2.

2.2. Dynamic Factor Models

Following Peña and Poncela (2004), we take an N-dimensional vector of observable series $\mathrm{Y_t}$. We assume that every time series can be written as a linear combination of r factors capturing the common dynamics and m specific components:

$$Y_t = \mathrm{P} f_t + n_t, \tag{5}$$

with f_t being a vector of common factors, of dimension r×1, P their loading matrix and n_t an N×1 vector of specific components. We assume that the common factors follow a VAR representation:

$$\Phi(B)f_t = a_t, \tag{6}$$

where $\Phi(B) = I - \Phi(1)B - \cdots - \Phi(p)B^p$ is an r×r polynomial matrix, B is the lag operator (By_t = y_{t-1}) and $a_t \sim N_r(0, \Sigma_a)$ and is serially uncorrelated. The model can be written in state space form as in the case of the VAR (see equations 3 and 4), with the s×1 state vector containing the common factors and their lags.

Forecasting is also done by applying Kalman equations to get first the state vector forecast h periods ahead and its VCV matrix and then use the measurement equation to get the observables forecasts altogether with their second moment.

2.3. Structural Models

2.3.1. DSGE Models

Once approximated by a log-linearization process around the steady state, DSGEs solution gives the laws of motion that can be cast into a state space form as the transition equation:

$$Z_t = f(Z_{t-1}, W_t; \theta), \tag{6}$$

where Z_t represents endogenous variables, W_t is the innovations vector and θ includes all structural parameters. In order to fulfill this process, we use Sims (2002b) algorithm[2]. This method requires a specific initial matrix representation of the model, such as

$$\Gamma_0 Z_t = \Gamma_1 Z_{t-1} + C + \psi V_t + \Pi\eta_t, \tag{7}$$

where V_t represents the structural innovations, while η_t introduces expectational errors. Its reduced-from representation is given by:

$$Z_t = \Theta_c + \Theta_0 Z_{t-1} + \Theta_1 W_t, \tag{8}$$

where the Θ_c and Θ_0 and Θ_1 matrices depend on the structural parameters and

[2] Which matlab version gensys.m is available at his personal website, http://www.princeton.edu/sims/.

summarize the dynamic behavior of the model, with a generic representation following equation (6). Moreover, the model estimation needs a measurement equation in order to complete the state space representation. Again, the observed variable Y_t will be a linear function of the state variables, with measurement errors e_t as in equation (4). If we assume Gaussian white noise perturbations, the Kalman Filter will evaluate the LF of our model, as for DFM and VAR models, by specifying the parameter vector to estimate ζ.

We will estimate the model using Bayesian techniques. The posterior density is made up of two components, the LF and a prior distribution over the structural parameters,

$$\mathrm{g}(\zeta|Y_t) = \mathrm{L}(Y_t|\zeta)\, g(\zeta), \tag{9}$$

and is obtained by Metropolis-Hastings numerical approximation methods, more specifically the Random Walk Metropolis (RWM) algorithm[3].

2.3.2. DSGE-VAR Models

Following Del Negro and Schorfheide (2004), we incorporate prior information from Smets and Wouters (2005) model to a VAR representation in order to proceed with the Bayesian estimation, by minimizing the divergence between an UVAR representation of the observable variables and the DSGE-VAR model [Kullback-Leibler discrepancy]. The *a priori* density function will have a hierarchical structure, as the DSGE model depends on unknown structural parameters. The density function will therefore be given by a marginal distribution of the θ parameters and a conditional distribution of the VAR (φ, Σ) parameters, given θ,

$$\mathrm{p}(\varphi, \Sigma, \theta) = \mathrm{p}(\varphi, \Sigma|\theta)\, p(\theta), \tag{10}$$

Symmetrically, the joint posterior density function will depend on the posterior density of the VAR parameters and the marginal posterior density of the θ,

$$\mathrm{p}(\varphi, \Sigma, \theta|Z) = \mathrm{p}(\varphi, \Sigma|Z, \theta)\, p(\theta|Z). \tag{11}$$

[3] See the Dynare website (http://www.dynare.org) for a more detailed explanation on the RWM approximation method used.

All in all, the DSGE-VAR representation implies many restrictions on the VAR parameters. In order to allow for possible misspecification problems, a new parameter λ rescales the VCV of the VAR and assesses the weight of the structural versus the non-structural component.

3. Data and estimation results

Both structural and non-structural models are estimated for Spain, the euro area and the US, using seven macroeconomic aggregates: Gross Domestic Product, GDP (Y), Private Consumption (C), Private Investment (I), Employment (L) or Total Hours worked (H), GDP deflator (Pr), Wages (W) and the interest rate (i). The data sample covers the period ranging from 1980:Q1 to 2010:Q4, thus including the latest subprime crisis. GDP and its components as well as wages are expressed in real terms and then divided by a population index to obtain per capita variables.

More specifically, for the US, in order to remain close to the original Smets and Wouters (2005) analysis, Real Gross Domestic Product, Nominal Personal Consumption Expenditures, Fixed Private Domestic Investment and the Implicit price deflator of GDP come from the US Department of Commerce - Bureau of Economic Analysis databank (BEA) while the Index of average weekly hours for the Nonfarm Business sector (NFB) and their Real Hourly compensation are taken from the Bureau of Labour Statistics (BLS). Moreover, hours are adjusted with a civilian employment index to take into account the limited coverage of the NFB sector. Table 1 and its graphical representation (figure 1) present more detailed information and the specific transformations of the US series for structural as well as non-structural models.

Table 1: US data, basic information

Acronym	Units	Transformation	Source
Y	B. of dollars	$\Delta(100*\ln(Y))$	BEA databank
C	B. of dollars	$\Delta(100*\ln(C/Pr))$	BEA databank
I	B. of dollars	$\Delta(100*\ln(I/Pr))$	BEA databank
H	Index (2005=100)	$\Delta(100*\ln(H))$	BLS databank
Pr	Index (2005=100)	$\Delta(100*\ln(Pr))$	BEA databank
W	Index (2005=100)	$\Delta(100*\ln(W/Pr))$	BLS databank
i[a]	Percentage	i/4	Federal Reserve

[a] The short-term interest rate is the Federal Funds rate

For the euro area, all time series are taken from the latest update of the Area Wide Model (AWM) database at the ECB, originally developed in *Fagan et al.* (2001). Private Consumption and Total Gross Investment are deflated with their own deflator. According to Smets and Wouters (2005), total employment data was used due to the lack of availability on hours worked. In order to compensate for this in the estimation of the DSGE model, an auxiliary equation will link labour services with observed employment following a hiring mechanism à la Calvo, as suggested in Smets and Wouters (2003).

Table 2 and its corresponding graph (figure 2) present more detailed information and the specific transformations of the series for structural as well as non-structural euro area models.

Table 2: Euro area data, basic information

Acronym	Units	Transformation	Source
Y	M. of ECU/euro	Δ(100*ln(Y))	QNA Eurostat
C	M. of ECU/euro	Δ(100*ln(C))	QNA Eurostat
I	M. of ECU/euro	Δ(100*ln(I))	QNA Eurostat
H	Thousands of	Δ(100*ln(H))	ECB Monthly Bulletin
Pr	Index (1955=100)	Δ(100*ln(Pr))	ECB Monthly Bulletin
W	M. of ECU/euro	Δ(100*ln(W/Pr))	ECB Monthly Bulletin
i[a]	Percentage	i/4	ECB Monthly Bulletin

[a] Three months interest rate

Lastly, data for the Spanish economy comes mainly from the National Statistics Institute (NSI) Quarterly National Accounts (QNA) except for the interest rate, provided by the Bank of Spain. Data on employment full-time equivalents require an auxiliary equation to the DSGE model, following the euro area specification. Table 3 and its accompanying graph (figure 3) provide a full description and representation of the Spanish data.

Table 3: Spain data, basic information

Acronym	Units	Transformation	Source
Y	M. of euro (2000)	Δ(100*ln(Y))	QNA INE
C	M. of euro (2000)	Δ(100*ln(C))	QNA INE
I	M. of euro (2000)	Δ(100*ln(I))	QNA INE
H	Thousands of	Δ(100*ln(H))	QNA INE
Pr	Index (2000=100)	Δ(100*ln(Pr))	QNA INE
W	M. of euro	Δ(100*ln(W/Pr))	QNA INE
i[a]	Percentage	i/4	Banco de España

[a] Non-transferable three month deposits

As can be seen in figures 1 to 3, stationarity is an issue for the employment, inflation and interest rate series. Their differenced versions are represented in dashed lines and used in the estimation of the non-structural models in order to obtain the best performing one in terms of forecasting accuracy[4]. Moreover, as the original Minnesota prior was designed for non-stationary data, the code was adapted following a procedure similar to Lütkepohl (2005) whenever stationary series were used. Four lags were selected for the VAR estimation following standard criteria and they were kept also for the Bayesian counterpart for consistency reasons.

[4] Detailed descriptions of the data and its transformations are available from the authors upon request.

Figure 1. US series

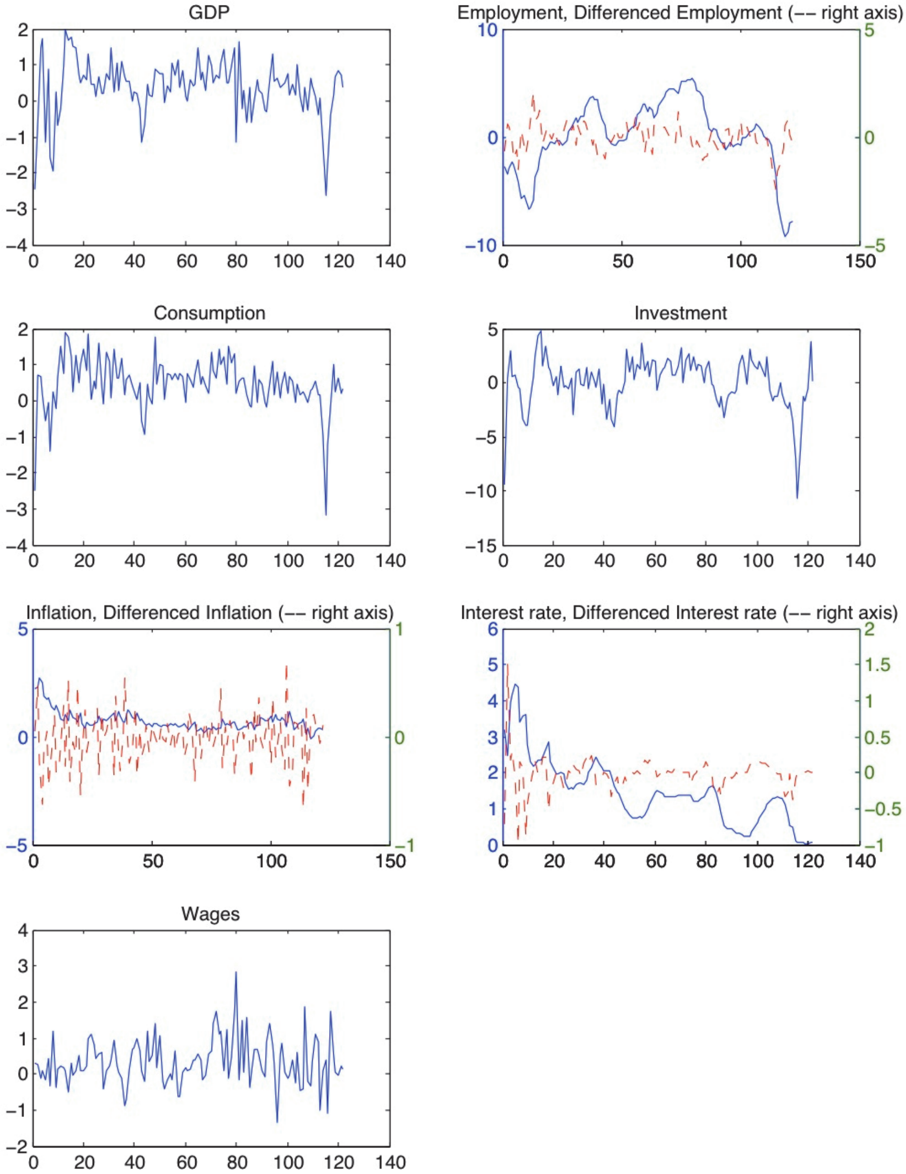


Figure 2. Euro area series

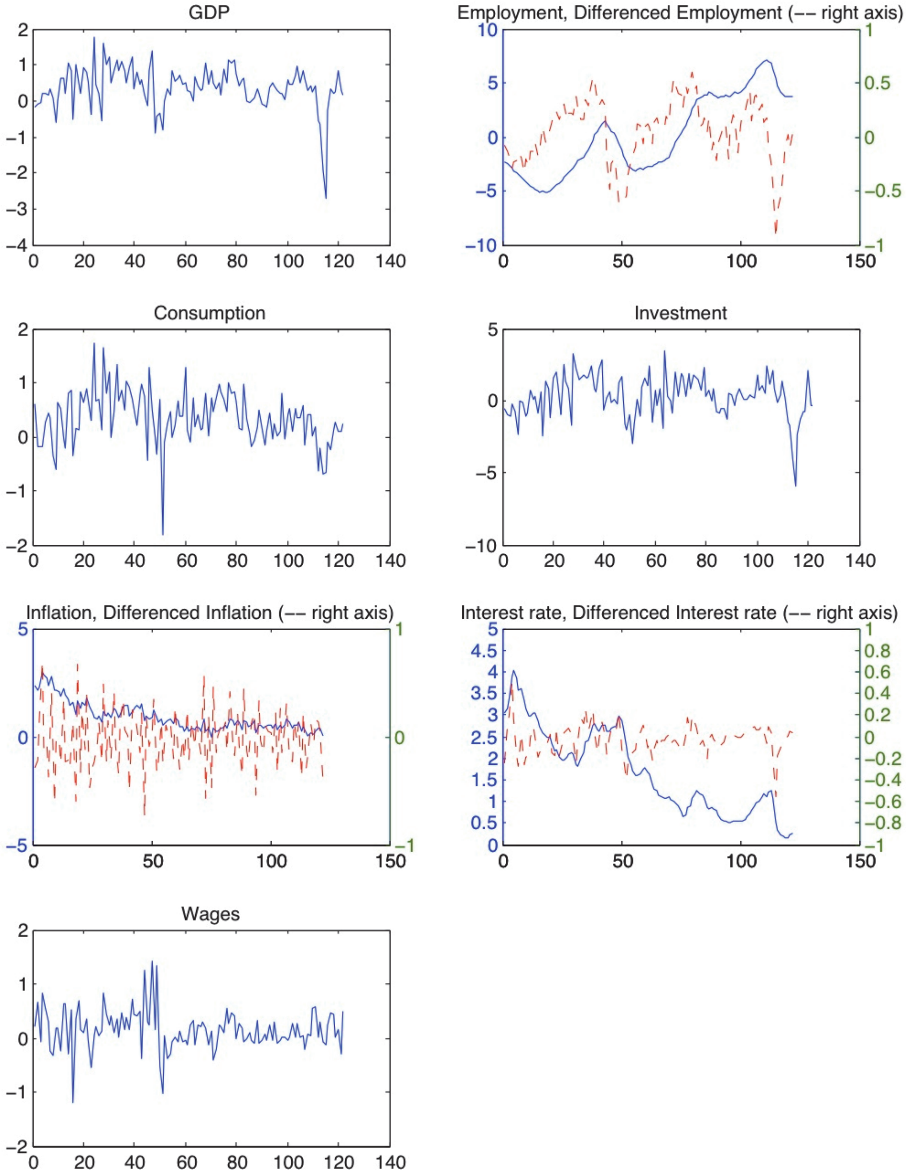
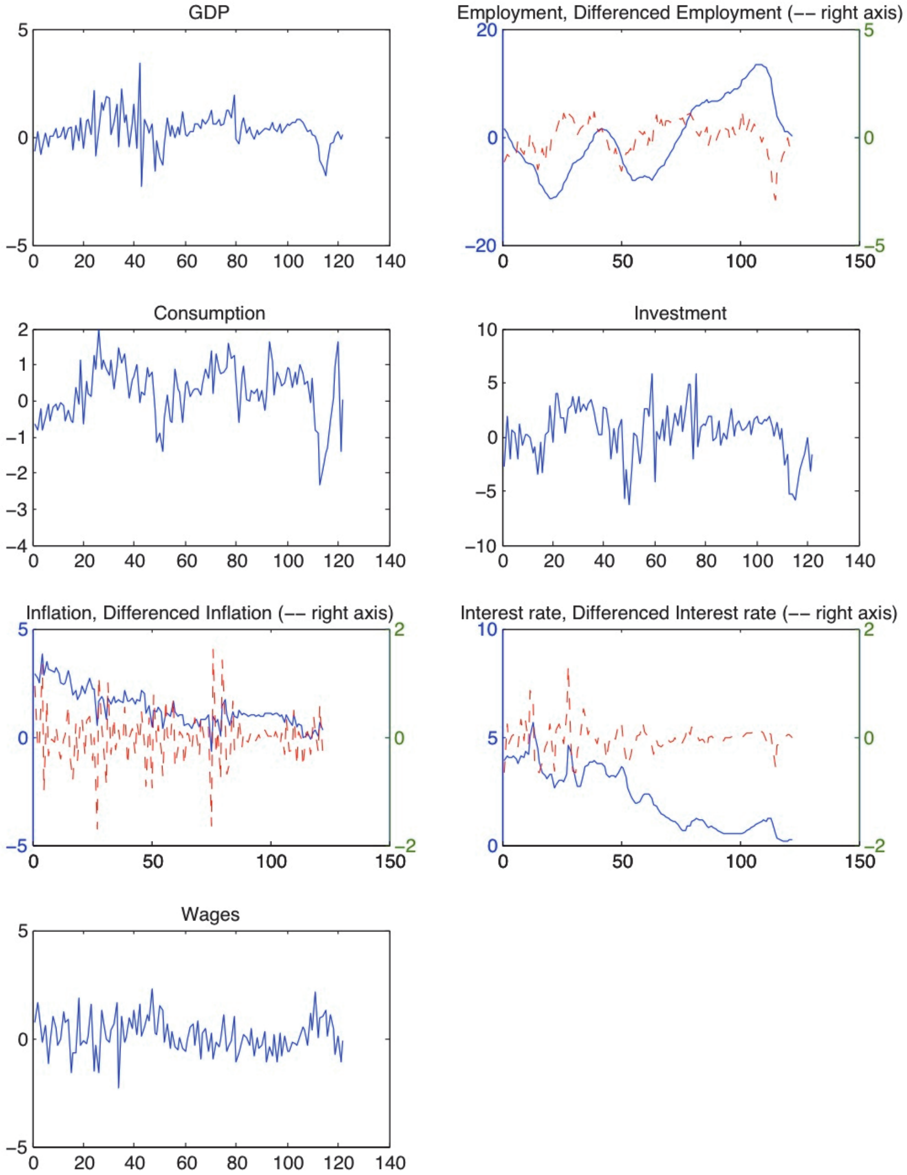


Figure 3. Spain series

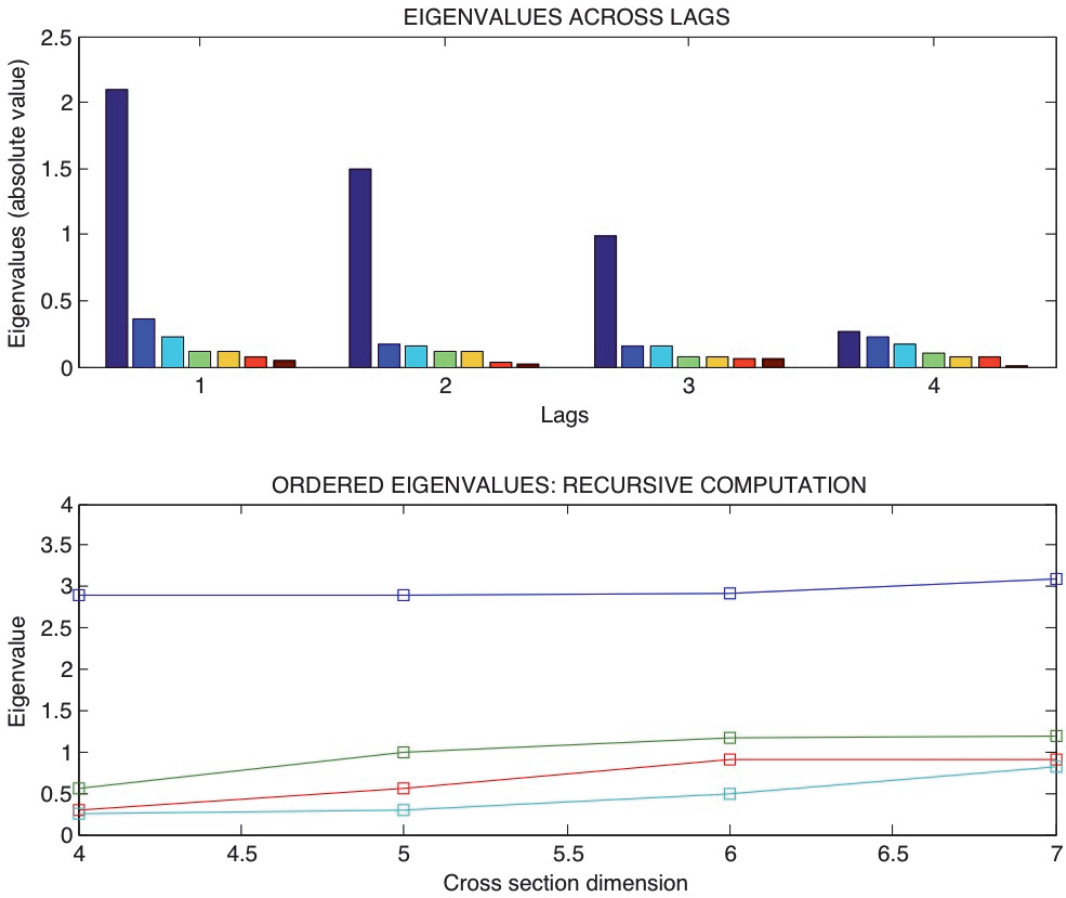


Turning to the first panel of figure 4, one eigenvalue of the cross-correlation matrix stands out when compared across lags for the US case. Moreover, following Peña and Box (1987) procedure, one factor was identified for each of the three countries as the corresponding stability requirements for the first associated eigenvector are met as well. Interestingly, performing Forni *et al.* (2005) methods, initially designed for large scale factor models, yields the same identification results as can be seen in the second panel of figure 4, with the eigenvalues computed recursively. By looking at the factor loadings, the extracted factor rests mainly on real series (GDP, consumption, investment and employment). However, nominal interest rate also stands as an important driver of this ”real” factor for the US and the euro area as can be seen in table 4. The impact of interest rate in the Spanish factor is much lower, possibly pointing to the importance of the ”one size fit all” monetary policy at the euro area level where the Spanish weight is around 10%. Real wages generally stand out with a lower loading in the common factor and interestingly, with a negative sign for Spain, confirming Messina *et al.* (2009) results on counterciclicality of the Spanish real wages, irrespectively of the deflator used. This real interpretation of the first common factor goes in line with Sargent and Sims (1977), who in a much larger set of series also assess that additional second and third factors share the nominal content of the GDP deflator, nominal wages and money supply variables.

Table 4: Factor loadings

Series	United	Euro	Spain
Y	0.75	0.77	0.85
C	0.86	0.89	0.85
I	0.90	0.91	0.77
H	0.85	0.78	0.90
Pr	0.14	0.09	0.10
W	-0.04	0.34	-0.31
I	0.51	0.52	0.26

Figure 4. DFM US identification results according to the cross-correlation matrix



The estimation of the structural DSGE-VAR sheds some light on possible misspecification issues. The posterior mode of the paramater λ for the US, the euro area and Spain is 1.66, 1.31 and 1.47, respectively. Following Consolo *et al.* (2009) we can define $\frac{\lambda}{1+\lambda}$ as the weight attached to the DSGE-generated data. The corresponding weights are 62%, 57% and 60%, confirming overall that the DSGE model restrictions are broadly supported by the data for the three countries. Although the US model is better specified as could be expected, the divergence between the three countries is minor despite their very different structures. The estimation of the Smets and Wouters (2005) model, tailored for a large-closed economy seems to work equivalently for the euro area and even for Spain, a small open economy without independent monetary policy.

4. Forecasting results

The design of the out-of-sample experimental forecasts follows a recursive procedure. Considering data until 2002Q4, we perform consecutive estimations up to 2010Q4, keeping one to eight steps out-of sample forecasts. We divide the forecasting period into two different samples of 18 data points for the one-step ahead forecasts, covering a smooth growth period [2003Q1-2007Q2] and a recession phase [2006Q3-2010Q4]. The forecasting performance can be dissected through four different dimensions: a time dimension (from one to eight quarters ahead), a contextual dimension (smooth growth period versus crisis period), a country-specific dimension (results for Spain, USA and the Euro area) and a model-specific dimension.

Abstracting first from model-specific aspects, figures 5 to 7 present RMSE results for the United States, the euro area and Spain, respectively. RMSEs are shown for the smooth as well as for the crisis period, signaling the first and third quartile, as well as the median and existing outliers. Several general conclusions appear to be robust across the different countries. First, as expected, the smooth growth period is characterized by smaller RMSEs at all horizons, especially for real variables. Second, the performance of the different models during the crisis worsens throughout the forecast horizon, contrary to the results for the stable period, with the exception of employment. Third, the relative dispersion of the results is bigger in the short-term for the stable period, confirming that bad forecasting performance is generalized during the crisis. Fourth, independently of the country considered, the dynamic factor model clearly outperforms the rest in terms of interest rate forecasts and the mixed DSGE-VAR obtains significantly worse results in predicting inflation and salaries This can be seen in the form of an ”outlier” in the box-plot diagrams. Fifth, the size of the RMSEs is similar across countries for the different variables, with a notable exception, employment, which exhibits abnormally large RMSEs for the Spanish and the European case, confirming possible misspecification problems.

Figure 5. US RMSE analysis, different forecast horizons (1 to 8 steps ahead)

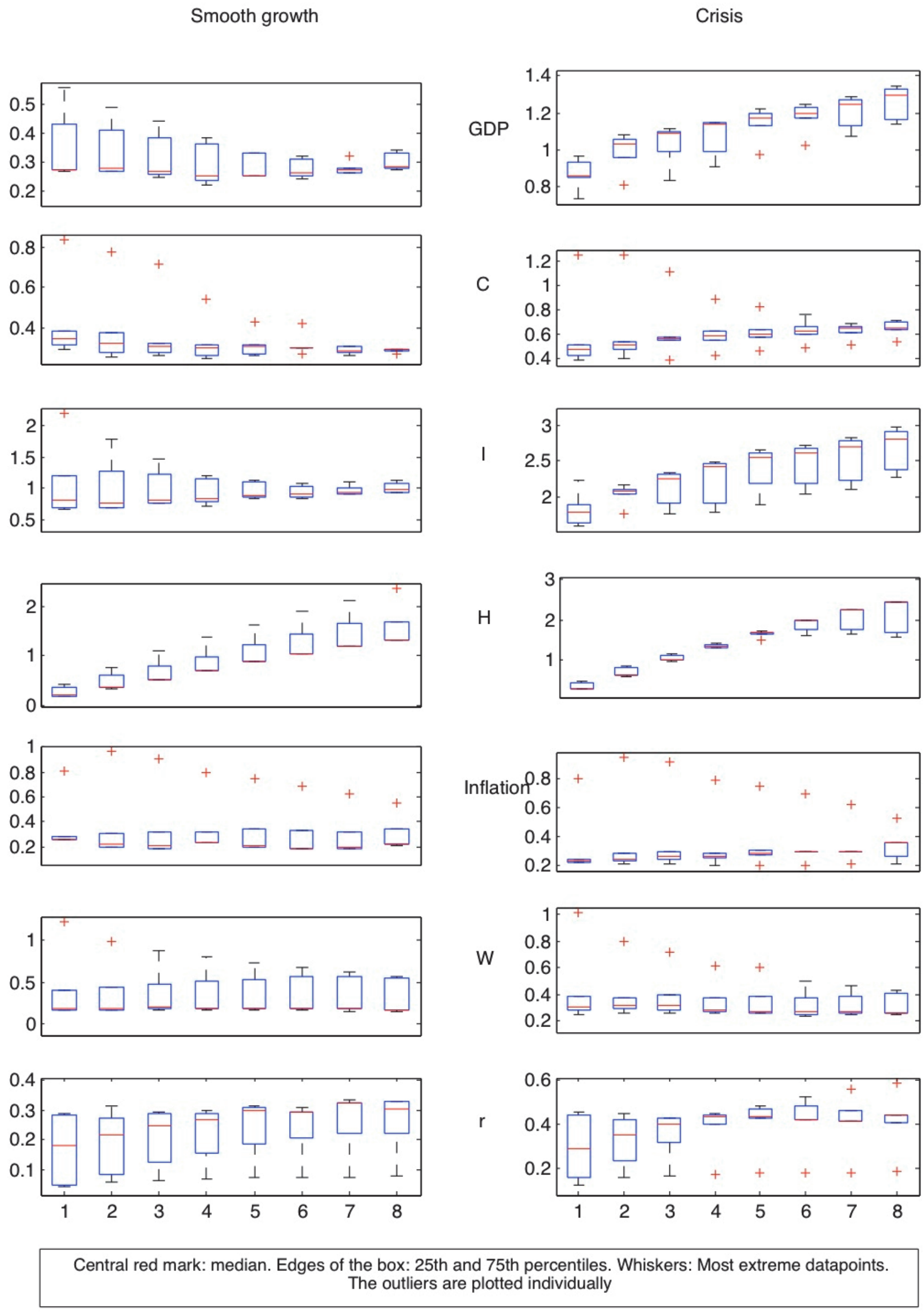


Figure 6. Euro area RMSE analysis, different forecast horizons (1 to 8 steps ahead)

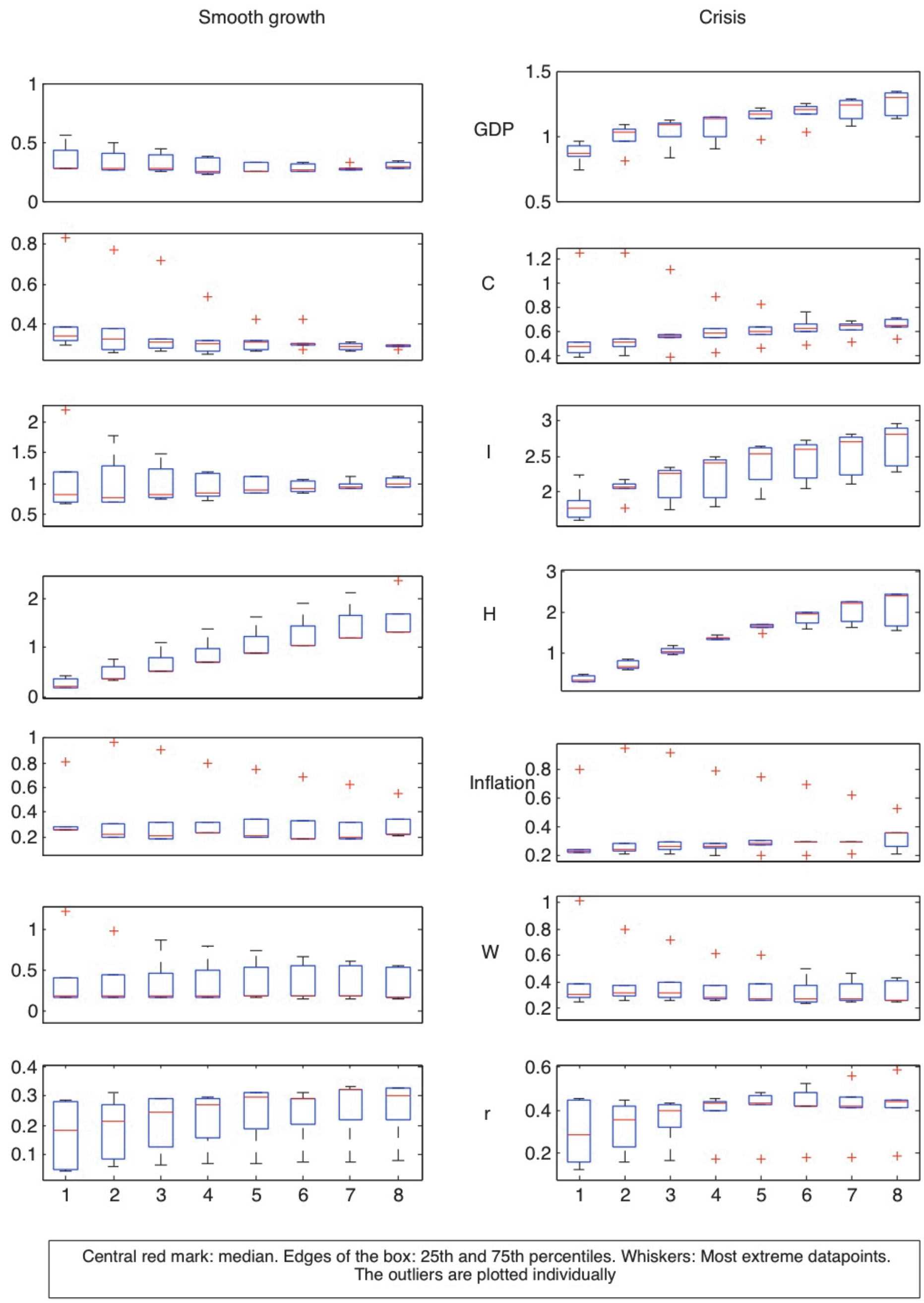
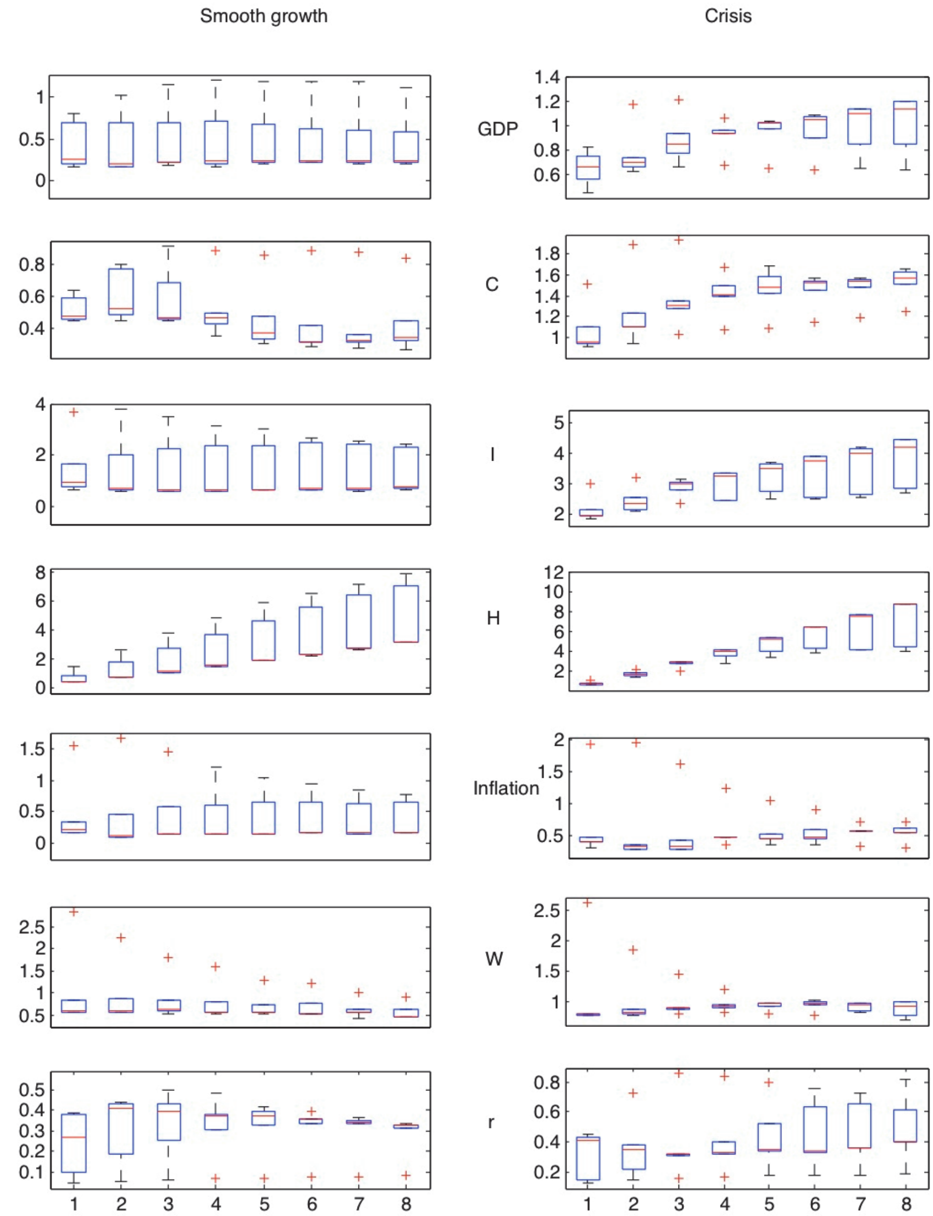


Figure 7. Spain RMSE analysis, different forecast horizons (1 to 8 steps ahead)



Central red mark: median. Edges of the box: 25th and 75th percentiles. Whiskers: Most extreme datapoints. The outliers are plotted individually

As stated during the methodological section, we will compare the predictive results from purely econometrical models [AR, VAR models as well as their bayesian estimated version with statistical priors] as well as structural models [DSGE], considering the DSGE-VAR as a natural, mixed benchmark. The DSGE-VAR framework presents a good fit to aggregated data and moreover retains the theoretical prescriptions from DSGE models. Figures 8 to 10 provide such a comparison across models, taking into account the time-dimension and also differences between countries. The percentage deviations of the different models with respect to the DSGE-VAR root mean square errors at the different horizons are represented. A negative value implies a gain in forecast accuracy and therefore lower RMSEs than the benchmark DSGE-VAR.

Independently of the country or the period considered, we observe accuracy gains from the DSGE-VAR along the time dimension. It follows that its relative forecasting results turn out to be better in the medium to long-run than in the short-run.

The performance of structural models during both periods is markedly better for the real variables, where they obtain relative predictive gains. Misspecification concerns arise when looking at employment forecasts during the smooth growth period, as its behavior is closer to nominal variables than to GDP, consumption and investment. Nominal variables, in turn, are dominated by the DFM model, which scores better for inflation, wages and interest rates across countries and periods, independently of the forecast horizon.

The relative performance of structural models improves during the crisis period for all countries and all variables. This is particularly striking for Spain, where DSGE fore- casting results are relatively poor during the smooth growth period. This finding could point out to possible misspecification problems of the Smets and Wouters (2005) model for the Spanish economy.

Figure 8. RMSE gap (%) with respect to the DSGE-VAR, US (1 to 8 steps ahead)

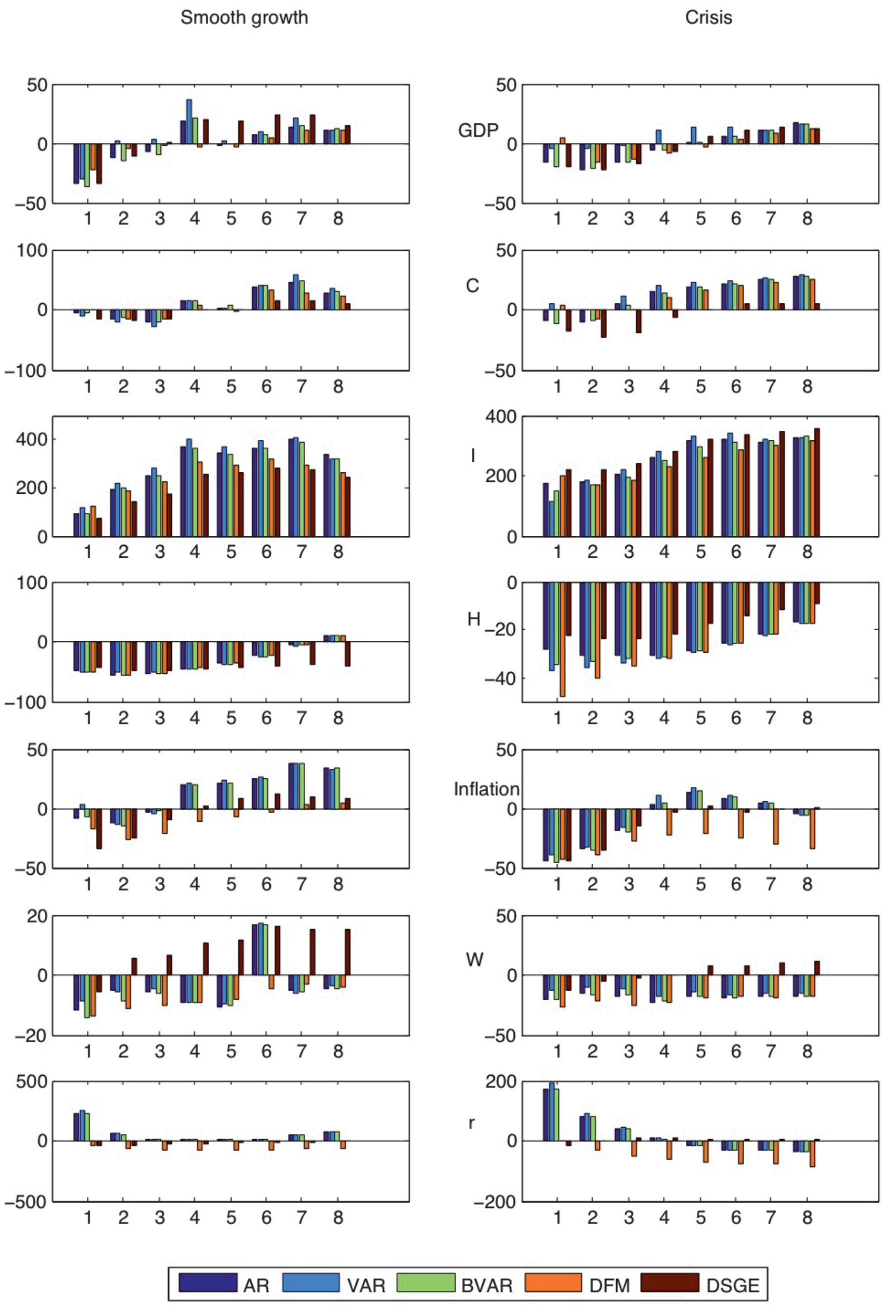


Figure 9. RMSE gap (%) with respect to the DSGE-VAR, euro area (1 to 8 steps ahead)

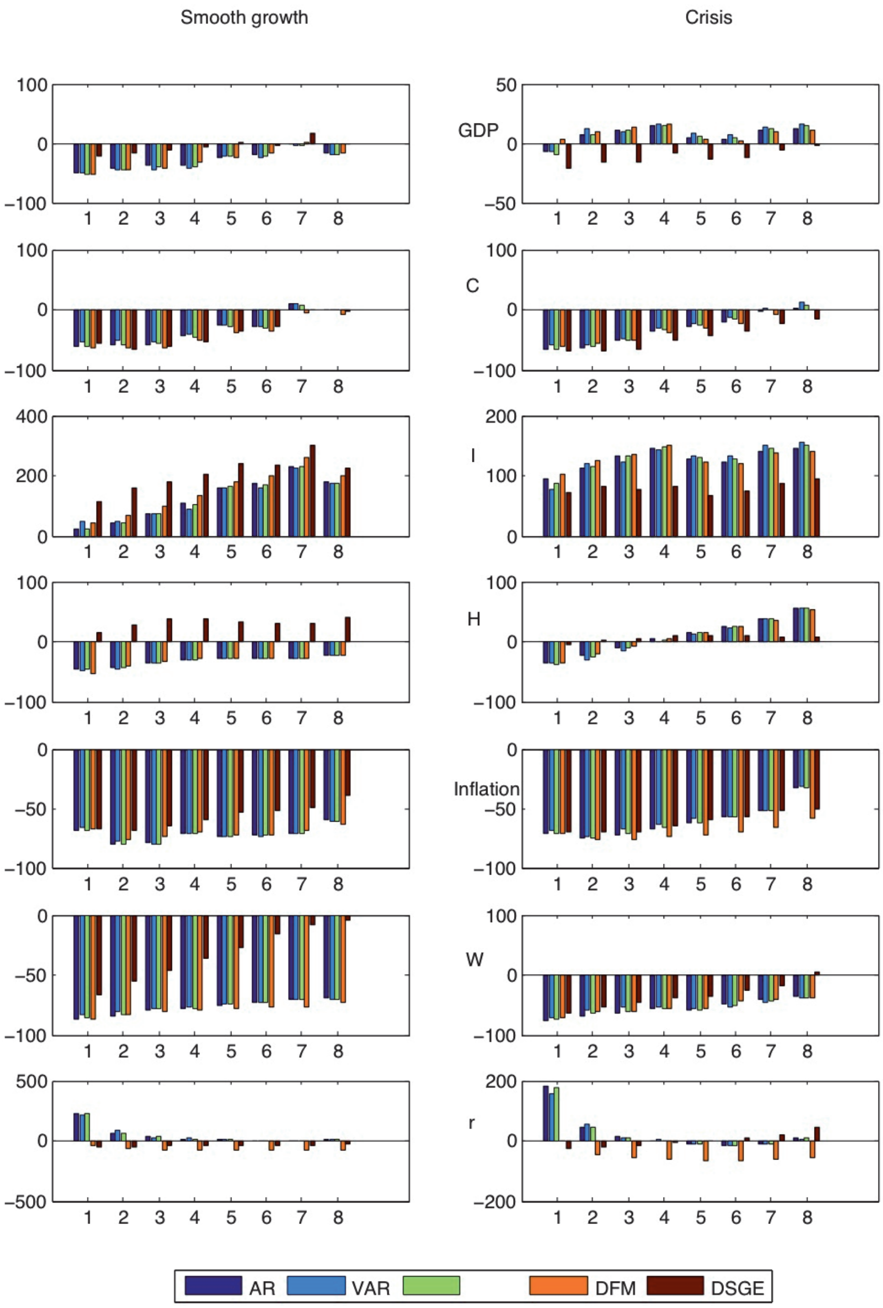
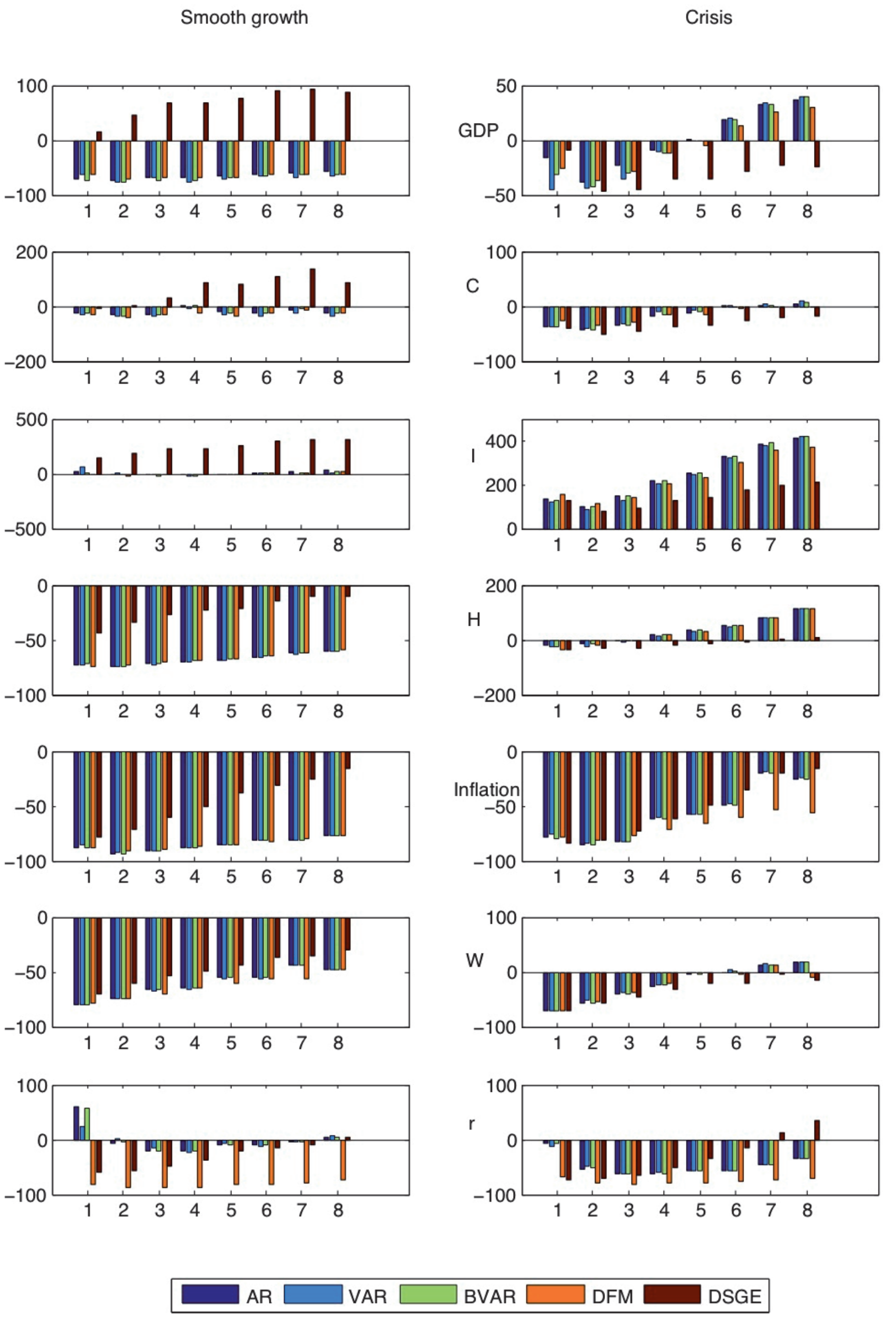


Figure 10. RMSE gap (%) with respect to the DSGE-VAR, Spain (1 to 8 steps ahead)



All in all, non-structural models perform relatively well during the smooth growth period at all forecast horizons, for all variables. This results would clearly point to the benefits of theoretical foundations for forecast accuracy in a period of higher volatility.

There seems to be a trade-off between considering stationary series in the non-structural models (with an extra difference in wages, inflation, interest rate and employment) and sticking to theoretically relevant concepts. The former strengthens the performance in the smooth growth periods and the short to medium term, while the latter is particularly relevant while facing turbulent times and high volatility in the economic cycle.

5. Concluding remarks

We conducted a comparative analysis of the out-of-sample forecasting performance of structural and non-structural models with quarterly data covering the period 1980Q1 to 2010Q4 for seven macroeconomic aggregates: Gross Domestic Product (GDP), private consumption, private investment, employment or total hours worked, the GDP deflator, real wages and the nominal interest rate. The forecasting performance was assessed using a recursive procedure through four different dimensions: a time dimension (from one to eight quarters ahead), a contextual dimension (smooth growth period [2003Q1-2007Q2] and recession phase [2006Q3-2010Q4]), a country-specific dimension (results for Spain, USA and the Euro area) and a model-specific dimension (comparison against traditional benchmarks such as VARs and BVARs).

All in all, there is supporting evidence for forecasting accuracy gains from structural models in the medium to long-run, while non-structural models perform generally better in the short-run. The benefits of structural models increase at all time horizons when considering disruptive times, with behavioral restrictions leading to higher parsimony. Indeed, during the ”Great Moderation” preceding the financial crisis, all models seemed to predict reasonably well at the different forecast horizons but this regularity was broken with the onset of the crisis as we observe increases in the relative performance of DSGE models for all countries and all variables. It would thus seem that forecasters should beware of too stable periods as the performance of the different models might not reflect their accuracy in terms of capturing the underlying economic developments but rather the regularity of the latter.

Moreover, Bayesian restrictions seem successful in shrinking the parameter space and providing better forecasting results. The results for the structural model, in turn, are robust across the different countries. Although the Smets and Wouters (2005) was initially tailored for the US economy (large and relatively closed), its forecasting gains with respect to non-structural models also apply to Spain and the euro area, despite their very different structures. It is thus natural at this stage to wonder whether misspecification issues have a sizable impact on the forecasting performance of these models. The latest theoretical refinements might not prove worth the effort when the ultimate goal is not better knowledge of the transmission channels of the different shocks but simply forecasting. A complementary line of research would mimic these results in a non-linearised environment, to check for the influence of the log-linearisation process in watering out modeling refinements.

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