Forecasting Industrial Production and the Early Detection of Turning Points[1]

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Abstract

In this paper we propose a simple model to forecast industrial production in Italy. We show that the forecasts produced using the model outperform some popular forecasts as well as those stemming from a trading days- and outlier-robust ARIMA model used as a benchmark. We show that the use of appropriately selected leading variables allows to produce up to twelve-step ahead reliable forecasts. We show how and why the use of these forecasts can improve the estimate of a cyclical indicator and the early detection of turning points for the manufacturing sector. This is of paramount importance for short-term economic analysis.

Keywords: Forecasting, VAR Models, Industrial production, Cyclical indicators.

JEL classification: C53, C32, E32.

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1 Introduction

Forecasting the industrial production index is an important issue in short-term economic analysis. This is still true in the nowadays economies where services are undertaking an increasing weight. In fact, the industrial sector is still important in explaining aggregate fluctuations, also because some of the services activities (business services) are closely linked to the industrial ones. The interest in the topic is witnessed by the continuous effort devoted to investigating it (see, *e.g.*, Bodo *et al.*, 2000; Huh, 1998; Marchetti and Parigi, 2000; Osborn *et al.*, 1999). In addition, forecasts of industrial production can be useful in more general forecasting models. Furthermore, cyclical indicators of the manufacturing sector may be derived from the industrial production index series: this is commonly done by applying signal extraction techniques, and accurate forecasts of the series to be filtered are essential in order to obtain reliable estimates at the end of the series.

From a general standpoint it should be said that most of the existing models in Italy offer an early estimation of the industrial production, rather than a pure forecast. Indeed, the official indicator is released by the National Statistical Institute (ISTAT) 45 days after the end of the reference month, so that a two-step ahead prediction is necessary to achieve a *nowcast* of the indicator itself. This is the case, in particular, for two popular predictions released monthly by CSC and IRS,[1] respectively. In the second half of month t (when the official indicator is available up to month $t-2$), CSC releases a preliminary survey-based estimate of month t and a revised estimate for month $t-1$. A similar dissemination scheme is followed also by IRS, which however uses a model based on electricity consumption to produce its projections; at the half of month t a preliminary estimate of the same month is released, and a final one is published at the beginning of month $t+1$.

In this paper, we propose a simple model able to produce satisfactory forecasts of the industrial production index well beyond the two-step ahead *nowcasts.* We show that the projections deriving from a simple VAR using appropriately selected leading variables can well compete with the two aforementioned accredited forecasts in terms of predictive ability. We also show how our projections can be used successfully to reduce uncertainty in the estimation of a cyclical indicator. Finally, we use these predictions also to

[1] CSC (Centro Studi Confindustria) is the research department of Confindustria, the Confederation of Italian Industry. IRS (Istituto per la Ricerca Sociale) is an independent no-profit social research centre. Speaking of "models" when referring to CSC projections is, strictly speaking, inappropriate, given that they are derived from a survey. However, for brevity we will refer to the different forecasting devices as "models". This should cause no confusion.

improve substantially on the timely detection of turning points in the manufacturing sector. We actually think that this is a major result of this paper. Though the empirical analysis is carried out on Italian data, we feel that the implications are far reaching and the arguments developed in the paper are potentially of interest to an international audience.

The paper is organized as follows: in the next section we illustrate some preliminary analyses carried out on the time series used in the forecasting exercise; Section 3 describes the forecasting model and Section 4 is devoted to the evaluation of its predictive ability, also in comparison with the competing predictions released by CSC and IRS. Section 5 discusses the improvements deriving from using our forecasts in estimating a cyclical indicator for the manufacturing sector, and in turning point detection. The final Section concludes.

2 Preliminary analysis

One of the first logical steps in a modelling strategy is the review of the available information. Econometric models already available in Italy to forecast industrial production mainly use coincident indicators of industrial activity, such as electricity consumption (see *e.g.* Marchetti and Parigi, 2000), which have the advantage of an earlier release with respect to the industrial production index, making it possible to formulate up to two-period ahead predictions (nowcasts).

The goal of obtaining "true" forecasts of the industrial production index (*IPI*), makes it necessary to forecast the official figures at least three months ahead. This is why it might be sensible to give priority, in the search of variables which will be used in the forecasting model, to those characterized by a leading pattern. A comprehensive analysis of the properties of many Italian economic time series has been carried out by Altissimo *et al.* (1999). In part using the results contained in that paper, and after restriction of a considerably higher number of candidates, we find that two variables seem particularly interesting as potential predictors of the industrial production in Italy: the ISAE business surveys series[2] of future production prospects (*PP*) and the quantity of goods transported by railways (*TON*).[3] The first variable

[2] See Pappalardo (1998) and the references therein for a description of the uses of ISAE business surveys in forecasting models of Italian industrial production.

[3] The industrial production index (*IPI*) is released monthly by ISTAT, the Italian National Statistical Institute. Press releases and recent data can be found at `http://www.istat.it/Anews/proind.html`. Future production prospects (forecasts, *PP*) are released monthly by ISAE, the Institute for Studies and Economic Analyses.

represents industrial entrepreneurs' opinions about future production. More precisely, the entrepreneurs are asked if the production in the following three-four months is expected to go "up", to be "stable", or to go "down". The answers are then synthesized with a *balance*, *i.e.* the share of "up" less the share of "down" answers. The variable obtained (PP) is therefore bounded in the interval [-100, 100], and it is a natural candidate in a forecasting model given its timely availability,[4] its explicit link with the variable to be forecast, and its long lead over the industrial production series. The usefulness of the second variable in a forecasting model is due to the fact that the merchandises transported by rail are mainly intermediate goods and raw materials used as inputs by manufacturing industries. Indeed, this variable is characterized by a fairly stable lead over the industrial production index, as well as by a short delay in its availability.

In the next sub-section a preliminary analysis of the univariate characteristics of the series of interest is presented. A log transformation is used for the series IPI and TON while PP is rendered unbounded using the transform[5]

$$-\log\left(\frac{200}{PP+100}-1\right). \tag{1}$$

2.1 Description of the series

Figure 1 plots the series used in this paper. They are characterized by rather heterogeneous patterns. The industrial production index shows strong seasonality, with some cyclical fluctuations around an upward trend. About the same can be said of the railways transport of goods, which, however, seems to feature more pronounced cyclical movements, together with some potential outliers. A rather different pattern characterizes the production expectations of industrial firms: in this case the cyclical movements are clearly predominant with respect to the other components, even though some seasonality seems to be present.

The relative importance of trend, seasonality and cyclical movements can be better appreciated by resorting to the Fourier representation of time series. An estimate of the spectrum[6] of the series is showed in Figure 2. This

Recent data and updates can be found at `http://www.isae.it/english.html.` The time series for tons of goods transported by railways (TON) and its updates are kindly provided by Ferrovie dello Stato, the Italian State railways company.

[4] At the beginning of month t, the results for $t-2$ are released.

[5] For ease of exposition we avoid creating further acronyms and, from now on, we maintain the same names for the transformed variables. This should create no confusion.

[6] Actually, the spectral density is correctly defined only for stationary time series; in particular, when some unit roots are present in correspondence of certain frequencies, the

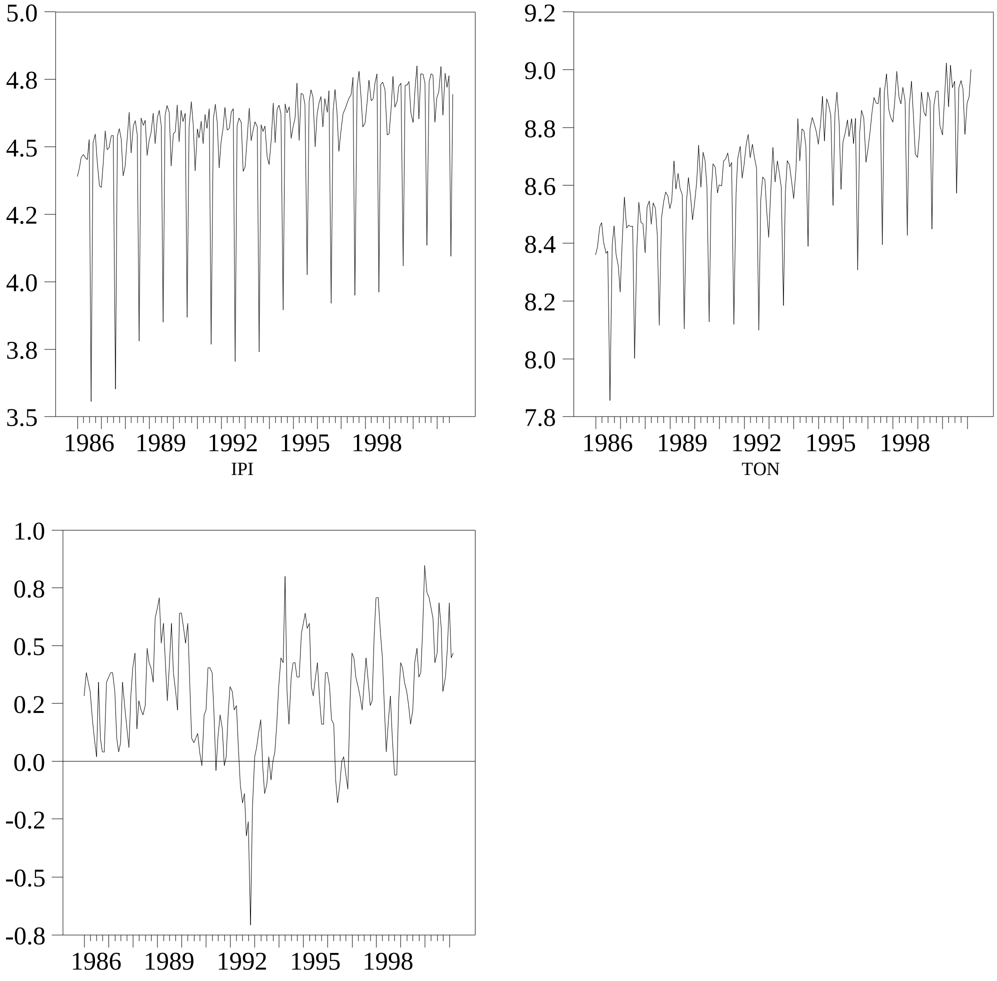

Figure 1: The time series.

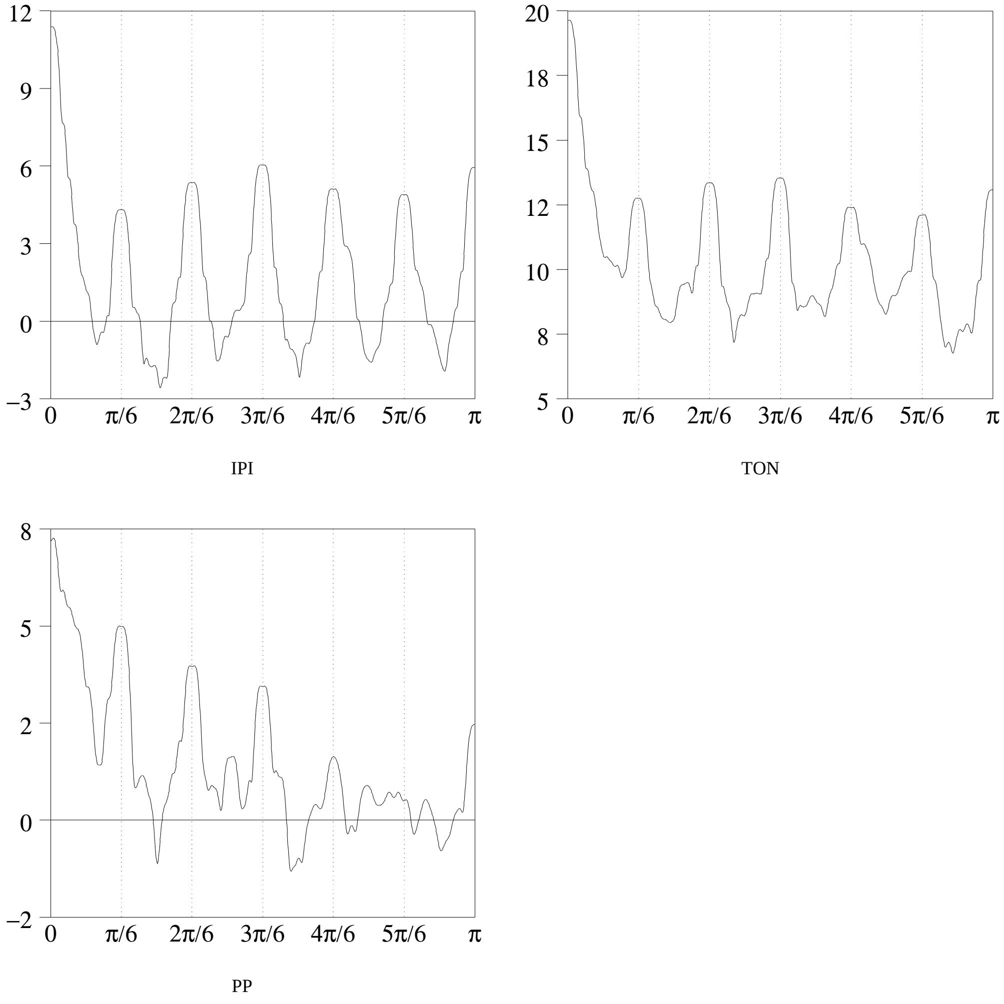

Figure 2: The spectra.

Table 1: Tests for unit roots

frequency	*IPI* (3 lags)		*TON* (3 lags)		*PP* (no lags)	
0	-2.64		-3.51	*	-2.68	
$\pi/6$	5.71		5.49		15.75	**
$\pi/3$	2.59		6.48		13.74	**
$\pi/2$	6.90	*	11.69	**	14.16	**
$2\pi/3$	12.00	**	8.74	*	14.16	**
$5\pi/6$	6.28		7.55	*	16.48	**
π	-1.90		-1.76		-2.45	

t-tests for the 0 and π frequencies, F-tests for the others. Values significant at 5% and 1% are indicated by '*' and '**', respectively.

confirms the previous observations, showing a similarity between the spectra of *IPI* and *TON*, with a concentration of power at low and seasonal frequencies. A slightly different pattern emerges when *PP* is considered: its long-run component has a smaller peak in the spectrum, while considerable power seems to be present at the cyclical frequencies. Important peaks are present also at the fundamental seasonal frequency and, though less important, at frequencies associated with periodicity of 4 and 6 months.

2.2 Stochastic properties

The stochastic properties of the three series can be examined in a more formal manner, testing for the presence of unit roots. Given the nature of the data we are dealing with, it is natural to test for unit roots at the zero and at the seasonal frequencies, and we do so using the approach detailed in Beaulieu and Miron (1993). The test regression includes a constant, a trend, eleven seasonal dummies, and the number of lags of the dependent variable sufficient to whiten the residuals. The results reported in Table 1 reject the presence of a unit root at frequency zero for *TON*, while it is not rejected for the others. In all cases the presence of all the twelve unit roots is strongly rejected.

The conclusion we draw from the previous evidence is that the presence of unit roots at some seasonal frequencies cannot be overall excluded. The

usual expression for the spectral density would take the value $+\infty$ at those frequencies; nevertheless, discarding them the so called *pseudo spectrum* (Bell, 1984) can be considered. In our case the spectrum has been estimated by smoothing the periodogram using a rectangular spectral window with spectral bandwidth equal to 0.048π. A cosine taper has been applied to the data. y-axis values are expressed in logs.

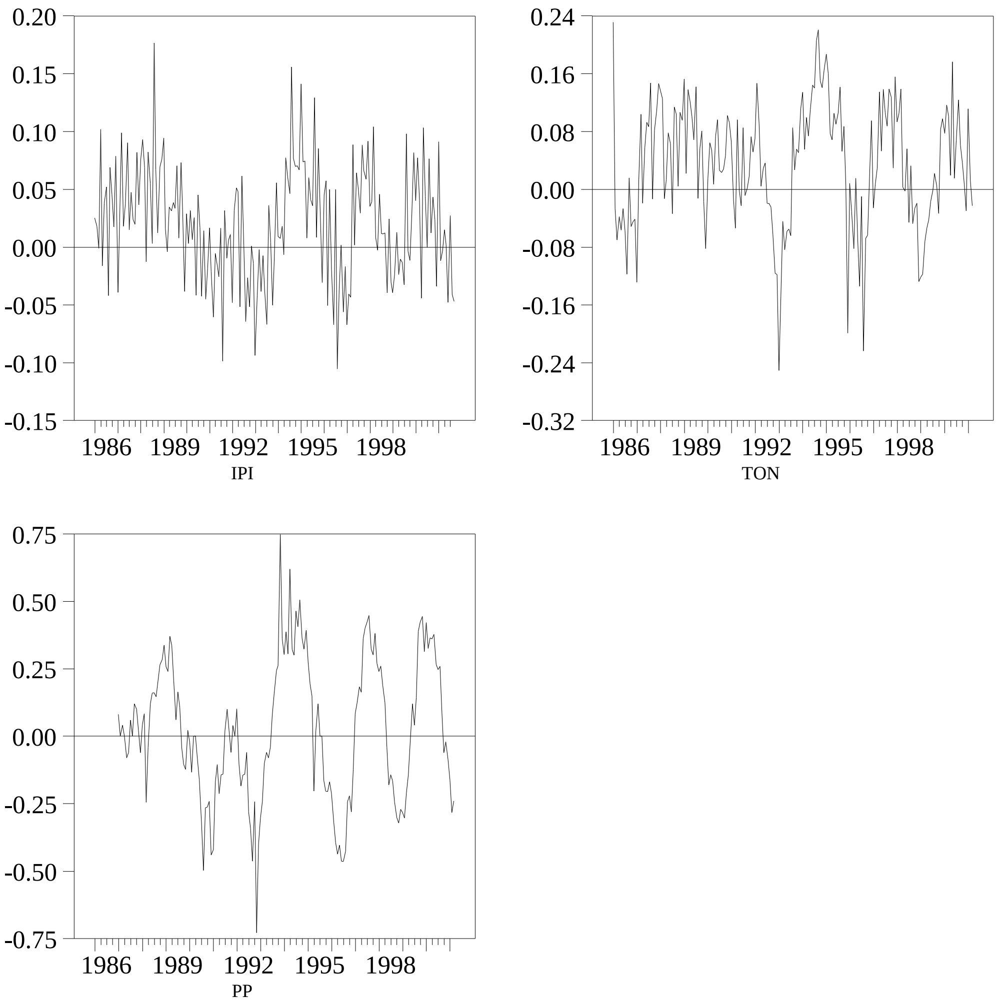

Figure 3: Differenced series.

Table 2: Tests for unit roots in seasonally differenced series

	ADF_c		$ADF_{c,t}$		BG2	BG3	BG4	BG5	BG6
$\Delta_{12}IPI$	-3.589	**	-3.566	*	0.061	0.061	0.061	0.860	0.713
$\Delta_{12}TON$	-4.105	**	-4.332	**	0.318	0.315	0.316	0.791	0.114
$\Delta_{12}PP$	-3.684	**	-3.666	*	0.632	0.586	0.588	0.328	0.219

Δ_{12} denotes the seasonal difference. The table reports the results from the Augmented Dickey-Fuller (with constant, ADF_c, and with constant and trend, $ADF_{c,t}$) tests for zero-frequency unit roots and from the Bierens-Guo (BG2-BG6) tests for stationarity (Bierens and Guo, 1993) applied to the seasonally differenced series. For the ADF tests, "*" and "**" denote values that are significant at 5% and 1%, respectively. For the Bierens-Guo tests, p-values are reported.

application of the seasonal difference operator produces the series plotted in Figure 3, where the cyclical pattern emerges more clearly, especially for *PP*, while *IPI* and *TON* are characterized also by strong irregular movements, some of which can be reasonably attributed to trading days effects. This feature is even clearer if we consider the spectral representation depicted in Figure 4, where seasonal peaks have disappeared and the long term component at the zero frequency is less pronounced. This makes much more evident the presence of a trading days pattern for the series *IPI* and *TON*, represented by peaks in the spectrum at the frequencies highlighted by vertical lines in the figure (on this aspect see Cleveland and Devlin, 1982). In order to investigate the presence of multiple unit roots at the zero frequency, we carry out formal tests on the annual differences of the series. The results of these tests are reported in Table 2. Both the unit roots and the stationarity tests seem to indicate that the annual differences of our series do not contain unit roots at frequency zero.

2.3 Cyclical characteristics

One important characteristic of the Italian industrial production index is represented by its strongly cyclical behavior. Indeed, this is perhaps the most important feature one is normally interested in when formulating the forecasts. This is the reason why series used to help forecasting industrial production index should be characterized by a regular lead on the latter at cyclical frequencies. We show that this is the case for the series considered in this paper.

We extract the cyclical component of each series by means of the *band-*

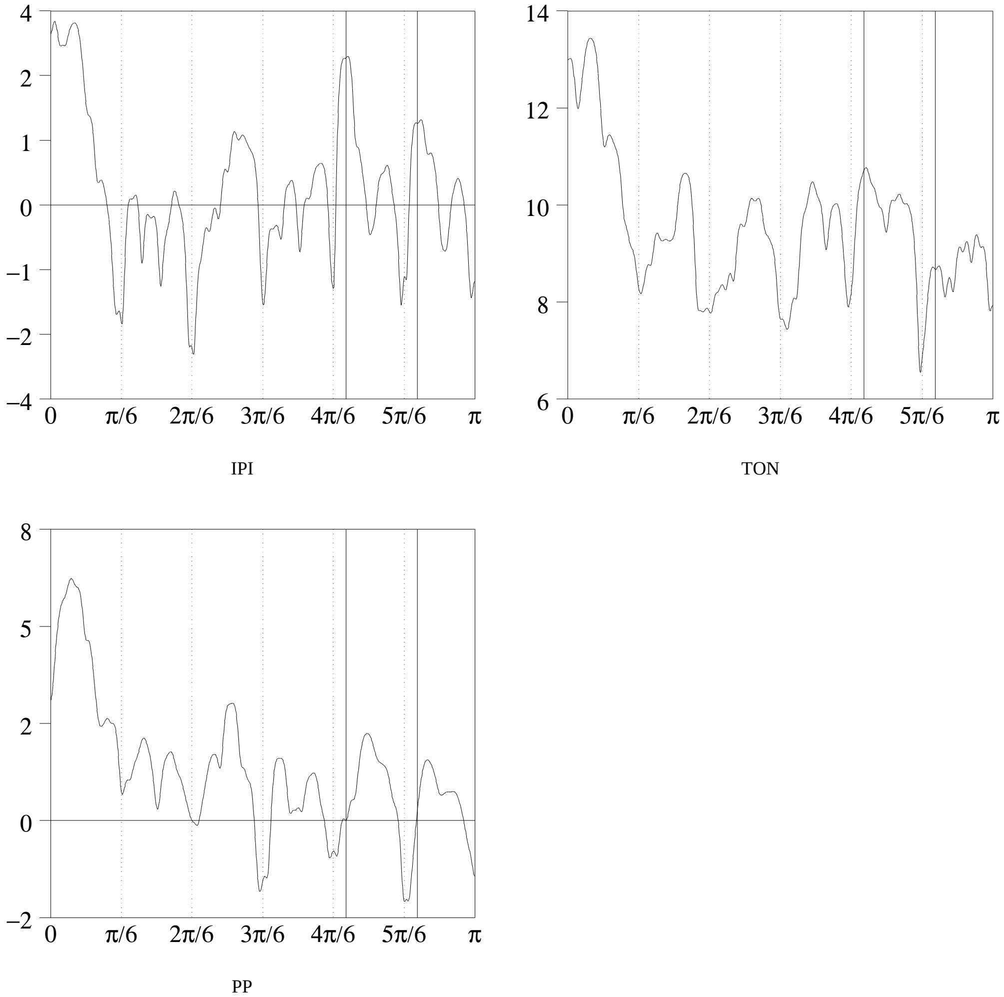

Figure 4: Spectra of the differenced series.

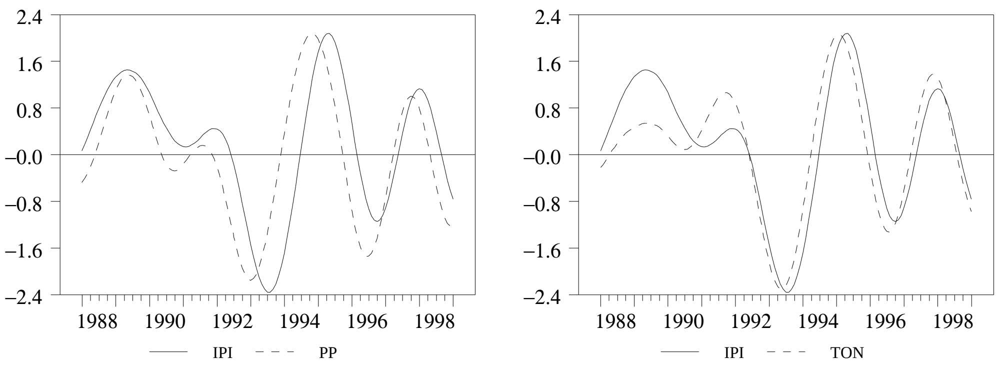

Figure 5: Cyclical components (2-8 year period component, standarized).

Table 3: Correlation of the cyclical components of TON and PP with that of IPI

	$\rho(0)$	$\rho_{\max}$	lead (+) or lag (-)
TON	0.87	0.92	+2
PP	0.68	0.90	+5

"$\rho(0)$" is the correlation between the series and IPI; "$\rho_{\max}$" is the maximum cross-correlation; the last column reports the time interval (in months) at which the maximum cross-correlation is observed.

pass filter developed by Baxter and King (1999). The estimated components are plotted in Figure 5 over the period 1988:1 -1999:1, leaving out two years of observations at the beginning and at the end of the available sample, since the band-pass filter is a symmetric one and the filtered observations at the extremes cannot be estimated. The figure shows a clear and regular lead of the production prospects over the industrial production, which is consistent with the nature of the series. The leading nature of the *TON* series is less evident but nevertheless holds on average, as highlighted in Table 3. The cyclical component of *PP* is confirmed to lead consistently that of *IPI*, on average by 5 months, with a high correlation. The correlation is even more pronounced if *TON* is considered, even though with a shorter lead (2 months on average).[7]

3 The forecasting model

An explicit goal of this study is that of finding a simple but reliable model to forecast the monthly Italian industrial production index. On the one hand, empirical evidence on the forecasting performance of nonlinear models is mixed (see *e.g.* Clements and Krolzig, 1998; Huh, 1998; Marchetti and Parigi, 2000; Simpson *et al.*, 2000). Franses and van Dijk (2001) suggest that linear models with simple seasonal components offer advantages over more complicated ones in terms of their short-term forecasting accuracy. On the other hand, we feel that the single-equation framework often used to forecast the industrial production index (see *e.g.* Marchetti and Parigi, 2000; Simpson *et al.*, 2000) offers an oversimplified option and does not allow for multi-step dynamic forecasts. For all these reasons our investigation rests on the well established VAR framework.

Given that we use seasonal time series, an aspect that deserves special attention is the parameterization of the VAR. According to the results listed in Section 2, the three time series that we consider have different seasonal properties: all display the presence of at least one unit root, but none of them seems to possess all the seasonal roots equal to unity. This implies that if we parameterize the VAR in seasonal differences, we are likely to over-difference the series at some frequencies. However, we believe that unit-roots pretesting is useful for forecasting, despite the potentially low power of the tests (see Diebold and Kilian, 2000, for a discussion related to unit roots at the zero-frequency). Indeed, not much is known about the effects on forecasting performance deriving from imposing all the seasonal roots at unity when this

[7]The use of cross-correlations estimated among filtered series might be questioned (see, among others, Canova, 1998). However, we utilize them only as descriptive devices.

Table 4: Main VAR diagnostics: estimation period 1988.3-1997:12

	σ	Corr(Act., Fit.)	AR 1-12	Norm.
$\Delta\Delta_{12}IPI$	0.020	0.962	0.355	0.310
$\Delta\Delta_{12}TON$	0.045	0.868	0.283	0.620
$\Delta\Delta_{12}PP$	0.110	0.761	0.217	0.418
VAR			0.221	0.601
Parameter constancy forecast tests (1998:1-2001:2)				
$\mathsf{F}_{\mathbf{\Omega}}$	0.524			
$\mathsf{F}_{\mathsf{V}(\mathbf{e})}$	0.944			
$\mathsf{F}_{\mathsf{V}(\mathbf{E})}$	0.959			

The Table reports the standard error of each equation in the VAR (σ), the correlation of actual and fitted values (Corr(Act., Fit.)), the p-value of the LM test for residuals autocorrelation up to the twelfth order (AR 1-12), and the p-value of the test for residuals normality (Norm.). The p-values of the tests on the residuals of the VAR as a whole are also reported in the row labelled "VAR". The values reported for the parameter constancy forecast tests are p-values of the tests in their F-form. The first one ($\mathsf{F}_{\mathbf{\Omega}}$) does not consider parameter uncertainty.

is not the case in reality. To the best of our knowledge, the empirical evidence does not offer a definitive answer, though there are indications that filtering out only the correct unit roots in general does not produce superior forecasts (see *e.g.* Clements and Hendry, 1997; Gustavsson and Nordström, 1999; Lyhagen and Löf, 2001; Osborn *et al.*, 1999; Paap *et al.*, 1997). In particular, Lyhagen and Löf (2001) suggest that when the model is not known and the aim of the modeling exercise is forecasting, a VAR in annual differences may be a better choice than a seasonal error correction model based on seasonal unit roots pre-testing. Given that the series considered in this paper appear to be at most $\mathsf{I}(1)$, and considering also that differencing partially protects forecasting from structural breaks (Clements and Hendry, 1999), we parameterize our VAR in seasonal differences. Also note that, following the standard short-term economic analysis practice, we are mostly interested in forecasting the annual growth rates, not the levels of the series.

As far as model selection is concerned, we rely on the general-to-specific approach and we start from a fairly heavily parameterized VAR with 14 lags and some deterministic components. In particular, we want to simplify the model sequentially by excluding non significant lags, starting from the least significant, while checking at each simplification step the statistical properties of the residuals. Though seasonal differences do effectively filter out most of the seasonal component of the series, nevertheless they still show high and slowly decreasing autocorrelations which make it difficult to find a valid (subset) reduction of the starting model (see also Krolzig, 2001). In order to obtain quasi-orthogonal regressors so to ease the reduction process, we reparameterize our stationary VAR into the isomorphic form

$$\Delta\Delta_{12}\mathbf{y}_t = \boldsymbol{\beta}'\Delta_{12}\mathbf{y}_{t-1} + \sum_{j=1}^{13}\boldsymbol{\gamma}_j'\Delta\Delta_{12}\mathbf{y}_{t-j} + \boldsymbol{\phi}'\mathbf{d}_t + \boldsymbol{\varepsilon}_t \tag{2}$$

where $\Delta = (1 - L)$, $\Delta_{12} = (1 - L^{12})$, L is the usual lag operator such that $L^p z_t = z_{t-p}$, $\mathbf{y}_t = (IPI_t,\ TON_t,\ PP_t)'$, and $\mathbf{d}_t$ are the deterministic components. We find successful not to include all the seasonal dummies in the model (the p-value of the test for the exclusion of all the seasonal dummies, except that for August, is 0.991). Rather, $\mathbf{d}_t$ includes, besides the constant and the dummy for August, three specific impulse dummies, and two special dummies for August and December that take the value 1 when production prospects (PP) are positive and -1 when they are negative.[8] This approach represents an attempt to take into account possible interactions between seasonal variations and business cycle in industrial production. These interactions can be theoretically justified by economic theory (see *e.g.* Cecchetti *et al.*, 1997), and can produce observable implications. In fact, it is common practice for firms in Italy to adjust production to demand by prolonging (shortening) summer and Christmas holidays when demand is low (high). Furthermore, $\mathbf{d}_t$ includes also $\Delta_{12}\log(TD_t)$ and $\Delta_{12}\log(TD_{t-1})$, with TD_t the number of trading days in month t. As is well known, the number of trading days significantly influences manufacturing activity. While the use of $\Delta_{12}\log(TD_t)$ is common in models for industrial production, the insertion of $\Delta_{12}\log(TD_{t-1})$ is fairly non-standard. However, in the presence of particularly unfavorable (favorable) trading days configurations, it is legitimate to expect that firms tend to compensate lower (higher) realized production in the following month. Indeed, the estimated coefficients of $\Delta_{12}\log(TD_t)$ and $\Delta_{12}\log(TD_{t-1})$ in our VAR are both highly significant and seem to confirm this view.

The VAR is sequentially simplified to obtain a more parsimonious parameterization. Even if the (subset) restricted VAR is more parsimonious than

[8] Strictly speaking, the use of these dummies is such that the model is no longer a VAR. However, given the rather special role of these variables, for brevity we prefer to continue denoting our model as "VAR". We used also a parameterization in which positive and negative dummies were separated, but the attached coefficients resulted not significantly different in absolute value. For this reason we prefer the more compact form described in the text.

Table 5: Main VAR diagnostics: estimation period 1988.3-2001:2

	σ	Corr(Act., Fit.)	AR 1-12	Norm.
$\Delta\Delta_{12}IPI$	0.019	0.962	0.103	0.097
$\Delta\Delta_{12}TON$	0.043	0.865	0.056	0.774
$\Delta\Delta_{12}PP$	0.106	0.706	0.254	0.258
VAR			0.286	0.268

The Table reports the standard error of each equation in the VAR (σ), the correlation of actual and fitted values (Corr(Act., Fit.)), the p-value of the LM test for residuals autocorrelation up to the twelfth order (AR 1-12), and the p-value of the test for residuals normality (Norm.). The p-values of the tests on the residuals of the VAR as a whole are also reported in the row labelled "VAR".

the starting one, nevertheless it is still rather highly parameterized including lags from 1 to 5, lag 9, and lags from 12 to 13. The p-value of the reduction is 0.8026, which indicates that no significant information is lost in the sequential simplification process. The main statistics and diagnostics of the VAR estimated over the period 1988:1-1997:12 are reported in Table 4.[9] The tests for parameter constancy, calculated over the forecast evaluation sample (see next section), do not reject structural stability.

For completeness, in Table 5 we report the main statistics and diagnostics of the VAR estimated over the full sample (1988:3-2001:2).

The final model we use to actually produce forecasts is further simplified by eliminating non significant deterministic elements from individual equations.[10]

4 Forecast evaluation

In this section we evaluate the forecasting ability of our VAR as opposed to an ARIMA model and to the forecasts released by CSC and IRS over a fairly long period (1998:1-2001:2).[11] Given that we are especially interested in forecasting industrial production annual growth rates, all forecasts

[9] The results have been obtained using Pc-Fiml 9.30 (see Doornik and Hendry, 2000).

[10] This further simplification and the procedure to routinely produce the forecasts are implemented in WinRATS 5.00 (see Doan, 2000).

[11] Time series of past CSC forecasts have been obtained from the Confindustria Website (`http://www.confindustria.it/DBImg.nsf/HTMLPages/CSCCongSIHome`) and start from 1998:3. IRS forecasts have been retrieved from the articles published in the financial newspaper *Il Sole 24 Ore*.

comparisons refer to this variable. To make the evaluation more interesting, the ARIMA model is rendered robust to potential outliers and is enriched with a deterministic part that includes trading days and Easter effects. This model is estimated recursively by maximum likelihood and the forecasts are produced using TRAMO (see Gómez and Maravall, 1998). Such an ARIMA constitutes a very robust benchmark to beat.

Perfectly fair forecasts comparisons would require the use of homogeneous forecasting criteria among the competing models (see *e.g.* Tashman, 2000). However, we want to compare our forecasts with those from a model of which we do not know many details, and even with those derived from a survey. For this reason we believe that, while perfectly homogeneous conditions are essential when comparing the forecasting performance of alternative *methods*, they cannot be imposed when comparing *real world forecasts*. However, to increase comparability both the ARIMA and the VAR forecasts are based on a recursive scheme. Parameters are estimated with data ranging from 1 to t_0, and forecasts are produced for $t_0+1, \ldots, t_0+n$ $(n \geq 1)$; then parameters are estimated on the sample ranging from 1 to t_0+1, and forecasts are produced for $t_0+2, \ldots, t_0+n+1$, and so forth. In our application the forecast evaluation sample runs from January 1998 to February 2001: the estimation sample is adjusted in such a way that for each forecasting horizon we have 38 out-of-sample observations.

Comparisons are somewhat complicated by some peculiarities in the CSC and IRS forecast samples. In fact, CSC does not produce one-step ahead estimates for the month of July and two-step ahead projections for the month of August of each year. IRS does not release two-step ahead forecasts for the month of August of each year; additionally, we could not retrieve IRS forecasts for a couple of dates.[12]

Macroeconomic analysts might be interested, more than on the numerical indications arising from the forecasts, on their signs, since these can be perceived as warnings of expansions or contractions. For this reason we think that it is useful to start the investigation of the forecasting performance of our VAR model from an analysis of the directional forecasts. In our forecast sample there are 23 observations for which the industrial production annual growth rates are positive, 14 for which they are negative and one (July 2000) in which the annual growth rate is zero. We assume that if a prediction has wrong sign, but the difference with the actual growth rate is less than one percentage point, the sign of the forecast is correct. This avoids considering

[12] This happened in corrispondence of dates for which IRS released only the seasonally adjusted figures.

Table 6: Directional forecast errors: wrong predictions as percentage of valid observations.

steps ahead	1	2	3	6	12
ARIMA	13.16	13.16	7.89	31.58	13.16
VAR	13.16	13.16	13.16	10.53	10.53
NAIVE	45.95	41.67			
CSC	15.15	12.12			
$\text{ARIMA}_{\text{CSC}}$	12.12	15.15			
VAR_{CSC}	12.12	12.12			
IRS	21.62	23.53			
$\text{ARIMA}_{\text{IRS}}$	13.51	14.71			
VAR_{IRS}	13.51	11.76			

Wrong directional forecasts as percentage of valid observations. The one-step and two-step ahead "NAIVE" directional forecasts are given by $\text{sign}(\Delta_{12}IPI_{t-1})$ and $\text{sign}(\Delta_{12}IPI_{t-2})$, respectively. "ARIMA" and "VAR" with the subsctript "CSC" and "IRS" denote the statistics calculated on the forecasts from the ARIMA benchmark and the VAR model over the same sample used for the CSC and IRS forecasts, respectively.

as wrong outcomes close to zero.[13] The results from this comparison are reported in Table 6 and indicate that the gain with respect to a naive forecasts defined as $\text{sign}(\Delta_{12}\widehat{IPI_t}) = \text{sign}(\Delta_{12}IPI_{t-i})$ with $i = (1, 2)$ is substantial. Furthermore, the ARIMA, CSC, and VAR projections show similar directional errors, while IRS prediction errors nearly double the others. Finally, the ARIMA benchmark shows a large fraction of errors corresponding to the six-step ahead forecasts. It should be noted, however, that contingency tables-based tests on the directional forecasts[14] are always very significant, indicating that direction-of-change forecasts are informative for all the predictions considered in this paper. This remains true even for the forecasts of $\Delta\Delta_{12}IPI_t$.

Figure 6 plots the estimated densities of the forecasting errors for the different models, computed on the largest common sample. Note that IRS

[13] However, results do not change qualitatively if we use a strict criterion.

[14] The detailed results are not reported for brevity. On the characteristics of the tests see *e.g.* Diebold and Lopez (1996).

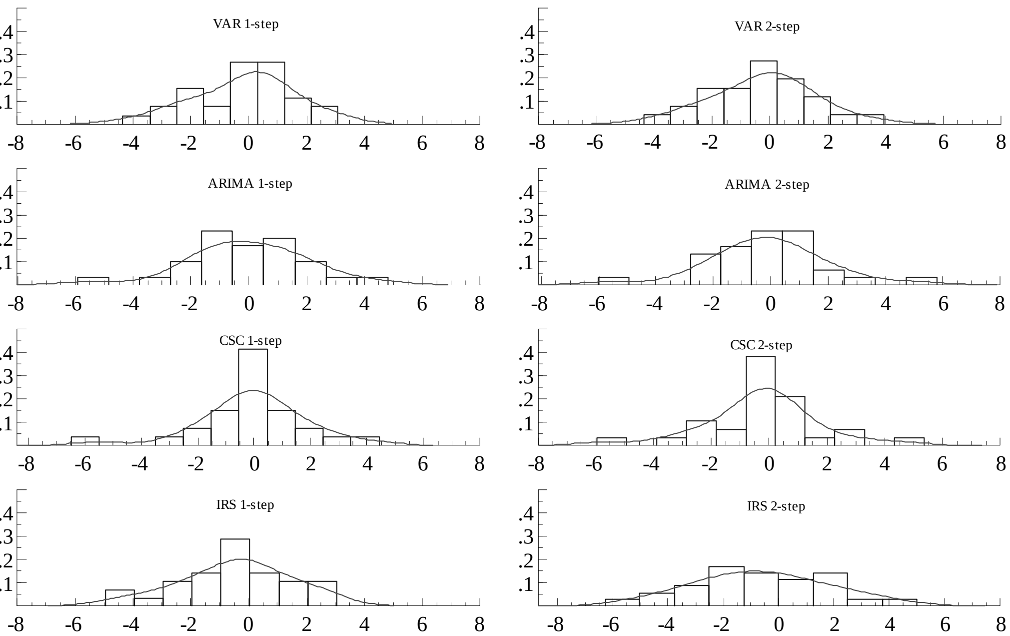

Figure 6: Estimated densities of forecasting errors.

Table 7: Mean absolute errors and mean errors of yearly growth rates (percent) forecasts of industrial production

steps ahead	1	2	3	6	12
ARIMA	1.64	1.62	1.65	2.16	3.31
	0.16	0.17	0.16	0.23	-0.26
VAR	1.25	1.27	1.31	1.54	1.49
	0.11	0.17	0.22	0.18	0.18
CSC	1.33	1.26			
	-0.12	0.32			
$\text{ARIMA}_{\text{CSC}}$	1.59	1.31			
	0.05	0.01			
VAR_{CSC}	1.23	1.21			
	0.03	0.13			
IRS	1.46	1.94			
	0.32	0.50			
$\text{ARIMA}_{\text{IRS}}$	1.56	1.32			
	0.12	0.07			
VAR_{IRS}	1.29	1.24			
	0.11	0.20			

For each number of steps ahead, the first row reports the mean absolute error (MAE), while the second one shows the mean error (ME). "ARIMA" and "VAR" with the subsctript "CSC" and "IRS" denote the statistics calculated on the forecasts from the ARIMA benchmark and the VAR model over same sample used for the CSC and IRS forecasts, respectively.

forecasts seem to be the most uncertain, while CSC projections are the most concentrated around zero, even if they have fatter tails than the VAR predictions. Though informative, Figure 7 could be not very significant on statistical grounds, given that the sample is rather short. The features of the forecasts in terms of their mean absolute error (MAE) and mean error (ME) are reported in Table 7. Our VAR model's forecasts uniformly outperform the others in terms of MAE. Furthermore, note that the ratio of the twelve-step ahead to the one-step ahead MAE is 2.02 for the ARIMA model and only 1.19 for the VAR. This shows how important is the cyclical information embodied in *PP* and *TON*.

The ability of traditional forecasting evaluation measures over long fore-

Table 8: Performance in forecasting annual percentage growth rates.

	1998	1999	2000
Actual	+1.9	+0.1	+3.2
VAR	+2.8	-0.4	+3.2
Benchmark	+4.7	-0.8	+2.2

casting horizons has been questioned by García-Ferrer and Queralt (1997). As an alternative, they suggest comparing the forecasts of future annual growth rates. In the case considered here, this comparison is not feasible between VAR model forecasts and the short term forecasts issued by CSC and IRS, because the latter are produced just for the two subsequent months; nevertheless, the performance of the VAR model can be compared with the ARIMA forecasts and the actual values. We made this evaluation estimating both the VAR and the ARIMA until December of year t, generating then 12 forecasts, and calculating the forecasted growth rate of year $t+1$ with respect to year t. This figure is then compared to the actual growth rate. The results showed in Table 8 confirm the ability of the VAR to track fairly accurately the actual annual rate of change of the industrial production index; also, for the three years considered in the table, the VAR consistently outperforms the benchmark, sometimes remarkably. For the year 2000 the VAR forecast is equal to the realized value.

In order to get a better assessment of the relative forecasting ability of our VAR as opposed to the other forecasts, we perform formal tests of (pair-wise) equal forecasting performance and forecast encompassing. Both classes of tests are variants of the test for predictive accuracy proposed by Diebold and Mariano (1995).[15] Suppose one has two series of n forecasts each to be compared. Let $\{e_{it}\}_{t=1}^{n}$ be h-step ahead forecast error deriving from model i. Denote by $d_t = g(e_{it}) - g(e_{jt})$ with $g(\cdot)$ some arbitrary (non necessarily symmetrical) pre-specified function. The null hypothesis of equality of expected forecast performance is $\mathsf{E}(d_t) = 0$. It is natural to consider $\bar{d} = n^{-1}\sum_{t=1}^{n} d_t$, so that $\sqrt{n}(\bar{d} - \mu_d) \xrightarrow{\text{d}} \mathsf{N}\left(0, 2\pi f_d(0)\right)$, where μ_d is the population mean of d_t and $f_d(0)$ is the spectral density of d_t at frequency zero, which in turn is $f_d(0) = (2\pi)^{-1}\sum_{\tau=-\infty}^{\infty}\gamma_d(\tau)$ with $\gamma_d(\tau)$ the lag-τ autocovariances. Diebold and Mariano (1995) propose basing the test of equal forecasting accuracy on

$$DM = \frac{\bar{d}}{\sqrt{n^{-1}2\pi\widehat{f_d(0)}}} \tag{3}$$

[15] The variants are those introduced by Harvey *et al.* (1997, 1998).

which, under the null, tends to $\mathsf{N}(0,1)$ when $\widehat{f_d(0)}$ is a consistent estimate of $f_d(0)$. In order to correct for the size distortions noticed in the test based on DM, Harvey *et al.* (1997, 1998) propose modifying the test as

$$DM^* = \left(\frac{n+1-2h+n^{-1}h(h-1)}{n}\right)^{1/2} DM \tag{4}$$

and comparing the results with the critical values from the Student's t distribution with $(n-1)$ degrees of freedom.

When comparing forecasting accuracy, in this paper we use $d_t = |e_{it}| - |e_{jt}|$: when performing tests of forecast encompassing, d_t becomes $d_t = e_{it}(e_{it} - e_{jt})$ (see Harvey *et al.*, 1998). Under the null, forecast i encompasses forecast j and $\mathsf{E}(d_t) = 0$: under the alternative, forecast i could be improved by incorporating some of the features present in forecast j.

In this paper we use the DM^* version of the tests. In order to obtain a consistent estimate of $f_d(0)$, we follow the recommendations contained in Diebold and Mariano (1995) and Harvey *et al.* (1997) and use an unweighted sum of the sample autocovariances up to $h-1$, that is $2\pi\widehat{f_d(0)} = \widehat{\gamma_0} + 2\sum_{\tau=1}^{h-1}\widehat{\gamma_\tau}$, with $\widehat{\gamma_k}$ the lag-k sample autocovariance.

Two remarks are important at this stage. First, given that CSC and IRS forecasts present some missing values, in the computation of $2\pi\widehat{f_d(0)}$ we use (see Harvey, 1989, p.329; Robinson, 1985)

$$\widehat{\gamma_k} = \frac{\sum_{t=1}^{n-k}(d_t^\dagger - \bar{d})(d_{t+k}^\dagger - \bar{d})}{\sum_{t=1}^{n-k} a_t a_{t+k}} \tag{5}$$

where $d_t^\dagger$ is d_t with zeros replacing the missing values, and $a_t = 1$ when d_t is observed and $a_t = 0$ otherwise. Second, West (2001) demonstrates that when forecasts are based on estimated models and parameters estimation uncertainty is neglected, the forecast encompassing test tends to reject too often. This size distortion depends, among other things, on the number of out-of-sample forecasts used to compute the test. When the fraction n/t_0 is small, the distortion is likely to be small. In our case, $n/t_0 \approx 0.25$: this implies that a nominal 5% t test should slightly over-reject, but the actual size should not exceed 8%.[16] Given that correction of DM^* to take into account parameters uncertainty entails knowledge of both the models to be compared, we cannot in practice use the modifications suggested by West (2001).

[16] A nominal 3% should not exceed actual 5%. These computations follow West (2001, p.30) and are based on some rather unrealistic technical conditions. However, the values obtained in this way seem to act as upper bounds in the simulations carried out by West (2001, p.31).

Table 9: Predictive accuracy tests

		1-step ahead forecasts			2-step ahead forecasts	
$\mathsf{M}_i \setminus \mathsf{M}_j$	VAR	CSC	IRS	VAR	CSC	IRS
ARIMA	1.306 (0.200)	1.325 (0.195)	0.800 (0.429)	1.193 (0.241)	0.981 (0.334)	−2.268 (0.030)
VAR		−0.342 (0.735)	−0.628 (0.534)		−0.165 (0.870)	−2.245 (0.032)
CSC			−0.383 (0.704)			−3.127 (0.004)
		ARIMA vs VAR (3 to 12 steps ahead)				
steps		3	4	5	6	7
		1.053 (0.299)	2.003 (0.053)	1.823 (0.076)	1.230 (0.226)	2.299 (0.027)
steps		8	9	10	11	12
		2.450 (0.019)	2.701 (0.010)	2.267 (0.029)	4.911 (0.000)	6.829 (0.000)

Modified Diebold-Mariano tests based on $d_t = |\hat{e}_{it}| - |\hat{e}_{jt}|$. The DM^* statistics and their p-values under the null (in brackets) are reported.

In order to evaluate the forecasting accuracy of the various models, in Table 9 we report the results of the comparisons carried out on the different projections. The table shows that our model on average produces more precise forecasts than the others. However, the comparisons suggest that the difference is statistically significant only with respect to the two-step ahead forecasts released by IRS, and with the 4-to-12-step ahead ARIMA forecasts.[17]

The tests of forecast encompassing reported in Table 10 show less clear-cut results. From our viewpoint it seems relevant to note that, though the average quality of our forecasts is superior to that of IRS projections, nevertheless these embody some pieces of information that could potentially improve both the one-step and the two-step ahead VAR forecasts. Note also that the converse apply even more strongly: in fact our VAR forecasts incorporate information that would be useful for improving IRS projections (and this, in the light of the previous findings, is an expected result). CSC predictions do not encompass ours. On the other hand, it is not entirely clear if our VAR forecasts encompass those elaborated by CSC: taking into account possible size distortions, the results seem to suggest that encompassing probably takes

[17]The test is significant (at the 10% significance level) for the 4- and 5-step ahead forecasts: it is not significant for the 6-step ahead forecasts (due to a single badly mispredicted value at 1999:12).

Table 10: Tests for forecast encompassing

$M_i \setminus M_j$		ARIMA	VAR	CSC	IRS
		1-step ahead forecasts			
ARIMA			1.805 (0.079)	1.395 (0.172)	1.979 0.056
VAR		2.246 (0.031)		1.832 (0.076)	2.745 (0.009)
CSC		1.188 (0.244)	2.945 (0.006)		1.475 (0.150)
IRS		2.085 (0.044)	3.603 (0.000)	1.524 (0.138)	
		2-step ahead forecasts			
ARIMA			2.088 (0.044)	3.460 (0.002)	3.222 (0.003)
VAR		1.999 (0.053)		2.094 (0.044)	2.405 (0.022)
CSC		2.084 (0.045)	2.822 (0.008)		−0.521 0.606
IRS		3.118 (0.004)	3.777 (0.000)	2.383 (0.023)	
		ARIMA vs VAR (3-12 steps)			
	3	4	5	6	7
	2.431 (0.020)	2.637 (0.012)	2.229 (0.032)	1.947 (0.059)	2.761 (0.009)
	8	9	10	11	12
	2.661 (0.014)	2.520 (0.016)	2.340 (0.025)	3.576 (0.000)	3.282 (0.002)
		VAR vs ARIMA (3-12 steps)			
	3	4	5	6	7
	2.148 (0.038)	2.089 (0.044)	0.960 (0.343)	1.189 (0.242)	0.737 (0.466)
	8	9	10	11	12
	0.622 (0.538)	1.164 (0.252)	1.660 (0.105)	0.762 (0.451)	0.197 (0.845)

Modified Diebold-Mariano tests based on $d_t = \hat{e}_{it}(\hat{e}_{it} - \hat{e}_{jt})$. The DM^* statistics and their p-values under the null (in brackets) are reported. The null hypothesis is that the forecasts produced by model M_i (column-wise) encompass those produced by model M_j (row-wise).

place for the one-step ahead predictions. The comparisons with the ARIMA indicate that the benchmark do possibly encompass the VAR forecasts only at the shortest horizon (one-step): on the contrary the predictions from the VAR encompass those from the ARIMA from the five-step ahead onward.

5 Using the forecasts to improve trend-cycle estimates

This section illustrates some possible uses of the results of the forecasts derived from the model described in Section 3. While the main purpose of the model lies in the pure forecast of the raw industrial production index, nevertheless its results can be used to improve the construction of a cyclical indicator, reducing the revisions implied in its calculation. Moreover, the forecasts can be of some help in reducing the delay with which a turning point is recognized.

5.1 Characterization of the problem

Let us consider the series IPI as composed by three (unobserved) components:

$$IPI_t = T_t + S_t + I_t. \tag{6}$$

The three elements are the trend (T_t), the seasonal (S_t) and the irregular component (I_t). The first represents the long term evolution of the series, together with oscillations associated with the business cycle. It should then be more correctly defined *trend-cycle*. The seasonal component represents movements which repeat themselves on a regular basis every year, while the irregular is a stationary, highly volatile and unpredictable component.

When looking for a cyclical indicator, one is normally interested in eliminating the seasonal and the irregular component, leaving only the trend-cycle. To estimate the latter, many criteria have been proposed. Some of them, such as the X-12-ARIMA seasonal adjustment procedure (Findley *et al.*, 1998), do not provide a statistical model for the components: others do provide an explicit characterization of the components. Among the latter, there are the structural time series approach (Harvey, 1989), and the ARIMA model-based approach (see Maravall, 1995, and the references therein).

The ARIMA model-based approach will be retained here, because of some appealing features. It is in fact quite simple to apply,[18] and is consistent

[18] We apply it using the software TRAMO-SEATS (Gómez and Maravall, 1998).

with the seasonally adjusted figures officially provided by ISTAT (see ISTAT, 1999).

As already pointed out, the trend-cycle component contains also movements associated with the business cycle. In this paper we are not interested to disentangle long term trend and business cycle oscillations, given that the data we are dealing with feature frequent cycles in the classical sense, so that no detrending is necessary. Indeed, whenever we refer to business cycle in this section, we do so in the sense of *classical cycle* and not in the sense of *growth cycle* (which consider detrended series). In addition, we are also interested in the absolute level of the series. Finally, the separation of the trend from the cycle has often been questioned on statistical grounds (see, *e.g.*, Canova, 1998). At any rate, if the forecasts produced by our model are useful to improve the construction of a trend-cycle indicator, they are likely to be equally useful with respect to the calculation of a purely cyclical indicator.

5.2 Main features of the trend extracted by TRAMO-SEATS

The ARIMA model-based approach implies the identification and estimation of an ARIMA model for the observed series, with some possible deterministic components, such as trading days effects and outliers. ARIMA models for the components are then derived, using some identifying assumptions; among them there is the independence of the components in (6). Below, the main steps of the procedure are summarized.

Consider, for the aggregate monthly series y_t the following seasonal ARIMA representation:[19]

$$(1-L)\left(1-L^{12}\right)y_t = (1+\theta_1 L)\left(1+\theta_{12}L^{12}\right)\varepsilon_t \tag{7}$$

where L is the lag operator, and $\varepsilon_t \sim \mathsf{NIID}(0,\sigma_\varepsilon^2)$. The autoregressive part of the model can be factorized as follows:

$$(1-L)\left(1-L^{12}\right) = (1-L)^2\left(1+L+L^2+...+L^{11}\right). \tag{8}$$

The first element in the right hand side of (8) implies two unit roots at the zero frequency of the spectral representation of y_t, while the second factor

[19] This example has been developed using the so called Airline model, which for monthly data is an ARIMA $(011)(011)_{12}$. It has been chosen, both because it is the default model used by TRAMO-SEATS and it is the one actually used by ISTAT to seasonally adjust the IPI series.

has eleven roots centered at the seasonal frequencies $2k\pi/12$, $k = 1, 2, ..., 11$. The first two unit roots, which are associated with the long term evolution of the series, can then be assigned to the trend component, while the other eleven are associated with the seasonal component. Restricting our attention to the trend, we obtain that it is given by the following ARIMA model:

$$(1-L)^2 T_t = \theta_T(L)\,\varepsilon_{T,t} \tag{9}$$

with $\varepsilon_{T,t} \sim \mathsf{NIID}(0, \sigma^2_{\varepsilon_T})$ and $\theta_T(L)$ a lag polynomial of order two. Nevertheless, the precise expression for the trend is not yet obtained, because there are many, possibly infinite, decompositions consistent with the aggregate model (7), which differ in the values of the coefficients in $\theta_T(L)$. The criteria used by TRAMO-SEATS to identify the components is to make them *canonical, i.e.* specifying them as free of noise as possible. The trend obtained is, therefore, the smoother trend among those obtainable from model (9) given the aggregate model (7). The canonical condition, together with the independence assumption, guarantees a unique decomposition.

Figure 7 plots the trend of the industrial production index estimated by TRAMO-SEATS. It clearly contains the long term evolution of the series, but also short term cyclical movements, up- and downturns.

Optimal (in a mean squared sense) estimation of the trend can be obtained using the Wiener-Kolmogorov filter (Maravall, 1995). Denote with $S(L) = (1 + L + L^2 + ... + L^{11})$, and with $\theta(L)$ the lag polynomial in (7). Considering an infinite realization of y_t, $Y = \{y_t\}_{t=-\infty}^{\infty}$, the trend estimator $\widehat{T_t}$ is given by the following expression:

$$\widehat{T_t} = \mathsf{E}(T_t|Y) = \frac{\sigma^2_{\varepsilon_T}}{\sigma^2_{\varepsilon}} \frac{\theta_T(L)\,\theta_T(L^{-1})S(L)\,S(L^{-1})}{\theta(L)\,\theta(L^{-1})} y_t = \nu(L)y_t. \tag{10}$$

From equation (10) it is evident that the estimator of the trend is obtained applying to the original series a symmetric, bidirectional, infinite filter $\nu(L)$. Moreover, invertibility of $\theta(L)$ ensures that the filter is convergent. This allows to render the procedure operational, approximating the infinite filter by truncation. Let now assume that the truncated filter length is equal to $2r + 1$, so that r observations are lost at the end of the observed series. The usual solution is to extend the latter with the predictions coming out from the ARIMA model (7). This means that we actually have a sequence of preliminary estimates of the trend $\mathsf{E}(T_t|y_{t+i})$ $(i = 0, 1, ..r)$ which gradually converge to the final one as long as predictions are replaced by true values.

In this section we claim that the use of the forecasts coming out from the model described in Section 3 dramatically improves the preliminary estimate of the trend of IPI, thus making it a much better device in order to monitor

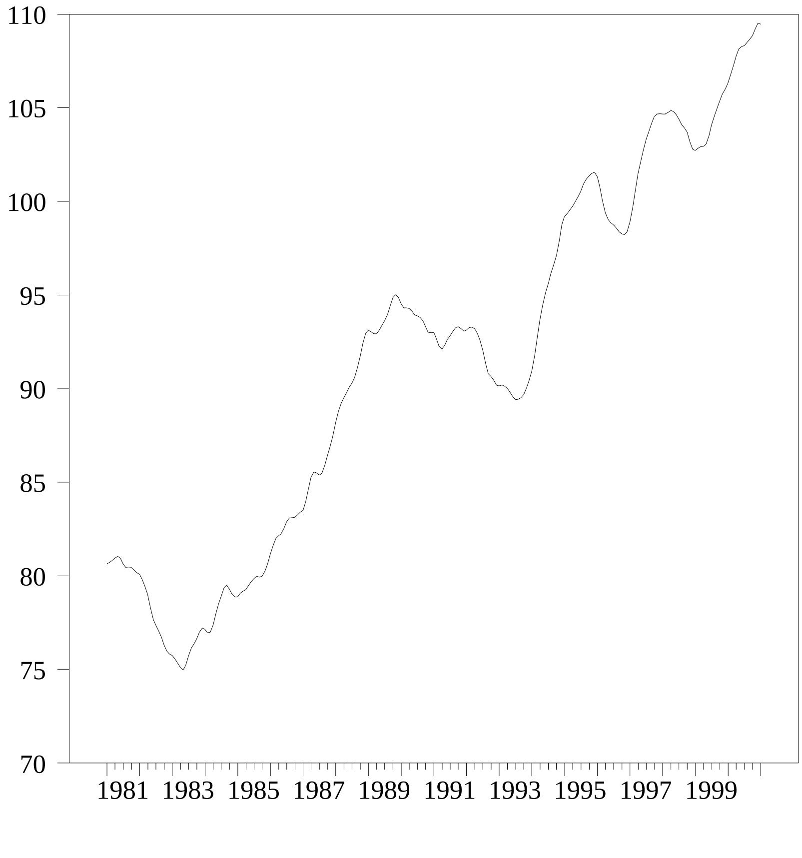

Figure 7: Estimated trend component of the industrial production index.

the evolution of this variable. To justify the need for such an exercise, we rely on Bruno (2001), that shows that the revisions of the trend component of IPI can be unacceptably large. In particular, Bruno (2001) shows that the trend extracted by TRAMO-SEATS, while representing a good historical representation of the cyclical development of the industrial production index, is characterized by a deep worsening of its performance near the end of the series, which is the point economic and business analysts are mainly concerned with.

5.3 Revisions in the trend series

In order to check the importance of revisions in trend the estimates, and to evaluate the advantages deriving from using our model's forecasts instead of the standard routine, we perform a historical simulation from January 1996 to February 2001, estimating the trend component with TRAMO-SEATS at every period, for the original series and for series extended with 3, 6 and 12-step ahead forecasts obtained from our VAR model.

The measure used to illustrate the revisions process is the following. Let $\widehat{T}_{t|t+k}$ be the estimate of the trend component at time t when a series of length $t+k$ $(k \geq 0)$ is observed: the so called *concurrent estimate* is obtained when $k = 0$. The quantity

$$\hat{r}_k = \widehat{T}_{t|t+k} - \widehat{T}_{t|t+k-1} \qquad k = 1, 2, ... \tag{11}$$

represents, for every k, the monthly revision in $\hat{T}_t$, k months after the concurrent estimate.

We compute (11) for every month from January 1996 onward, obtaining a distribution of revisions for every k ranging from 1 (with 61 observations) to 61 (just one observation). In practice we are usually interested in, say, $k \leq 12$. We can therefore derive summary statistics of the monthly revisions: in particular it is interesting to check their variances, to see if the use of our forecasts improves the revision process. Figure 8 shows clearly how effective is the improvement in the revision pattern using the forecasts from our model. The black bar (labelled 'Original') is the variance associated with the monthly revision after k periods (x-axis) using the standard procedure: the variance behavior is characterized by a sharp decrease after the first five months, when it becomes negligible. The use of three-step ahead forecasts from our model reduces the variance of revision of about 35% during the first three periods, and of about 30% during the fourth and fifth month. Six-step ahead forecasts improve on this result, reducing by 50% the variance of revisions in the first three periods. Using 12-step ahead forecasts does not seem to give further significant gains.

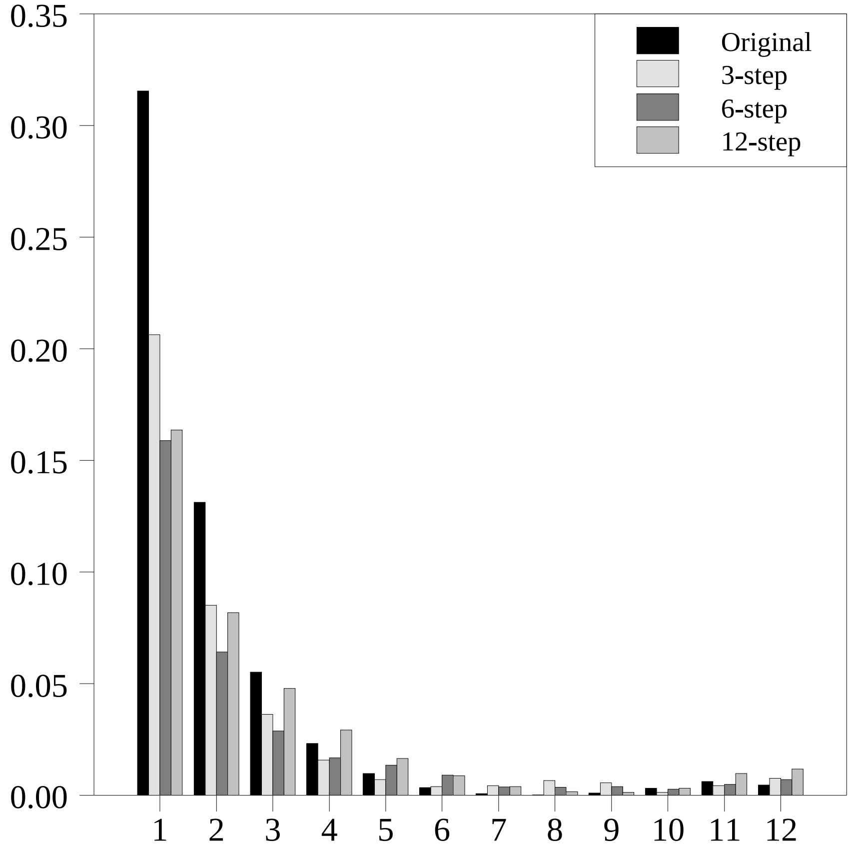

Figure 8: Variances of revisions of the trend.

5.4 Detection of turning points

In order to assess how important is the improvement in the revision process showed in the previous sub-section, it is possible to analyze if it helps, for example, in the timely detection of turning points. In order to do this we perform again a historical simulation, from 1996 onwards, applying a routine to detect the turning points in the trend-cycle component estimated over the original series and over that extended with, 3, 6, and 12 forecasts coming from our model. As a reference, we considered definitive turning points those identified with the observed series ending in February 2001.

It is important to stress that our aim is not to find out the best approach to signal turning points, nor to forecast them; we rather want to verify if the use of our forecasts helps in detecting past turning points earlier. In fact, turning points are essentially non-linear phenomena, and non-linear methods are natural candidates to forecast them (Camacho and Perez-Quiros, 2000). On the contrary, here we want to show that our forecasts can be used to recognize turning points earlier.[20]

The turning points identified by the procedure over the period 1996-2000 using the trend estimated over the actual data up to February 2001, are four, two peaks and two troughs: they are few, but going back further would have led to a too pronounced loss of data in order to estimate our model. The historical simulation is performed, again, reproducing as closely as possible a real world situation, that is re-estimating each month the model, leaving its structure unchanged. Table 11 shows the main results. The dates in the first column represent the turning points estimated as of February 2001, while the others are the months where the turning points were first detected. Dates in brackets represent the estimated locations of the turning points when first identified.

The mean lag in the detection of turning point with the original trend series is about 10 months. The use of our model's three-step forecasts improves the detection of the turning points, leading to a mean lag of 6.6 months. A further improvement is obtained with a longer forecasting horizon. With a six-step ahead forecast the turning point is detected, on average, after 4.3 months, while using a twelve step-ahead forecast it reduces to just 1.3 months. Figure 9 shows four cases in the neighborhood of the actual turning points where the performance of the trend obtained by extending the industrial production with 12-step forecasts of the VAR model (labelled 'Forec.' in the figure) is compared with the ordinary output of TRAMO-SEATS (labelled 'Concurrent'). The first appears to follow more closely the final estimate of

[20]The identification of turning points is carried out following the procedure proposed by Bry and Boschan (1971).

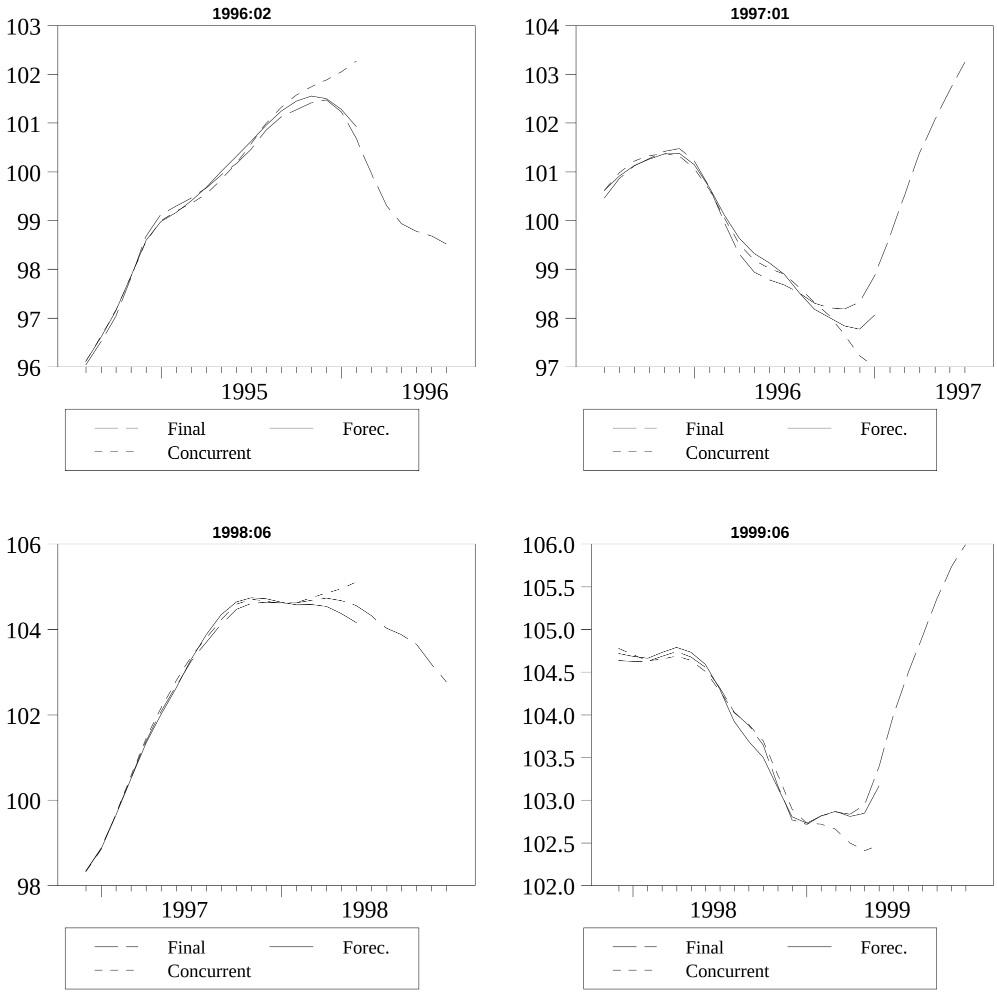

Figure 9: Turning points detection under the standard procedure ('Concurrent') and using 12-step ahead VAR forecasts ('Forec.'). The date on the top of each panel indicates the last observation used in the trend estimation. The trend estimate based on actual data up to February 2001 ('Final') is also reported for comparison.

Table 11: Detection dates of turning points with different forecasts

Turning points	TRAMO-SEATS	3-step ahead	6-step ahead	12-step ahead
1995:12 (p)	96:09 (95:10)	96:05 (95:10)	96:03 (95:10)	96:02 (95:11)
1996:11 (t)	97:08 (96:12)	97:05 (96:12)	97:04 (96:12)	96:12 (96:12)
1998:04 (p)	99:02 (97:12)	98:10 (97:12)	98:07 (97:11)	98:04 (98:06)
1999:01 (t)	99:12 (99:01)	99:10 (99:01)	99:07 (99:01)	99:03 (99:04)
Mean lag	9.9	6.6	4.3	1.3

The table reports the dates of first detection of the turning points: ”p” denotes a peak, ”t” a trough. In the first column are listed the dates of the turning points as estimated using the whole time series up to February 2001. In parenthesis are reported the turning point locations as estimated at the detection date. Mean lag is the average lag of the detection.

the trend (’Final’), obtained using the observed time series as of February 2001. Visual inspection confirm very clearly the results illustrated in Table 11 and in Figure 8, that is the gain in precision in the trend estimate, in particular around turning points.

The procedure of Bry and Boschan in this context proves particularly robust against false signals, which never occur in our sample. Some problems emerge for the detection of the peak in 1998:4, which is sometimes located at the end of 1997; this is due to the ”flatness” of the industrial production during that period. In addition, with the use of the twelve-step ahead forecast the turning point 1999:1 is identified the first time in January 1999, but not in the subsequent month: this is why in the table we report the value of March (from that month onward this turning point is always reported).

6 Concluding remarks

In this paper we propose a simple VAR model to forecast Italian industrial production. We test for the predictive accuracy of our model over a fairly long forecast evaluation sample. We show that our VAR predictions outperform those produced on the basis of a robust ARIMA model, are on average at least as good as the survey-based projections elaborated by CSC, and more accurate than those deriving from the IRS econometric model. Furthermore, we show that using the VAR we are able to produce reliable forecasts on longer horizons. The forecast encompassing tests highlight that the different predictions embody different pieces of information that could be exploited to obtain even better forecasts. As long as one is interested only in forecasting horizons of at most two periods, this opens the possibility of investigating the opportunity of combining the forecasts: given that one of our goals is to produce multi-step dynamic forecasts, we do not pursue this route in the present paper.

We argue that obtaining good forecasts is essential to derive a reliable cyclical indicator using signal-extraction (smoothing) techniques. We show that this is the case by comparing the variance of revisions of a cyclical indicator estimated using our VAR's forecasts with that of the same indicator estimated using standard procedures: the information embodied in our predictions halves the uncertainty in the concurrent estimate of the cyclical indicator. This is also fundamental to timely detect turning points: the average gain in the delay with which a turning point is detected when using our forecasts is about nine months! We guess that a clear indication to pratictioners and economic analysts arise from these results: multi-step dynamic forecasts can improve substantially on the perception we can gain not only on the future, but also on the current phase of the economy.

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