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Keskusteluaiheita
Discussion papers

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A NOTE ON FORECAST EVALUATION AND CORRECTION

No 309 27.12.1989

ISSN 0781-6847

ILMAKUNNAS, Pekka: A NOTE ON FORECAST EVALUATION AND CORRECTION. Helsinki: ETLA, Elinkeinoelämän Tutkimuslaitos, The Research Institute of the Finnish Economy, 1989. 13 p. (Keskusteluaiheita, Discussion Papers, ISSN 0781-6847 ; 309)

ABSTRACT: The question whether macroeconomic forecasts are useful as inputs to forecasting models for other variables is studied. A simple t-test is suggested for testing whether the use of a forecast is better than the use of the mean of the variable to be forecasted. The test is shown to be related to some other criteria that have been suggested, but has the advantage that critical values are easily available. Also some methods for forecast correction by end users are discussed and their limitations are studied.

KEY WORDS: Forecast accuracy, linear correction of forecasts

1. Introduction

When published macroeconomic or industry-level forecasts are used as explanatory variables in forecasting models for less aggregate variables, e.g. regional or firm-level, one faces the problem that these forecasts may be fairly inaccurate and therefore not necessarily useful. It would be important to be able to test whether a forecast is accurate enough in such situations. For example, in a recent survey of forecasting methods in marketing, Armstrong et al. (1987) posed the question "what are acceptable levels of forecast errors for environmental inputs to market forecasting models?" as one issue where more research is needed. This paper attempts to provide an answer to that question.

Ashley (1983, 1985) suggested a new criterion for the evaluation of forecasts in such situations. The idea is to compare use of the forecast in the model and use of the mean of the variable as an alternative forecast. In some cases this can be interpreted as omission of the variable in question from the model in the forecast period. For example, when the model is expressed in terms of deviations from the means, this is the same as using a zero forecast of the deviation.

If the published forecast is poor, some kind of correction may be applied by the forecast user for improving it. Typically this kind of revisions are based on past forecast performance, either informally or by estimating adjustment factors. Ashley (1985) suggested improvement of a forecast by combining it with the mean of the variable to be forecasted.

The purpose of this note is to show that both Ashley's test criterion and forecast correction are closely related to a familiar forecast evaluation approach. Also a simpler form of the test is derived. In the concluding section reservations about this kind of forecast revision are discussed.

2. Testing the usefulness of forecasts

Consider a simple model $y_t = \alpha + \beta x_t + e_t$, $t = 1,\dots,T-1$, where y is e.g. a firm's sales and x is some activity variable, e.g. disposable income. To use this model for forecasting the firm's sales for period T, a forecast for the macro variable x_T is needed. α and β are assumed to be known, although this does not affect the results.

For simplicity, it is better to consider the model when the constant has been removed by expressing the variables as deviations from their means: the period T model is

$$y_T - E(y) = \beta(x_T-E(x)) + e_T. \tag{1}$$

Consider as a first alternative the use of a published macro forecast $\hat{x}_T$ in (1). The forecasted value of y_T is $\hat{y}_{T1} = E(y) + \beta(\hat{x}_T-E(x))$, which has mean squared error (MSE)

$$MSE_1 = E(y_T-\hat{y}_{T1})^2 = \beta^2E(x_T-\hat{x}_T)^2 + \sigma_e^2 = \beta^2MSE(\hat{x}_T) + \sigma_e^2. \tag{2}$$

An alternative is to forecast x_T to be equal to E(x), the mean of x. The corresponding forecast for y is $\hat{y}_{T2} = E(y)$, which has MSE

$$MSE_2 = E(y_T-\hat{y}_{T2})^2 = \beta^2E(x_T-E(x))^2 + \sigma_e^2 = \beta^2\sigma_x^2 + \sigma_e^2. \tag{3}$$

Hence use of the mean $E(x)$ rather than the forecast $\hat{x}_T$ is justified if $k = MSE(\hat{x})/\sigma_x^2 > 1$, which result was derived in Ashley (1983). For optimal forecasts $k < 1$ always holds, since then $E(x) = E(\hat{x})$, $Cov(x,\hat{x}) = \sigma_{\hat{x}}^2$, and hence $MSE(\hat{x}) = \sigma_x^2 - \sigma_{\hat{x}}^2 < \sigma_x^2$ (Granger and Newbold, 1986, p. 131).

A disadvantage of this approach is that it is difficult to find a significance level or critical values for the test. However, Ashley (1985) notes that since in practice the parameter β has to be estimated and on the other hand, the forecasting model is likely to have other explanatory variables besides x, leaving out x (or using E(x) as the forecast) may be justified even when k is close to one. One may argue, however, that the properties of the forecast might not stay constant over time so that a test based on past forecast performance is not quite reliable. This issue is discussed further in the concluding section, but here an alternative test is derived which is more conservative in the sense that the usefulness of the forecast is less likely to be rejected. Consider a third alternative forecast $\hat{x}_T$ + bias = $\hat{x}_T + E(x-\hat{x})$, i.e. a forecast where a known additive bias in the original forecast has been removed (cf. Ashley, 1985). The forecast of y_T is now $\hat{y}_{T3} = E(y) + \beta(\hat{x}_T - E(\hat{x}))$, which has MSE

$$
\begin{aligned}
MSE_3 = E(y_T-\hat{y}_{T3})^2 &= \beta^2 E(x_T-E(x)-\hat{x}_T+E(\hat{x}))^2 + \sigma_e^2 \\
&= \beta^2(\sigma_x^2 + \sigma_{\hat{x}}^2 - 2Cov(x,\hat{x})) + \sigma_e^2 \\
&= MSE_2 + \beta^2(\sigma_{\hat{x}}^2 - 2Cov(x,\hat{x})). \qquad (4)
\end{aligned}
$$

This shows that $MSE_3 < MSE_2$ if $Cov(x,\hat{x})/\sigma_{\hat{x}}^2 = b > 1/2$, where b is the slope coefficient from a regression of x on $\hat{x}$ and a constant. On the other hand, it is easy to show that MSE_1 exceeds MSE_3 by $\beta^2(E(x)-E(\hat{x}))^2$.

The interpretation of these results is the following. When x is regressed on $\hat{x}$ and a constant, for optimal forecasts the constant should be zero and the slope coefficient b should be 1. If $1 < b < \infty$, $\hat{x}_T$ + bias has a downwards slope bias, but it is still more accurate than the mean E(x) as a forecast. If $b < 1$, the forecast $\hat{x}_T$ + bias has an upwards slope bias, and if $b < 1/2$, $\hat{x}_T$ + bias is too inaccurate compared to E(x) to be useful in forecasting y_T. If $b < 0$, even the direction of changes (assuming that the forecasts refer to changes) are typically forecasted incorrectly. In the remainder of this paper it will be assumed that b is positive.

Especially, if the variables are changes, a large value of b may be interpreted as resulting from underestimation of the magnitude of changes although the signs of the changes are correctly forecasted. However, "underestimation of changes" or downward slope bias can result simply from the fact that optimally the variance of the forecasts should be smaller than that of the actual values (Granger and Newbold, 1986, p. 286). Correspondingly, $b < 1$, or "overestimation of changes" may imply inefficient forecasts for which $\sigma_x^2 < \sigma_{\hat{x}}^2$.

Since $MSE_3 \leq MSE_1$ always holds, it follows that when the mean forecast E(x) is superior to the additive bias corrected forecast $\hat{x}_T - E(x-\hat{x})$, it is also superior to the original forecast $\hat{x}_T$. There may, however, be cases where $MSE_3 < MSE_2 < MSE_1$, so that one cannot choose between E(x) and $\hat{x}_T$ on the basis of the $b < 1/2$ criterion.

Compared to Ashley's test the present test has the advantage that it can easily be performed using a one-sided t-test. In practice a complication may arise from the fact that the error term in the regression of x on $\hat{x}$, i.e. the forecast error, is not necessarily uncorrelated with $\hat{x}$. In fact, $E(\hat{x}(x-\hat{x})) = 0$ holds for optimal forecasts, but not necessarily for sub-optimal ones. Therefore instrumental variable estimation may be applied. In addition, the errors of especially several periods ahead forecasts can be serially correlated because of information lags and data revisions after the forecast was made (Brown and Maital, 1981, Berger and Krane, 1985). In this case the testing of b should be done using a consistent estimate of its standard error.

If the model is formulated so that the variables refer to changes, another useful benchmark forecast is a zero, or no-change forecast. Decomposing $MSE(\hat{x}_T) = E(x_T-\hat{x}_T)^2$, it can be seen that $MSE(0) = E(x^2)$ is smaller than $MSE(\hat{x}_T)$ if $b' = E(x\hat{x})/E(\hat{x}^2) < 1/2$, i.e. the slope coefficient from a regression of x on $\hat{x}$ without a constant is smaller than 1/2. If $b' > 1/2$, the original forecast is better than the no-change forecast.

It can also be noted that, when the variables are changes, the above tests are closely related to some other test statistics. Theil (1966) suggested using $U = (\sum_i(x_i-\hat{x}_i)^2/n)^{1/2}/(\sum_i x_i^2/n)^{1/2}$ as a measure of forecast accuracy. The U statistic compares the sample analogues of the root mean squared errors of forecast $\hat{x}$ and a no-change forecast. If $E(x) = 0$, i.e. the no-change forecast is unbiased, this can be written as $U = (MSE(\hat{x}_T))^{1/2}/(\sigma_x^2)^{1/2} = k^{1/2}$. Both Ashley's test and Theil's U-statistic can hence be used for testing whether a published forecast of an exogenous variable is accurate enough compared to a no-change forecast. In fact, the $U < 1$, $k > 1$ and $b' > 1/2$ criteria all give the same ranking to $\hat{x}_T$ and a no-change forecast in the special case $E(x) = 0$. However, $U < 1$ and $b' > 1/2$ give the same result even when the no-change forecast is biased.

Nelson (1961) has suggested another test criterion for choosing between a forecast and the mean of the variable to be forecasted, $R^2 = 1 - E(x_T-\hat{x}_T)^2/E(x_T-E(x))^2 = 1 - MSE(\hat{x}_T)/\sigma_x^2 = 1-k$. It measures the proportion of the total variation in x, i.e. the proportion of the variance of x, which is explained by the forecast; the unexplained part of the variation is measured by the MSE of the forecast. Also Pierce (1975) has suggested a R^2-type measure of forecast evaluation in the case where the forecast $\hat{x}_T$ is obtained from a time series model. As shown above, R^2 is directly related to the k-statistic. If the mean E(x) is used as a forecast, $R^2 = 0$, and for perfect forecasts, $R^2 = 1$. Assuming that $E(\hat{x}) = E(x)$, R^2 can be written also as $R^2 = 2(b-1/2)\sigma_{\hat{x}}^2/\sigma_x^2$, which relates the $R^2 > 0$, criterion to the $b > 1/2$ criterion. Theoretically, inequality $1 \geq R^2 \geq 0$ should hold, which requires that the forecast has to be optimal in the sense that $MSE(\hat{x})$ is always smaller than σ_x^2. In practice, there is no guarantee for this to hold, since forecasts may be biased and inefficient.

Nelson derived the $R^2 > 0$ criterion in an example from the theory of the firm. The similarity to the econometric results shown above follows the quadratic structure of both problems: here a MSE criterion has been used for comparing the forecasts, whereas Nelson assumed a quadratic profit function for the firm (see also Ilmakunnas (1987)).

3. Optimal correction of forecasts

When the macroeconomic forecast does not satisfy the $b > 1/2$ or $k < 1$ conditions, it may be possible to combine the forecast with some other forecast to reduce the MSE.

Ashley (1985) shows that this can be done by taking a linear combination of $E(x)$ and the bias adjusted forecast $\hat{x}_T + E(x-\hat{x})$. An interpretation of his results is given in terms of the methods of evaluating forecast accuracy that were discussed in section 2. The combined forecast is $x_T^* = \lambda E(x) + (1-\lambda)(\hat{x}_T + E(x-\hat{x})) = E(x) + (1-\lambda)(\hat{x}_T - E(\hat{x}))$, which is unbiased. The MSE of x_T^* is therefore equal to the variance of the forecast error $x_T - x_T^* = x_T - E(x) - (1-\lambda)(\hat{x}_T - E(\hat{x}))$:

$$MSE(x_T^*) = \sigma_x^2 + (1-\lambda)^2\sigma_{\hat{x}}^2 - 2(1-\lambda)Cov(x,\hat{x}). \qquad (5)$$

Minimization of (5) with respect to λ yields optimal value $\lambda^* = 1-Cov(x,\hat{x})/\sigma_{\hat{x}}^2 = 1-b$. The optimal combined forecast is $x_T^* = (1-b)E(x)+ b(\hat{x}_T-E(x-\hat{x}))$ $= a + b\hat{x}_T$, where $a = E(x) - bE(\hat{x})$. This is the same as Theil's (1966) optimal linear correction of forecasts: x is regressed on $\hat{x}$ and a constant and the estimated slope and intercept are used for correcting the original forecast. Using Theil's forecast error decomposition, the MSE of a forecast can be broken down to variance, slope and bias portions. The optimal linear correction removes the bias and slope parts of the error within the sample period.

The above result is the same as that obtained when $E(x)$ is combined with $\hat{x}_T$ without correcting $\hat{x}_T$ for bias and without constraining the weights to add up to one. In this case the combined forecast is $x_T^* = \lambda_1 E(x) + \lambda_2\hat{x}_T$. However, $\lambda_1 E(x)$ can be replaced by a constant, a. The problem is then to choose a and λ_2 so that the MSE of x_T^* is minimized. The optimal values are $\lambda_2^* = b$ and $a=E(x)-bE(\hat{x})$.

These results can be compared to those in Section 2. When one is indifferent between using $\hat{x}_T+E(x-\hat{x})$ and $E(x)$ as a forecast of x_T, i.e. when $b = 1/2$, the optimal combination gives a weight 1/2 to each forecast.

Above the simple case of combining a macro forecast with the mean of the variable was discussed. In practice forecast users are likely to have available both own information of the economic situation and alternative macroeconomic forecasts. Combination of several forecasts has been extensively discussed in the forecasting literature and shown to improve forecasts. As one alternative, Granger and Ramanthan (1984) suggested estimating the weights to combine forecasts by regressing realized values on a constant and the alternative forecasts without any constraints on the slope parameters. Clearly, this is the same as Theil's optimal linear correction when there is only one forecast.

Clemen (1986) has criticized the Granger-Ramanathan procedure on the grounds that although it improves within-sample forecasts, out-of-sample forecasting may be more efficient when an adding-up constraint on the weights is used, i.e. they add up to one, and the model has no constant term. This combined forecast may be biased, but an efficiency gain arises from the fact that less parameters have to be estimated. Clearly this argument applies also to the present case. The corrected forecast $x_T^* = a+b\hat{x}_T$ is unbiased, but may be less efficient out-of-sample than the original uncorrected forecast $\hat{x}_T$. The latter can be interpreted as x_T^* with constraints a=0, b=1, so that the MSE of x_T^* is increased compared to the MSE of $\hat{x}_T$ due to the variances of the estimates of a and b.

When the forecast correction is used only if a preliminary test of forecast accuracy indicates that the forecast is not useful, i.e., $b < 1/2$, the resulting forecast is actually a pretest forecast. Evaluation of such forecasts is discussed in Ilmakunnas (1989). [Trenkler and Liski (1986) consider pretesting in the case of a combination of alternative forecasts, but use a different preliminary test.]

6. Conclusions

Revisions of published forecasts should be carefully assessed. It could be argued that if there are systematic biases in the forecasts, the forecasters themselves would correct them. Hence there would be no need for the users of forecasts to make corrections. Further, each forecast reflects specific information that the forecasters have of the economic situation, and therefore uncritical adjustments by users may make the results much worse.

It is possible that although within-sample forecasts are necessarily improved by forecast revisions, out-of-sample forecast performance deteriorates. Examples of such negative effects of mechanical use of Theil's correction are given in Bohara, McNown and Batts (1987), and in the case of a combination of several alternative forecasts in Clemen (1986) and Holden and Peel (1986). The latter also suggest that forecast errors are not likely to have a constant mean and variance, since the models and the personnel of the forecasting units change over time. This may make the forecast adjustments unreliable, if they are based on past forecast performance only. As noted above, there are also statistical reasons for not always correcting forecasts that "underestimate" changes.

On the other hand, there is evidence to show that macroeconomic forecasts do not use all information efficiently (e.g. Berger and Krane, 1985) and hence there may be room for corrections by users of forecasts. Successful examples of improvement of out-of-sample forecast performance by Theil's correction can be found in Ahlburg (1984) and Brandon, Fritz and Xander (1983), and, in the case of several combined forecasts, in Granger and Ramanathan (1984). Ashley's (1985) positive results are based on within-sample forecasts.

One case where systematic biases typically appear is business or consumer survey measures of expectations, since those answering the surveys are not likely to make an explicit evaluation of their own past forecasts and therefore may not correct possible biases. If such expectations are used as variables in a forecasting model, correction of the bias is hence useful (see e.g. Figlewski, 1983).

Another case where adjustments are clearly justified is when the user of forecasts does not have the same loss function as the forecaster. For example, it may be irrelevant to a forecaster whether a positive or a negative error is made, but a firm that uses the macro forecast in a sales forecasting model may regard an overestimate of sales as more costly than an underestimate. Therefore the firm may prefer a downwards biased forecast which involves an adjustment of a published macroeconomic forecast. Formally this can be rationalized with an asymmetric loss function (see e.g. Zellner, 1986).

More complicated cases in forecast evaluation arise when dynamics are taken into account. In econometrics it is not common to give an explicit weighting, or discounting, to future forecast errors. In contrast, a firm, when deciding on whether to use a forecast or not, has to consider also e.g. how investments made today, using currently available demand forecasts, affect expected profits in future periods, and how the future profits are valued. Therefore also adjustment costs and discount rate should be taken into account in forecast evaluation by the forecast users. In this case the simple rules for forecast evaluation developed in this paper may break down. This is discussed in more detail in Ilmakunnas (1987). Another example, where use of a forecast may not be preferable, although the conditions given in this paper hold, is the case of a risk averse firm.

Blair and Romano (1988) show that risk averters value even perfect forecasts less than risk-neutral firms, because the use of the forecast information may increase the variability of profits.

Acknowledgement

I am thankful to Yrjö Jahnsson Foundation for financial support.

References

Ahlburg, D.A.. (1984), "Forecast evaluation and improvement using Theil's decomposition", Journal of Forecasting 3, 345-351.

Armstrong, J.S., Brodie, R.J. and Mcintyre, S.H. (1987), "Forecasting methods in marketing: review of empirical research", International Journal of Forecasting 3, 355-376.

Ashley, R. (1983), "On the usefulness of macroeconomic forecasts as inputs to forecasting models", Journal of Forecasting 2, 211-223.

Ashley, R. (1985), "On the optimal use of suboptimal forecasts of explanatory variables", Journal of Business and Economic Statistics 3, 129-131.

Berger, A.N. and Krane, S.D. (1985): "The informational efficiency of Econometric model forecasts", Review of Economics and Statistics 67, 128-134.

Blair, R.D. and Romano, R.E. (1988): "The influence of attitudes toward risk on the value of forecasting", Quarterly Journal of Economics 103, 387-396.

Bohara, A., McNown, R. and Batts, J.T. (1987), "A re-evaluation of the combination and adjustment of forecasts", Applied Economics 19, 437-445.

Brandon, C., Fritz, R. and Xander, J. (1983), "Econometric forecasts: evaluation and revision", Applied Economics 15, 187-201.

Brown, B.W. and Maital S. (1981), "What do economists know? An empirical study of experts' expectations", Econometrica 49, 491-504.

Clemen, R.T. (1986), "Linear constraints and the efficiency of combined forecasts", Journal of Forecasting 5, 31-38.

Figlewski, S. (1983): "Optimal price forecasting using survey data", Review of Economics and Statistics 65, 13-21.

Granger, C.W.J. and Newbold, P. (1986), Forecasting Economic Time Series, 2nd ed., New York: Academic Press.

Granger, C.W.J. and Ramanathan, R. (1984): "Improved methods for combining forecasts", Journal of Forecasting 3, 197-204.

Holden, K. and Peel, D.A.: "An empirical investigation of combinations of economic forecasts", Journal of Forecasting 5, 1986, 229-242.

Ilmakunnas, P. (1987), "On the profitability of using forecasts", Economics Letters 25, 345-349.

Ilmakunnas, P. (1989), "Forecast pretesting", paper presented at Econometric Society European Meetings, Munich, Sept. 4-8, 1989.

Nelson, R.R. (1961), "Uncertainty, prediction and competitive equilibrium", Quarterly Journal of Economics 75, 41-62.

Pierce, D.A. (1975): "Forecasting in dynamic models with stochastic regressors", Journal of Econometrics 3, 349-374.

Theil, H. (1966), Applied Economic Forecasting, Amsterdam: North-Holland.

Trenkler, K., and Liski, E.P. (1986), "Linear constraints and the efficiency of combined forecasts", Journal of Forecasting 5, 197-202

Zellner, A. (1986): "Biased predictors, rationality and the evaluation of forecasts", Economics Letters 21, 45-48.

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