

Duarte, Pablo; Süßmuth, Bernd

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## Implementing an approximate dynamic factor model to nowcast GDP using sensitivity analysis

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# Implementing an approximate dynamic factor model to nowcast GDP using sensitivity analysis

Pablo Duarte<sup>a</sup> and Bernd Süßmuth<sup>a, b</sup>

<sup>a</sup>*University of Leipzig*<sup>\*†</sup>

<sup>b</sup>*CESifo*

## Abstract

Dynamic factor models based on Kalman Filter techniques are frequently used to nowcast GDP. This study deals with the selection of indicators for this practice. We propose a two-tiered mechanism which is shown in a case study to produce more accurate nowcasts than a benchmark stochastic process and a standard model including extreme bounds fragile indicators. Nowcasting accuracy nearly measures up to the one of real-time forecasts by an institution with an interest in high-quality nowcasts.

JEL classification: C38, C53

Keywords: Dynamic factor, Kalman Filter, extreme bounds analysis

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<sup>\*</sup>Correspondence: University of Leipzig, Grimmaische Str. 12, D-04109 Leipzig, Germany  
Email contact: [suessmuth@wifa.uni-leipzig.de](mailto:suessmuth@wifa.uni-leipzig.de), [duarte@wifa.uni-leipzig.de](mailto:duarte@wifa.uni-leipzig.de)

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## 1 Introduction

GDP data is published quarterly with a lag of several weeks. Relying on relevant variables which are available earlier and at higher frequency, an “approximate” dynamic factor model (ADFM) can be used to deduce a common indicator to nowcast GDP. The outcome is a GDP series of higher frequency than its actual one (Aruoba et al., 2009, Camacho and Perez-Quiros, 2010, Camacho and Doménech, 2012, and Camacho and Garcia-Serrador, 2014).<sup>1</sup>

A central aspect of this class of small-scale nowcasting models is the pre-selection of indicators for the ADFM.<sup>2</sup> For example, Camacho and Perez-Quiros (2010, p. 672-674) suggest a selection procedure based on “successive enlargements.” Due to ordering in their screening out, they do not treat candidate variables symmetrically. Our contribution is to overcome this deficiency by proposing to subject the selection of indicators to an extreme bounds analysis (EBA); see e.g. Levine and Renelt (1992). The selection of indicators is illustrated for a sample case in two steps. First, we reduce a set of 258 monthly time series related to Spanish GDP to 27 (14) by keeping only the series showing a reasonable correlation with GDP, both contemporaneously and at its quarterly and yearly lags. In a second step, we apply an EBA procedure to drop “non-robust” indicators. To identify robust indicators, the strategy is to consider all possible regressions with quarterly GDP as dependent that can be estimated by taking combinations of a parsimonious subset (*micro-scale*) of pre-selected series (*small-scale*). The sensitivity analysis then consists in checking for each series whether the resulting distribution of parameter estimates has only positive or negative support, that is, whether the respective coefficient estimate does not change its sign in all regression runs. The latter should, at least, hold in an interval of so-called “extreme bounds” (Levine and Renelt, 1992).

We interpret and apply EBA as shrinkage device. Moreover, we prefer EBA to other more timely shrinkage devices such as the least absolute shrinkage and selection operator (LASSO) and ridge regressions or subset regressions based methods (see, e.g., Elliott et al., 2013) as for the latter the focus is on forecast performance, whereas it is on qualitative characteristics of series in the EBA case. EBA selects more economically meaningful indicator series into the micro-scale ADFM by assuring that they do not change, for instance, from procyclical to countercyclical (or from coinciding to either leading or lagging) indicators depending on covariate subset compositions

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<sup>1</sup>In practice, an ADFM modifies the exact dynamic factor model by Stock and Watson (1991) to account for problems of different frequency and asynchronous publication of series underlying the real-time forecast in applying a Kalman Filter strategy to fill up the series. For a more exact distinction see, e.g., Luciani (2014). Notably, Giannone et al. (2008), technically, were the first to develop the nowcasting methodology in large scale nowcasting models.

<sup>2</sup>We refer to small-scale as diffusion index forecasts (Stock and Watson, 2002) based on rather small two-digit numbers sets of indicators.

during shrinkage.

Against this backdrop, our central research questions can be summarized as follows:

1. Does it pay to consider a symmetric shrinkage device in the pre-selection process?
2. Does *robust* parsimony improve nowcasts, even if we rely on a rather ad hoc pre-selection device and a shrinkage device with focus on selecting economically meaningful indicators rather than on pure forecast performance in going from small-scale to micro-scale ADFM?

## 2 Approximate Dynamic Factor Model

The general idea behind an ADFM is that each series of a set can be written as sum of two stochastic components: a *common* factor  $f_t$ , which e.g. reflects the overall business cycle and an *idiosyncratic* component  $u_t$  which captures specific characteristics of each series. In general it reads

$$y_t = \mathbb{L}f_t + u_t, \quad (1)$$

where  $y_t$  is a  $K \times 1$  vector of variables,  $f_t$  an  $N \times 1$  vector of common factors with  $N < K$ , and  $\mathbb{L}$  a  $K \times N$  matrix of *factor loadings*. Furthermore,

$$f_t = A_1 f_{t-1} + \dots + A_p f_{t-p} + \eta_t; \quad u_t = C_1 u_{t-1} + \dots + C_q u_{t-q} + \varepsilon_t, \quad (2)$$

where  $\eta_t$  and  $\varepsilon_t$  represent white noise processes. To be estimated are matrices  $\mathbb{L}$ ,  $A_i$ , and  $C_i$  for  $i = 1, \dots, p$ . Growth rate  $\gamma$  then is forecasted by

$$\gamma_{t+h} = \mu + a(L)\gamma_t + b(L)\hat{f}_t + e_{t+h}, \quad (3)$$

where  $a(L)$  and  $b(L)$  denote lag polynomials in  $L$  of order  $p_a$  and  $p_b$ , i.e.  $a_0 + a_1 L + \dots + a_{p_a} L^{p_a}$  and  $b_0 + b_1 L + \dots + a_{p_b} L^{p_b}$  (Breitung and Eickmeier, 2006). An ADFM comprises three steps: (i) computation of the Gaussian likelihood using the Kalman Filter, (ii) estimation of  $\mathbb{L}$ ,  $A_i$  and  $C_i$ , and (iii) use of the Kalman Filter smoother to obtain estimates of  $f_t$ .

## 3 Implementation and in-sample analysis

Spanish GDP growth is stationary and available in quarterly frequency between 1975 and 2011 (solid line in Figure 1). A group of indicators ( $y_t$ ) driven by the same common factor as the growth rates of GDP must be selected to calculate the factor  $\hat{f}_t$ .

### 3.1 Implementation

Initially, we consider 258 variables related to Spanish GDP. All variables are available in monthly frequency and cover, at least, 10 years during the period at stake. All non-deseasonalized series were seasonally adjusted. Time series which were not stationary in levels were log first differenced. All monthly rates are tested stationary.

**Choice of core-set of indicators.** We first consider the correlation coefficients ( $\rho$ ) of each series and its quarterly and yearly lags with GDP growth (henceforth, denoted by  $\gamma_t$  representing first differences of the natural log of real GDP). To qualify it is required that  $|\rho| > .5$ . In our sample 27 series meet the requirement.

**Dynamic extreme bounds analysis.** In a second step we apply a non-standard EBA<sup>3</sup> to check if the remaining  $K = 27$  candidate indicators are “robust” in explaining GDP growth, i.e. if they are stable in their relationship with GDP growth. To this end, for each of the  $K$  candidate variables  $c$  a total number of  $J$  of the following regressions are estimated

$$\gamma_t = \delta_j + \beta_{c_j} c_t + \sum_{i=1}^k \beta_{x_{i,j}}(L) x_{t,i,j} + \varepsilon_{t,j}, \quad (4)$$

where  $x_{i,j}$  denotes a vector of up to three ( $k = 3$ ) additional indicator candidates, acting as randomly assigned conditioners in each of the  $J$  regression models. The order  $p_{x_\beta}$  of the polynomial  $\beta_{x_{i,j}}(L)$  in  $L$  is chosen on the basis of information criteria AIC and BIC for each  $k$ , respectively. The maximum lag order is set to 12. Note, if a variable is chosen in  $\geq 1$  lagged expression for the  $x$ -set, it is also considered both in contemporaneous form and in all approved lagged expressions for  $c_t$ . This is why we refer to our EBA version as “dynamic” EBA (DEBA). Hence, the minimum number of models to be checked  $J^{\min}$  is

$$J^{\min} = \binom{K-1}{k} = \frac{(K-1)!}{k!(K-1-k)!} \quad (5)$$

In our case ( $K - 1 = 26$ ;  $k = 3$ ),  $J^{\min} = 2,600$  and a total number  $K \times J^{\min} = 27 \times 2,600 = 70,200$  is to be estimated disregarding  $L$ . Considering  $L$ , we actually estimate a total of  $K \times J = 309,960$  models. For each model  $j$ , we obtain estimates  $\hat{\beta}_{c_j}$  and  $\hat{\sigma}_{\hat{\beta}_{c_j}}$  distributed across regressions  $j = 1, \dots, J$ . Next an upper and lower extreme

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<sup>3</sup>If standard EBA is defined by the seminal approaches advocated in Levine and Renelt (1992), in Sala-i-Martin (1997), sometimes referred to as “modified EBA”, and in Hoover and Perez (2004, p. 774-775), our approach is a hybrid version of all three of them. Following Levine and Renelt (1992), we consider one focus variable (in our notation  $c$ ). Like Sala-i-Martin (1997) we take into account *exactly* (and not *at most* as in Levine and Renelt, 1992) three “doubtful” variables in the conditioning set. And, as in Hoover and Perez (2004), we abstract from any other fixed subset of series (or, in general, variables) that we always include as covariates. However, we stick with the “strict” robustness definition of Levine and Renelt (1992) and also do not likelihood weight estimates.

bound ( $EB$ ) for each candidate indicator  $c$  and for each model  $j$  is defined:

$$\widehat{\beta}_{c_j} \pm 2\widehat{\sigma}_{\widehat{\beta}_{c_j}}. \quad (6)$$

A candidate indicator is said to be robust if for all models  $J$

$$\text{sign}(EB_L) = \text{sign}\left(\widehat{\beta}_{c_j} - 2\widehat{\sigma}_{\widehat{\beta}_{c_j}}\right) = \text{sign}(EB_U) = \text{sign}\left(\widehat{\beta}_{c_j} + 2\widehat{\sigma}_{\widehat{\beta}_{c_j}}\right); \quad (7)$$

in contrast, it is identified as fragile if

$$\text{sign}(EB_L) = \text{sign}\left(\widehat{\beta}_{c_j} - 2\widehat{\sigma}_{\widehat{\beta}_{c_j}}\right) < 0 < \text{sign}(EB_U) = \text{sign}\left(\widehat{\beta}_{c_j} + 2\widehat{\sigma}_{\widehat{\beta}_{c_j}}\right). \quad (8)$$

The intuition behind applying EBA is straightforward: whether a candidate indicator  $c_t$  behaves procyclically or countercyclically, i.e. is positively or negatively related with  $\gamma_t$ , is a defining property. This property should not change by including other indicator series as covariates, i.e. should not depend on whether  $\gamma_t$  is subjected to a randomly drawn set of further indicators in contemporaneous or contemporaneous and lagged expressions. In our sample case, we identify 14 variables as robust (Table 1).

1. INDUSTRY: MANUFACTURING - PRODUCTION TREND RECENT MONTHS
2. INDUSTRY: MANUFACTURING - ORDER BOOKS
3. INDUSTRY: MANUFACTURING - INDUSTRIAL CONFIDENCE
4. CONSUMER SURVEY: ECONOMIC SITUATION NEXT 12 MTH.
5. CONSUMER SURVEY: FINANCIAL SITUATION NEXT 12 MTH.
6. INDUSTRY SURVEY: PROD.EXPECTATION FOR MTH. AHEAD
7. INDUSTRY SURVEY: ORDER BOOK POSITION
8. INDUSTRIAL CONFIDENCE INDICATOR
9. INDUSTRY SURVEY: PROD. TRENDS IN RECENT MTH.
10. WORKERS ON LABOR REGISTER AFFILIATED TO SOCIAL SECURITY (First Differences)
11. INDUSTRIAL PRODUCTION - CONSUMER GOODS (First Differences)
12. INDUSTRIAL PRODUCTION - CAPITAL GOODS (First Differences)
13. INDUSTRIAL PRODUCTION - INTERMEDIATE GOODS (First Differences)
14. MADRID S.E - REAL ESTATE & FINANCE (First Differences)

Table 1: Selected variables to calculate the common leading factor

To sum up and to benchmark EBA against another widely used shrinkage device, consider the general LASSO objective

$$\min_{\beta \in \mathbb{R}^k} \|\gamma_t - X_t \beta\|^2 + \lambda_N \cdot \text{pen}(\beta) \text{ with } \text{pen}(\beta) = \sum_{j=1}^k \frac{|\beta_j|}{|\tilde{\beta}_j|},$$

where  $X_t$  comprises all potential indicator candidates or “target determinants” of  $\gamma_t$ . The LASSO can be interpreted as performing “all-in” penalized least squares (LS) estimations of linear models dropping “seemingly small” target determinants, whereas the EBA routine performs “all-in” LS estimations of linear models dropping “sign switching” target determinants. Economically, sign switching target determinants of  $\gamma_t$  are implausible indicator series that switch from procyclical (coinciding) to countercyclical (leading or lagging) indicators or the other way around depending on the subset composition of conditioners considered in the estimations.

**Statistical and economical properties of EBA.** Sensitivity analysis in the form of EBA has been mainly, although not exclusively (see, e.g. Sturm et al., 2005, and Moosa, 2009, analyzing IMF credit and FDI decisions, respectively), used in the context of model uncertainty with regard to cross-sectional growth regressions. This is also where the Bayesian foundations of this method are rooted (Leamer, 1982, 1983). It has been very rarely applied in the context of time series so far. Indeed, the only exceptions we are aware of are Fowles and Loeb (1995) and, more recently, Gassebner et al. (2016). Both studies are not dynamic in the above defined sense of DEBA, in that both do not consider lagged expressions of variables, neither in the conditioners set nor as focus indicators. In general, following Leamer (1982) and Fowles and Loeb (1995), EBA can be developed as an application of a Bayesian natural conjugate prior on the set of “doubtful” indicator-variables. For the normal linear regression model

$$\mathbf{Y} \sim \mathbf{N}(\mathbf{X}\beta, \sigma^2 \mathbf{I}),$$

the prior mean on the  $p$  doubtful variables is also normal (representing the crucial characteristic of a conjugate prior), centered at zero, with variance matrix  $\mathbf{H}^{*-1}$ . Hence,

$$\mathbf{R}\beta \sim \mathbf{N}(\mathbf{0}, \mathbf{H}^{*-1}),$$

where  $\mathbf{R}$  is a  $p \times k$  matrix of constants,  $\beta$  a  $k \times 1$  column vector of parameters, and  $\mathbf{0}$  a  $p \times 1$  zero vector, and  $\mathbf{H}^*$  is denoting a  $p \times p$  positive definite symmetric precision matrix (i.e. the inverse of the variance covariance matrix). As shown in Leamer (1982), extreme values of linear functions of the posterior mean for the  $k \times 1$  vector  $\tau$ ,

$$\tau \mathbf{b}^{**} = \tau' (\mathbf{H} + \mathbf{R}' \mathbf{H}^* \mathbf{R})^{-1} \mathbf{H} \mathbf{b},$$

are given by

$$\mathbf{a} + \tau^{*'} \mathbf{f} \pm \sqrt{\tau^{*'} \mathbf{A}^{-1} \tau} \mathbf{c}$$



when  $\mathbf{H}^*$  is constrained to fall between positive definite matrices  $\mathbf{V}\mathbf{I}$  and  $\mathbf{V}\mathbf{h}$ , and where

$$\begin{aligned}
\mathbf{a} &= \tau' \mathbf{b} - \tau' \mathbf{H}^{-1} \mathbf{R}' (\mathbf{R} \mathbf{H}^{-1} \mathbf{R}')^{-1} \mathbf{R} \mathbf{b}, \\
\tau^{*/'} &= \tau' \mathbf{H}^{-1} \mathbf{R}' (\mathbf{R} \mathbf{H}^{-1} \mathbf{R}')^{-1} \\
\mathbf{f} &= (\mathbf{h} + \mathbf{V}\mathbf{I}^{-1})^{-1} \left( \mathbf{h} \mathbf{R} \mathbf{b} + (\mathbf{V}\mathbf{I}^{-1} - \mathbf{V}\mathbf{h}^{-1})^* (\mathbf{h} + \mathbf{V}\mathbf{h}^{-1})^{-1} \mathbf{h} \mathbf{R} \mathbf{b} \frac{1}{2} \right), \\
\mathbf{A} &= (\mathbf{h} + \mathbf{V}\mathbf{h}^{-1}) (\mathbf{V}\mathbf{I}^{-1} - \mathbf{V}\mathbf{h}^{-1})^{-1} (\mathbf{h} + \mathbf{V}\mathbf{h}^{-1}) + (\mathbf{h} + \mathbf{V}\mathbf{h}^{-1}), \\
\mathbf{c} &= (\mathbf{R} \mathbf{b})' \mathbf{h} (\mathbf{h} + \mathbf{V}\mathbf{h}^{-1})^{-1} (\mathbf{V}\mathbf{I}^{-1} - \mathbf{V}\mathbf{h}^{-1}) (\mathbf{h} + \mathbf{V}\mathbf{I}^{-1})^{-1} \mathbf{h} \mathbf{R} \mathbf{b} \frac{1}{4}, \\
\mathbf{h} &= (\mathbf{R} \mathbf{H}^{-1} \mathbf{R}')^{-1}, \mathbf{b} = (\mathbf{X}' \mathbf{X})^{-1} \mathbf{X}' \mathbf{Y}, \mathbf{H} = s^{-2} \mathbf{X}' \mathbf{X}, \text{ and} \\
s &= \sqrt{\frac{1}{n-k} (\mathbf{Y} - \mathbf{X} \mathbf{b})' (\mathbf{Y} - \mathbf{X} \mathbf{b})}.
\end{aligned}$$

As Fowles and Loeb (1995, p. 363) note, the widest possible bounds occur at  $\mathbf{V}\mathbf{I} = \mathbf{0}\mathbf{H}^{*-1}$  and  $\mathbf{V}\mathbf{h} = \infty\mathbf{H}^{*-1}$ . As shown in Chamberlain and Leamer (1976) and Leamer (1982), priors are minimal since  $\mathbf{H}^*$ , i.e. the prior precision matrix, is only required to be positive definite symmetric. Matrix  $\mathbf{R}$  reduces the dimensionality of the prior space from  $k$  to  $p$ .

Apart from these desirable properties, (D)EBA has essentially three drawbacks. First, irrespective of its dynamic (i.e. lag) structure, the model size is quite restrictive (e.g. allowing only triplets of indicators as conditioning set). Secondly, models are equally weighted in the shrinkage procedure. And finally, standard EBA can be assessed as very or even “excessively” conservative; see Durlauf et al. (2009, p. 1132-1134). The standard cross-sectional and non-dynamic version of EBA has been shown in simulation studies by Hoover and Perez (2004) to be of very low empirical size (showing an overly low proportion of type I error), that is, the “algorithm almost never selects a variable that does not belong” (Hoover and Perez, 2004, p. 777), at the cost of a relatively low power. However, this trade-off needs not to apply to the same extent in the context of DEBA<sup>4</sup> and needs to be put into perspective in the context of using sensitivity analysis in the forecasting context as shrinkage device. Relatively high type II errors might hamper the precision of forecasts as “incorrect” indicators might survive the shrinking. On the other hand, the picking-out of economically meaningful (either coinciding or leading/lagging) indicator series is warranted with an extremely high probability. See the discussion in the preceding paragraph on DEBA.

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<sup>4</sup>It remains for future work to apply the numerical strategy used by Hoover and Perez (2004) in the cross-sectional growth regression context also in the (D)EBA time series context. Similarly, growing computational power will allow to alleviate the first caveat in our list.

### 3.2 In-sample performance

In Figure 1 the estimated common factor is plotted against the Spanish quarterly GDP growth series. The factor loadings are all estimated as significantly different from zero at standard levels. The estimated common factor  $\hat{f}_t$  is the final ingredient to determine the adequate lag structure of the forecast equation

$$\gamma_{t+h} = \mu + \alpha_0 \gamma_t + \alpha_3 \gamma_{t-3} + \beta_0 \hat{f}_t + \beta_1 \hat{f}_{t-1} + \beta_2 \hat{f}_{t-2} + e_{t+h}. \quad (9)$$

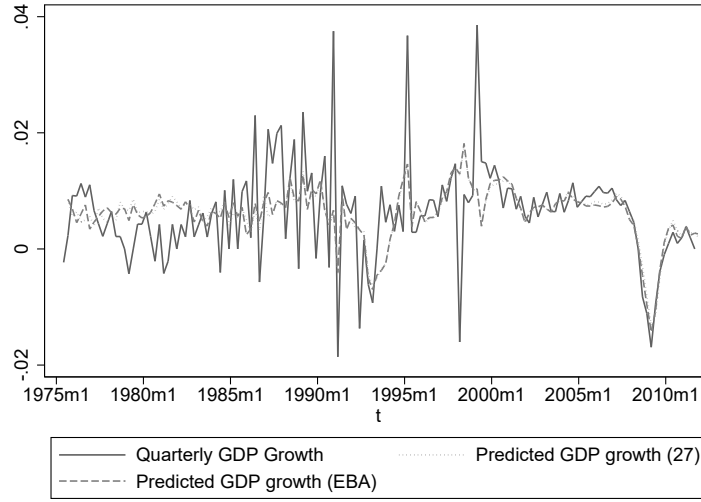


Figure 1: Predicted GDP growth –  $\gamma_{t+h} = \mu + a(L)\gamma_t + b(L)\hat{f}_t + e_{t+h}$

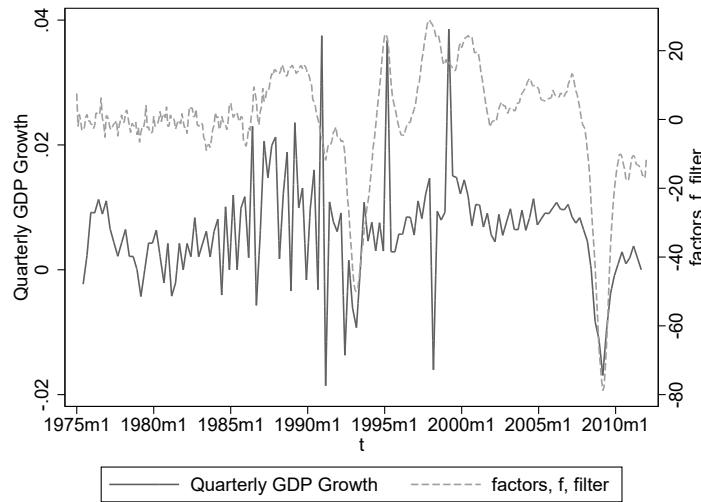


Figure 2: Estimated common factor  $\hat{f}$  and quarterly GDP growth

Our lag choice is based on information criteria.

Figure 2 shows observed and accordingly predicted (dashed line) values of Spanish GDP growth at quarterly frequency since 1975.

## 4 Out-of-sample performance

To examine and compare the out-of-sample characteristics of the two indices based on the core indicators on the one hand and on the selected candidates from the core set after screening out fragile indicators using EBA techniques on the other, we perform a “pseudo” real-time nowcasting exercise. It is “pseudo” (in contrast to “genuine”) because we are considering a method (ADFM) that was not available at the time of the nowcasts we are simulating. In our exercise, the focus is to see how our method would have performed in case it would have been available.<sup>5</sup>

In our exercise we simulate nowcasts for every quarter starting in 2007:Q1 until 2011:Q3. To simulate our pseudo real-time nowcasting procedure we assume that the information from the first quarter of 2007 onwards is not available at the moment of selecting the variables to be used to calculate the ADFM. This way we overcome a possible hindsight-bias regarding the selection of the variables for our procedure.

Additionally, we use GDP vintage data to simulate the nowcasting procedure with the information that one would have had at the time of the nowcast. Due to multiple data revisions the estimations of GDP are usually revised even some years after the quarter had passed. Figure 3 shows, for example, the GDP growth series nowcasters had access to in the first quarter of 2007 and the actual GDP growth series for the time between December 2000 (2000m12) and December 2006 (2006m12). Even if both series follow a similar pattern, which they obviously should, there are evident misalignments between the time series during the plotted period of time. This makes clear that even past values can change at the time of doing the nowcast and –by this circumstance– also the estimation results. For that reason we estimate our model every quarter anew using the (vintage) GDP growth series that was available at that quarter. The vintage as well as the final series are from the OECD.

Using only the information available until the end of 2006, the core set of indicators includes 14 (plus some lags). After applying the DEBA procedure, 9 variables are identified as robust. Interestingly, without the information from year 2007 afterwards, our selection procedure emphasizes financial variables such as a household consumer credit index –which does not qualify when including the information until 2011– and a real estate index.

As a benchmark to compare the performance of our nowcasting procedure we use two indicators. The first is a simple AR(2) process. An AR(2) –that is justified on

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<sup>5</sup>See Müller and Köberl (2012) for a genuine ex-ante nowcasting experiment.

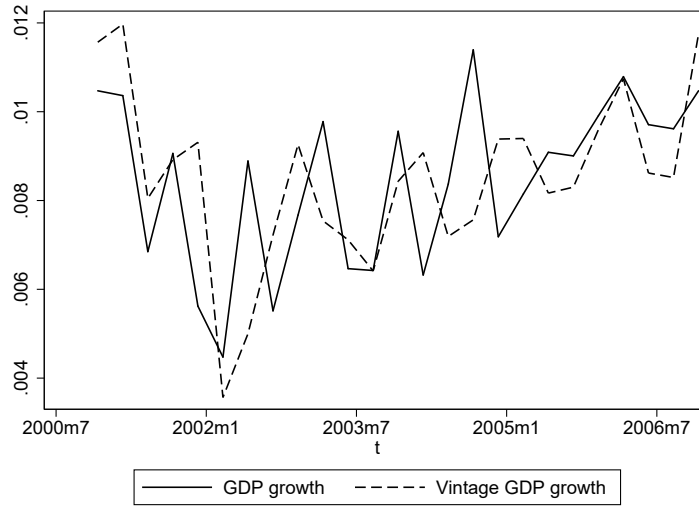


Figure 3: GDP growth vs vintage GDP growth

the basis of information criteria in terms of its order in our case study of Spanish GDP growth as well– is a very frequently used benchmark forecasting process in the context of forecasting real GDP growth. Chauvet and Potter (2013, p. 156) quote an exemplary series of four papers from the 2010s, where an AR(2) is employed in such a way, and state that a “simple linear univariate AR(2) offers an interesting comparison with the more complicated models as it has been shown in several papers to have comparable or better forecasting performance.” In addition, we went through the “Economic Forecast Reports” from the European Commission (EC) and collected the estimation for the current quarter and treated it as its nowcast. For example, if the report was published in April 2007, we treated the therein reported forecast of 2007:Q1 as the nowcast. Since the reports are only published biannually (once in spring and once in autumn), we took the values for the first two quarters of the year from the spring report and the last two quarters from the autumn report. The idea of using the EC official forecasts is to compare our model to a real-time forecast done by an institution with a clear interest in accurately nowcasting Spanish GDP.<sup>6</sup>

We proceed widely in line with the applied literature (see, e.g., Castle et al., 2009) in judging the out-of-sample forecast potential of the derived common factor. In order to assess the one-period ahead forecast properties, we treat the period starting the second quarter of 1975 to the last quarter of 2006 as given and successively extend this window by an additional quarterly observation to make one-quarter ahead projections for each successive step of prolonging the in-sample-“window” up to the penultimate quarter (using the vintage series of each quarter). Hence, our first projection uses equation (9)

<sup>6</sup>We are thankful to one of the anonymous referees to whom we owe this suggestion to referring to actual institutional forecasts as benchmark.

	Without EBA	With EBA	AR(2)	EC
Total period (averaged 1-step FEs)				
SFE	0.00006576	0.00001988	0.00006356	0.00000967
Relative SFE	0.8465	0.8236	–	
FE	-0.00386447	-0.00223941	-0.00484215	-0.00118044
Relative FE	0.798089692	0.462482575	–	
Excl. 2008:Q3 to 2009:Q2				
SFE	0.00004445	0.00001978	0.00002459	0.000005214
Relative SFE	1.807645384	0.804392029	–	
FE	-0.00240069	-0.00194073	-0.00285718	-0.00045629
Relative FE	0.840230577	0.67924667	–	

Table 2: Forecast errors across models

with  $h = 1$ ,  $t = 1$  (1975:Q2), ...,  $T_1$  (2006:Q4), and (justified by information criteria)  $p_a = p_b = 2$  to make a respective forecast of the GDP growth rate  $\gamma_{T_1+1}^F$ . We repeat this procedure for  $T_2$  (2007:Q1) to  $T_{34}$  (2011:Q2) in order to obtain a series of one-step ahead forecasts  $\gamma_{T_1+1}^F, \dots, \gamma_{T_{34}+1}^F$ . For the univariate benchmark AR(2) predictor  $a(L)' = (1 \ 0 \ 0)$ , while  $b(L)' = (0 \ 0 \ 0)$ .

From  $\gamma_{T_1+1}^F, \dots, \gamma_{T_{34}+1}^F$  we can calculate forecast errors (FE) for  $\hat{f}_t$  derived without EBA, i.e. FE (14), and with EBA, i.e. FE (EBA), respectively. Results along with the FE of the benchmark AR(2) are shown in the upper part of Figure 4. The dashed line refers to the EBA-robustly implemented DFM derived index based on 7 constituting indicators, the solid line to the one based on 14 constituting indicators, and the dotted one to the benchmark AR(2) predictor, respectively. Both DFM-based models generate lower forecast errors than the benchmark AR(2). In particular, this holds for times of crises. The error measure spike at the beginning of the great recession and at its onset is the lowest for the EBA-robust DFM index. Putting a stronger weight on larger deviations of forecasts from actual values, we obtain a similar picture for the squared forecast errors (SFE) as shown in the lower part of Figure 4.

Figures of these forecast errors averaged for different periods are summarized in Table 2, where “relative” errors refer to ratios of the respective errors to the corresponding errors of the AR(2) benchmark model. As can be seen from Table 2, while the nowcasting accuracy of the AR(2) measures up to the one of the small-scale ADFM without sensitivity analysis shrinkage (or even tops it in SFE-terms in case of excluding the crisis period), the real-time forecasts of the micro-scale ADFM implemented applying (D)EBA outperform both of them, i.e. AR(2) and small-scale ADFM, throughout and nearly measure up to the ones by the EC.

The common factor extraction including an EBA in the final pre-selection stage

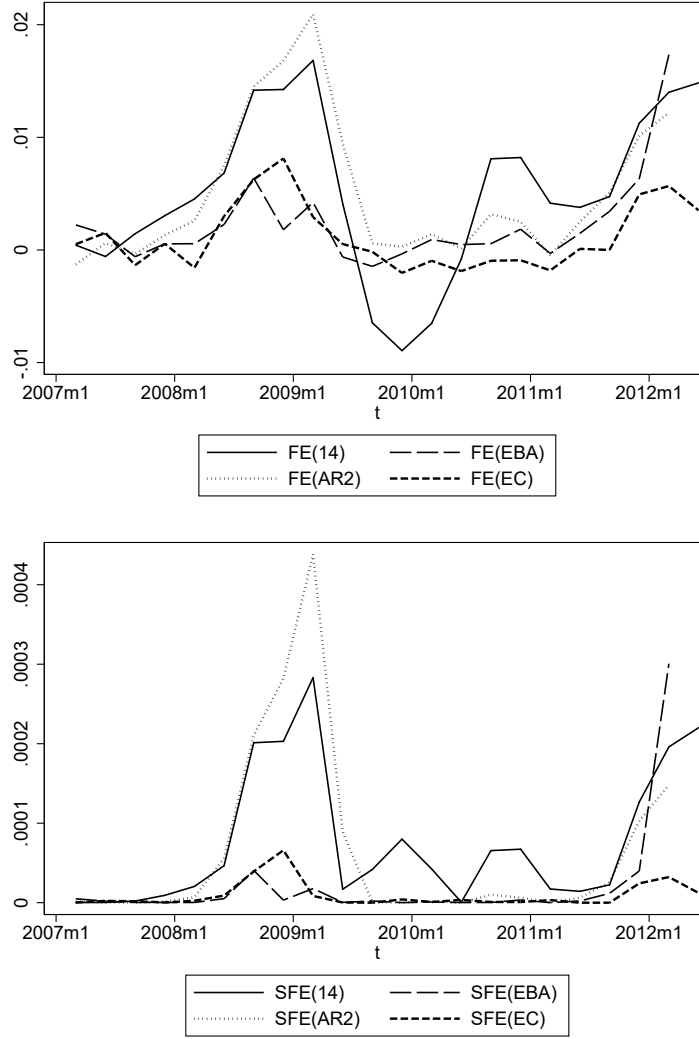


Figure 4: Forecast Error (FE) and Squared Forecast Error (SFE) at each predicted period, based on rolling window starting first quarter 2007

	Diebold-Mariano Test			Modified Diebold-Mariano Test		
	w/o EBA	with EBA	AR(2)	w/o EBA	with EBA	AR(2)
with EBA	-1.94 (0.052)			-1.719 (0.050)		
AR(2)	-0.178 (0.858)	1.13 (0.259)		-0.158 (0.438)	1.001 (0.164)	
EC	-2.879 (0.004)	-0.799 (0.424)	-1.587 (0.112)	-2.551 (0.009)	-0.708 (0.243)	-1.407 (0.087)

Table 3: Forecast accuracy tests between models

outperforms the one excluding it across all considered measures and periods shown in Table 2. We consider for the averaging of forecast errors the total forecast period from  $T_1$  (2007:Q1) to  $T_{34}$  (2011:Q2) as well as a period that leaves out the sub-period from the third quarter of 2008 to the second quarter of 2009, that is, the period when the shock corresponding to the crisis hit the economy.<sup>7</sup> Our results remain mainly untouched. Throughout forecast errors are the lowest for the common factor including the EBA exercise in the final step of its implementation. It is also this factor that throughout outperforms the univariate benchmark model by implying relative forecast errors  $< 1$ .

In Table 3, the first three columns with numerical entries show the test statistics of the pairwise test proposed by Diebold and Mariano (1995) to analyze whether empirical loss differences between competing models are statistically significant. It is generally seen to be the most influential and most widely used test in this context (see Carstensen et al., 2011, p. 87). The second triple-block of columns with numerical entries refers to the test statistics of the modified Diebold-Mariano (MDM) test suggested by Harvey et al. (1997) that corrects for small sample bias. It evaluates whether the average loss difference between a pair of models is significantly different from zero. Under the null the forecast accuracy of two models is identical; see e.g. Carstensen et al. (2011, p. 87) for detail. Note, each entry of Table 3 displays the respective test statistics for the forecasting performance of each model listed in the respective row vis-à-vis its respective competitor model in the header to each column. For instance, the combination of row “with EBA” and column “w/o, i.e. without, EBA” amounting to  $-1.94$  ( $-1.719$ ) for the (modified) Diebold-Mariano test corresponds to the respective test statistics of the MSE of model “with EBA” minus the corresponding MSE of model “w/o EBA.” A negative value implies that the model behind the subtractive MSE-part (here “w/o EBA”) performs worse in terms of forecasting accuracy. Corresponding p-values are given in parantheses. There are two differentials between models that are statistically significant for the standard Diebold-Mariano (SDM) test. Both nowcasts “with EBA” and the EC-one significantly outperform the model without DEBA pre-selection stage (i.e. the small-scale ADFM) at the ten percent and at the one percent level of significance, respectively. This holds also for the MDM test, where the first difference now turns out statistically significant even at a five percent level of significance. Additionally, the EC nowcast is statistically significantly better in terms of accuracy than the AR(2) benchmark, at least, at a ten percent level of significance, when correcting for small sample properties (MDM test). Moreover, neither for the SDM nor the MDM test, we

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<sup>7</sup>One reason for doing so is that one might be bothered that by merely dropping some indicators in the derivation of the common factor raises the weight of the stock exchange for real estate and finance index in the constitution of the factor and through this channel alone gives the (D)EBA based factor a comparative advantage. However, as can be seen from Table 2, the FE difference between the two factors remains if we leave out the recent crisis which is generally seen to be rooted in Spanish housing price dynamics. Interestingly, the FE for the model with the 14 core variables and the benchmark AR(2) process is almost the same excluding the crisis.

find a statistically significant difference in forecast accuracy between the EC nowcast and the micro-scale ADFM nowcast with DEBA pre-selection of indicators.

## 5 Conclusion

We examined the implementation stage of an ADFM to nowcast GDP growth using Spanish time series. To overcome a common deficiency of the literature, treating candidate indicators for the estimation of the common factor not symmetrically, we propose a two-tiered procedure. First, variables highly correlated with growth are identified as “core.” Secondly, they are checked for robustness according to a dynamic version of EBA standards to nowcast Spanish GDP growth. We find that our implementation clearly improves the accuracy of nowcasts compared to ADFM nowcasts implemented without the help of sensitivity analysis and to a benchmark AR(2) process. It leads to nowcasts almost measuring up to real-time forecasts by an institution with an interest in high-quality nowcasts (i.e. by the EC). Thus, we conclude by answering our central research questions raised in the introductory section: First, it can pay to consider a symmetric shrinkage device in the pre-selection process. Secondly, parsimony can improve nowcasts, even if we rely on a rather ad hoc pre-selection device and a shrinkage device with focus on selecting economically meaningful indicators rather than on pure forecast performance in going from small-scale to micro-scale ADFM.

It remains for future work to check our partially ad hoc framework for its robustness in selecting predictors with regard to the choice of parameters both at the pre-selection (in particular, concerning  $\rho$ ) and at the shrinkage stage (in particular, concerning  $k$ ). A further future task, that is beyond the scope of the present paper, is to scrutinize the trade-off between using all variables and adding more factors on the one hand (see, e.g. Giannone et al., 2008) and selecting variables a-priori and then using only a single factor on the other. Such a check might be based on the quasi maximum likelihood approach proposed in Doz et al. (2012).

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