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Reevaluating the prudence of economic forecasts in the EU: The role of instrument persistence

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Summary

Christodoulakis and Mamatzakis (2009, *Journal of Applied Econometrics* 24, pp. 583–606) estimate the EU Commission loss preferences for selected economic forecasts of 12 EU Member States. They employ the generalized method of moments (GMM) estimation procedure proposed by Elliott et al. (2005, *Review of Economic Studies* 72, pp. 1107–1125) and find the forecasts to be somewhat optimistic on average. However, this note shows the GMM estimator to possess nonstandard limiting distributions when some of the instruments are highly persistent, which is the case with one of the instruments employed by Christodoulakis and Mamatzakis. Standard distributions are recovered in some interesting particular cases which are relevant in practice. A reexamination of the EU Commission loss preferences using methods robust to persistence and a dataset extended to 2017 reveals that, while the conclusions of the original study are, by and large, still justified, the EU Commission loss preferences have become more symmetric over the whole studied period.

1 | INTRODUCTION

The evaluation of forecasts enables an important feedback loop in applied econometrics. It has, e.g., become routine to test realized forecast error series for unbiasedness or lack of serial correlation. However, standard tests rely on the assumption that the relevant optimality criterion is mean squared error (MSE) based and may be misleading if the loss function is, say, asymmetric. It is thus of interest to learn about the shape of the loss function underlying a *given* sequence of forecasts. Employing the GMM framework of Elliott, Komunjer, and Timmermann (2005), Christodoulakis and Mamatzakis (2008, 2009) find asymmetric preferences in series forecasts of EU institutions and countries.

One limitation of the GMM framework is that the instruments are assumed to be stationary. This may seem benign, but many typical instruments are, on the contrary, quite persistent. In their assessment of EU Commission loss preferences, Christodoulakis and Mamatzakis (2009) use the lagged levels of inflation, unemployment, government balance, investment, and current account as instruments, which are often regarded to show stochastic trends (see Section 3 for

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concrete evidence on the persistence of these variables). This note therefore reexamines the forecast preferences of the EU Commission in light of possibly highly persistent instruments.

We show in Section 2 that the limiting distribution of the GMM estimator is non-Gaussian in general. The J -tests discussed by Elliott, Komunjer, and Timmermann (2005) lose the chi-square limiting distributions, and decisions based on χ^2 critical values lead to overrejections of the respective null hypotheses. However, we also find interesting particular cases where the limiting null distribution of the t -statistic of the asymmetry parameter is standard normal even under high instrument persistence.

In Section 3, we extend the dataset used by Christodoulakis and Mamatzakis (2009) to cover the period 1970–2017 and use our theoretical insights to reevaluate the analysis in a robust fashion. Christodoulakis and Mamatzakis find that apparent irrational forecasting behavior of the EU Commission can be well explained by asymmetries in the loss preferences and interpret the asymmetries in light of specific forecasting preferences (i.e., optimism versus pessimism). For year-ahead forecasts, they find a tendency for optimism, while the current-year forecasts are somewhat pessimistic. Using robust methods, we confirm the conclusions of Christodoulakis and Mamatzakis for the original data. However, when considering the extended dataset, we observe an overall reduction of deviations from symmetry in the EU Commission forecasts, as well as a somewhat reduced evidence of forecast irrationality.

The final section concludes, while technical proofs, additional simulations, and further empirical findings are gathered in a supplement to this note.

2 | GMM INFERENCE UNDER HIGH INSTRUMENT PERSISTENCE

2.1 | Estimation of asymmetry

The one-step-ahead optimal predictor of a series y_t is given by $\hat{y}_t = \operatorname{argmin}_{y^*} E_{t-1}(\mathcal{L}(y_t - y^*))$, where \mathcal{L} denotes the relevant loss function, which should be quasiconvex (see, e.g., Granger, 1999), and E_{t-1} denotes the expectation taken with respect to the conditional forecast density given the available information. The class of functions proposed by Elliott, Komunjer, and Timmermann (2005) is given by

$$\mathcal{L}(u) = (\alpha + (1 - 2\alpha)\mathbf{1}(u < 0))|u|^p, \quad (1)$$

where $\mathbf{1}(\cdot)$ is the indicator function. Let $u_t = y_t - \hat{y}_t$ denote the realized forecast error at time t . Under forecast optimality, the so-called generalized forecast error

$$\mathcal{L}'(u_t) = p(\alpha - \mathbf{1}(u_t < 0))|u_t|^{p-1} =: \tilde{u}_t(\alpha)$$

is a martingale difference sequence, $E_t(\tilde{u}_t) = 0$ (Granger, 1999). We sometimes write $\tilde{u}_t(\alpha) = \tilde{u}_t$ in the following, should dependence on α not be of essence.

To estimate α , one employs for given p (typically chosen as $p = 1$ or $p = 2$) a set of D instrument variables gathered in the vector \mathbf{v}_{t-1} . The instruments cannot improve the forecasts when these are optimal, so $E(\tilde{u}_t | \tilde{u}_{t-1}, \dots, \mathbf{v}_{t-1}, \dots) = 0$ under rationality. This implies D moment restrictions,

$$E(\mathbf{v}_{t-1} \tilde{u}_t(\alpha)) = pE(\mathbf{v}_{t-1}(\alpha - \mathbf{1}(u_t < 0))|u_t|^{p-1}) = 0,$$

leading, for the loss functions from Equation (1), to the GMM estimator

$$\hat{\alpha} = \frac{\hat{\mathbf{h}}' \hat{\mathbf{S}}^{-1} \left(\frac{1}{T} \sum_{t=2}^T \mathbf{v}_{t-1} \mathbf{1}(u_t < 0) |u_t|^{p-1} \right)}{\hat{\mathbf{h}}' \hat{\mathbf{S}}^{-1} \hat{\mathbf{h}}} = \alpha - \frac{1}{p} \frac{\hat{\mathbf{h}}' \hat{\mathbf{S}}^{-1} \left(\frac{1}{T} \sum_{t=2}^T \mathbf{v}_{t-1} \tilde{u}_t \right)}{\hat{\mathbf{h}}' \hat{\mathbf{S}}^{-1} \hat{\mathbf{h}}},$$

where $\hat{\mathbf{h}} = \frac{1}{T} \sum_{t=2}^T \mathbf{v}_{t-1} |u_t|^{p-1}$ and $\hat{\mathbf{S}} = \frac{1}{T} \sum_{t=2}^T \mathbf{v}_{t-1} \mathbf{v}'_{t-1} (\mathbf{1}(u_t < 0) - \hat{\alpha})^2 |u_t|^{2p-2}$. For estimation, an iterative procedure is required since $\hat{\mathbf{S}}$ depends on α via $\tilde{u}_t = \tilde{u}_t(\alpha)$. Under the assumptions of Elliott, Komunjer, and Timmermann (2005),

$$\sqrt{T}(\hat{\alpha} - \alpha) \xrightarrow{d} \mathcal{N}(0, \text{plim} \hat{V}),$$

where $\hat{V}^{1/2} = (\hat{\mathbf{h}}' \hat{\mathbf{S}}^{-1} \hat{\mathbf{h}})^{-1/2}$, and its probability limit exists and is positive. Moreover, J -statistics are available,

$$\mathcal{J} = \frac{1}{p^2} \left(\frac{1}{\sqrt{T}} \sum_{t=2}^T \mathbf{v}'_{t-1} \tilde{u}_t(\cdot) \right) \hat{\mathbf{S}}^{-1} \left(\frac{1}{\sqrt{T}} \sum_{t=2}^T \mathbf{v}_{t-1} \tilde{u}_t(\cdot) \right);$$

under stationarity assumptions, $\mathcal{J}_{\hat{\alpha}}$ (which may be interpreted as a rationality test, see Corollary 5 of Elliott, Komunjer, & Timmermann, 2005) has a limiting null $\chi^2(D - 1)$ distribution (so it requires the use of at least two instruments). If testing the null $\alpha = \alpha_0$ (with α_0 capturing the case of symmetric loss), \mathcal{J}_{α_0} has a limiting null $\chi^2(D)$ distribution, whereas the t ratio satisfies under the null $\alpha = \alpha_0$ ¹

$$\mathcal{T} = (\hat{\alpha} - \alpha_0) / \hat{V}^{1/2} \xrightarrow{d} \mathcal{N}(0, 1).$$

The normal and χ^2 null distributions hinge on whether a central limit theorem may be applied to the term $\frac{1}{\sqrt{T}} \sum_{t=2}^T \mathbf{v}_{t-1} \tilde{u}_t$. The summands $\mathbf{v}_{t-1} \tilde{u}_t$ form by construction a martingale difference sequence, but a “classical” central limit theorem does not apply when elements of \mathbf{v}_{t-1} are highly persistent (exhibit stochastic trends, e.g., random walks) and are thus nonstationary.²

2.2 | Highly persistent instruments

To allow for both stationary and highly persistent instruments, partition $\mathbf{v}_{t-1} = (\mathbf{v}'_{0,t-1}, \mathbf{1}, \mathbf{v}'_{1,t-1})'$, where the D_1 persistent instruments $\mathbf{v}_{1,t}$ satisfy the following technical requirement:

Assumption 1. Let \mathbf{N}_T be a diagonal matrix with $N_{T,ii} \rightarrow \infty$, and assume that there exists a continuous-time vector process $\mathbf{X}(s)$ such that the weak convergence $\mathbf{N}_T^{-1} \mathbf{v}_{1,[sT]} \Rightarrow \mathbf{X}(s)$ holds in a space of càdlàg functions endowed with a suitable norm.

This allows, among others, for near-integrated instruments (see, e.g., Campbell & Yogo, 2006, and references therein); but $\mathbf{X}(s)$ may be a (multivariate) Ornstein–Uhlenbeck (OU) process with nonzero initial condition, corresponding in discrete time to a near-integrated process initiated with a draw from the unconditional distribution (as employed by Müller & Elliott, 2003). Also, $\mathbf{X}(s)$ may be a nonstationary fractional Brownian motion (see, e.g., Maynard & Phillips, 2001, and references therein). The D_0 stationary instruments $\mathbf{v}_{0,t}$ do not exhibit stochastic trends:

Assumption 2. Let $(\tilde{u}_t, \mathbf{v}'_{0,t-1})' \in \mathbb{R}^{D_0+1}$ be a zero-mean stationary, ergodic, uniformly $L_{4+\delta}$ -bounded sequence for some $\delta > 0$, such that $E(\tilde{u}_t | \tilde{u}_{t-1}, \dots, \mathbf{v}_{t-1}, \dots) = 0$.

Theorem 27.14 in Davidson (1994) then implies

$$\frac{1}{\sqrt{T}} \sum_{t=2}^{[sT]} \begin{pmatrix} \tilde{u}_t \\ \mathbf{v}_{0,t-1} \tilde{u}_t \end{pmatrix} \Rightarrow \begin{pmatrix} \tilde{W}(s) \\ \tilde{\mathbf{W}}(s) \end{pmatrix},$$

¹As noted by Elliott, Komunjer, and Timmermann (2005) and Christodoulakis and Mamatzakis (2009), the studied null hypotheses are joint with the hypothesis that the loss function is of the asymmetric power type.

²Weak persistence, e.g., stationary long memory, would not hinder the use of a CLT to achieve limiting normality.

with $\tilde{W}(s)$ and $\bar{W}(s)$ Brownian motions. Take this to be joint with the weak convergence in Assumption 1, and define for later reference

$$\text{Cov} \begin{pmatrix} \tilde{W}(1) \\ \bar{W}(1) \end{pmatrix} = \begin{pmatrix} \text{Var}(\tilde{u}_t) & E(\mathbf{v}'_{0,t-1}\tilde{u}_t^2) \\ E(\mathbf{v}_{0,t-1}\tilde{u}_t^2) & E(\mathbf{v}_{0,t-1}\mathbf{v}'_{0,t-1}\tilde{u}_t^2) \end{pmatrix} = \begin{pmatrix} \sigma_u^2 & \boldsymbol{\gamma}'_0 \\ \boldsymbol{\gamma}_0 & \bar{\boldsymbol{\Omega}}_0 \end{pmatrix};$$

in case of conditional homoskedasticity of \tilde{u} , $\bar{\boldsymbol{\Omega}}_0 = \sigma_u^2 \text{Cov}(\mathbf{v}_{0,t-1})$ and $\boldsymbol{\gamma}_0 = \mathbf{0}$.

Having a constant in the vector of instruments is on the one hand common in practice, and on the other hand, the constant stands in for stationary instruments with nonzero mean since their purely stochastic component would be dominated in the limit (see the proof of Proposition 1). Hence requiring that $\mathbf{v}_{0,t-1}$ have zero mean does not imply any loss of generality. We then have.

Proposition 1. *Under Assumptions 1 and 2, it holds as $T \rightarrow \infty$ that*

$$\sqrt{T}(\hat{\alpha} - \alpha) \Rightarrow \frac{1}{p} \frac{\mathbf{H}'\mathbf{S}^{-1}\mathbf{U}}{\mathbf{H}'\mathbf{S}^{-1}\mathbf{H}} \quad \text{and} \quad \mathcal{T} \Rightarrow \frac{1}{p} \frac{\mathbf{H}'\mathbf{S}^{-1}\mathbf{U}}{\sqrt{\mathbf{H}'\mathbf{S}^{-1}\mathbf{H}}}$$

where

$$\mathbf{H} \equiv \begin{pmatrix} E(\mathbf{v}_{0,t-1}|u_t|^{p-1}) \\ E(|u_t|^{p-1}) \\ E(|u_t|^{p-1}) \int_0^1 \mathbf{X}(s) ds \end{pmatrix}, \quad \mathbf{U} \equiv \begin{pmatrix} \bar{W}(1) \\ \tilde{W}(1) \\ \int_0^1 \mathbf{X}(s) d\tilde{W}(s) \end{pmatrix}$$

and

$$\mathbf{S} \equiv \frac{1}{p^2} \begin{pmatrix} \bar{\boldsymbol{\Omega}}_0 & \boldsymbol{\gamma}_0 & \boldsymbol{\gamma}_0 \int_0^1 \mathbf{X}'(s) ds \\ \boldsymbol{\gamma}'_0 & \sigma_u^2 & \sigma_u^2 \int_0^1 \mathbf{X}'(s) ds \\ \int_0^1 \mathbf{X}(s) ds \boldsymbol{\gamma}'_0 & \sigma_u^2 \int_0^1 \mathbf{X}(s) ds & \sigma_u^2 \int_0^1 \mathbf{X}(s) \mathbf{X}'(s) ds \end{pmatrix}.$$

Proof: See Section A of the Supplement.

Proposition 2. *Under the assumptions of Proposition 1, it holds under the respective null hypotheses that, as $T \rightarrow \infty$,*

$$\mathcal{J}_{\hat{\alpha}_0} \xrightarrow{d} \frac{1}{p^2} \mathbf{U}'\mathbf{S}^{-1}\mathbf{U}$$

and

$$\mathcal{J}_{\hat{\alpha}} \xrightarrow{d} \frac{1}{p^2} \mathbf{U}' \left(\mathbf{S}^{-1} - \frac{1}{\mathbf{H}'\mathbf{S}^{-1}\mathbf{H}} \mathbf{S}^{-1} \mathbf{H} \mathbf{H}' \mathbf{S}^{-1} \right) \mathbf{U}.$$

Proof: Analogous to the proof of Proposition 1 and omitted.

Above all, it should be emphasized that the estimator $\hat{\alpha}$ is still consistent; it is only the higher-order properties that are affected by instrument persistence. Yet, in spite of the nonstandard distribution of $\hat{\alpha}$, the GMM estimator is \sqrt{T} -consistent only, and not superconsistent as might have perhaps been expected in the presence of stochastically trending instruments.

The fact that the limiting distributions of the J - and t -statistics change if some instruments are highly persistent has implications for the behavior of tests based on these statistics. Moreover, the distributions derived above depend on the properties of the process $\mathbf{X}(s)$, so it is quite difficult to provide practitioners with correct critical values, as these would

depend on the particular data-generating process (DGP). In fact, it may well happen that the relevant characteristics of $X(s)$ cannot even be consistently estimated. This is, for instance, the case with the mean reversion parameter of a near-integrated process (Phillips, 1987).³ In such situations, standard bootstrap schemes will fail too in general, as they cannot replicate the correct limiting distributions when the distributions depend on parameters that cannot be estimated consistently.

Yet, standard limiting behavior may be recovered in particular cases. The first case is when the instruments $\mathbf{v}_{1,t}$ are “exogenous” in such a way that *mixed* Gaussianity of \mathbf{U} is given:

Corollary 1. *Let \mathbf{X} be independent of $(\tilde{W}, \tilde{W}')'$. Then, $\mathcal{J}_{\hat{\alpha}} \xrightarrow{d} \chi^2_{D-1}$, $\mathcal{J}_{\alpha_0} \xrightarrow{d} \chi^2_D$, and $\mathcal{T} \xrightarrow{d} \mathcal{N}(0,1)$.*

Proof: *Obvious and omitted.*

Note that such independence is not plausible when there is contemporaneous correlation between the forecast errors and the innovations to the instruments. Still, if the forecast errors fulfill certain restrictions on the serial dependence of the conditional higher-order moments, the following corollary shows that normality of \mathcal{T} is recovered even in cases where Corollary 1 does not apply:

Corollary 2. *Let $E(|u_t|^{p-1} | \tilde{u}_{t-1}, \dots, \mathbf{v}_{t-1}, \dots) = \mu_{p-1}$ and $E(\tilde{u}_t^2 | \tilde{u}_{t-1}, \dots, \mathbf{v}_{t-1}, \dots) = \sigma_u^2$ be constant. Then, $\mathcal{T} \xrightarrow{d} \mathcal{N}(0,1)$.*

Proof: *See Section A of the Supplement.*

The corollary essentially requires constant conditional scale of u_t . Importantly, this condition is automatically fulfilled in the case $p = 1$, where $\tilde{u}_t(\alpha)$ follows a two-point distribution only depending on α , $\Pr(\tilde{u}_t = \alpha | \tilde{u}_{t-1}, \dots, \mathbf{v}_{t-1}, \dots) = 1 - \alpha$ and $\Pr(\tilde{u}_t = \alpha - 1 | \tilde{u}_{t-1}, \dots, \mathbf{v}_{t-1}, \dots) = \alpha$. The presence of a constant instrument is however paramount for the result; see the proof for details.

Finally, the constant alone also eliminates nonstandard distribution components from \mathcal{T} . The essential requirement is that *only* persistent instruments are employed in conjunction with a constant:

Corollary 3. *Let $\mathbf{v}_{t-1} = (1, \mathbf{v}'_{1,t-1})'$. Then, $\mathcal{T} \xrightarrow{d} \mathcal{N}(0,1)$.*

Proof: *See Section A of the Supplement.*

Corollaries 2 and 3 are particularly relevant for applied work under uncertain persistence of the instruments, since they allow for standard inference in nonstandard situations.

In what concerns the J -statistics, the expressions in Proposition 2 do not appear to simplify under the conditions of either Corollary 2 or 3.

2.3 | Recommendations for practitioners

Section B of the supplement contains a Monte Carlo evaluation of the impact of neglected persistence on the behavior of the t - and J -statistics for the practically relevant cases $p = 1$ and $p = 2$.

In a nutshell, the J -statistics are not reliable as soon as at least one instrument is highly persistent, with oversized tests and increased chances of spurious rejection of the respective null hypotheses. Unless instrument persistence can be excluded, it is not recommendable to rely on J -statistics for inference. This already suggests either to remove persistent instruments or to apply some trend-removing transformation (such as differencing) before applying the Elliott, Komunjer, and Timmermann (2005) methodology to achieve robust inference with J -statistics.

The situation is more finely nuanced for the t -statistic. For the case $p = 1$, Corollary 2 applies and the t -statistic is reliable for all instrument DGPs. For the case $p = 2$, some care needs to be taken to ensure correctly sized inference. First, practitioners could conduct a test of the null hypothesis of constant conditional moments of \tilde{u}_t to be able to justify the application of Corollary 2. Alternatively, the set of instruments may be split to separate stationary from highly persistent instruments and run separate tests for the two sets of instruments. This would allow one to exploit Corollary 3 without worrying about the conditional homoskedasticity requirement.

³In fact, types of persistence other than near-integration generate similar behavior, and distinguishing among them, e.g., deciding between fractional integration and near integration, is difficult; see Müller and Watson (2008).

In our replication of the analysis of the loss preferences of the EU Commission, we therefore take the t -ratio at face value for $p = 1$ for all instruments used. For the J -statistics for $p = 1$, and for all statistics for $p = 2$, we distinguish between persistent and stationary instruments to prevent spurious findings.

3 | REASSESSING THE EU COMMISSION FORECASTS

This section reexamines the EU Commission forecast data first analyzed by Christodoulakis and Mamatzakis (2009). They examine forecasts of five macroeconomic variables for 12 EU Member States, namely inflation, unemployment, government balance, investment, and current account, over the period 1970–2004. Each variable is predicted twice a year: in the spring for the current year, and in the autumn for the upcoming year. The number of observations for each country varies from 35 to 18 depending on its year of entry into the EU (e.g., 1986 for Spain and Portugal). For this replication study, we augment the initial dataset with more recent data, up to 2017,⁴ which yields an increased number of observations of between 31 and 48 data points. To evaluate time evolution of asymmetries, we perform the evaluation for the original time span, for the full time span, and for the last available 20 years only, i.e., 1997–2017.

For estimation under asymmetric loss, Christodoulakis and Mamatzakis (2009) primarily discussed the case $p = 1$ (with results for $p = 2$ provided as robustness checks). They worked with $D = 3$ instruments: a constant, a lagged realization and a lagged forecast error; they also provided results for $D = 2$ with constant and lagged realization, and constant and lagged forecast errors, however focusing on the \mathcal{J} -statistics. The stationarity assumptions required by Elliott, Komunjer, and Timmermann (2005) are however overly optimistic for the data at hand. Table 1 gives the sums of estimated coefficients of autoregressive (AR) processes fit to the realized values of the variables to be forecast, as well as correlation between the AR residuals and the forecast errors. (The model order p is selected individually using the Akaike information criterion.) Quite often, the cumulated AR coefficients are seen to be close to, or even above, unity.⁵ Moreover, the contemporaneous correlation between the forecast errors and the innovations to the instruments is highly positive: for the year-ahead forecasts, the correlation ranges in a neighborhood of 0.8, while it somewhat decreases to about 0.5 for the current-year forecasts.

We first replicate the study of Christodoulakis and Mamatzakis (2009). We use the dataset and the MATLAB estimation routine of Christodoulakis and Mamatzakis.⁵ The results of the replication are given in Tables S5–S9 of Section C of the supplement (where we also report values for the T -statistic testing the null $\alpha = 0.5$). Our computations lead to somewhat different figures. In particular, the three \mathcal{J}_{α_0} -statistics are different by about 5–10% for most variables and both the current-year and the year-ahead forecasts. For government balance, some of the point estimates are also different, although the differences do not change the overall picture. As a robustness check, we conducted this replication in R, obtaining the same numbers as in Tables S5–S9. The differences are likely due to different MATLAB versions. We used MATLAB 6.1, as later versions did not run the codes without modifications, but it is not clear which version was used in the original study. It should be noted at this point that certain cases display convergence problems when $D = 3$. In particular, estimation for inflation, unemployment, and government balance sometimes yields values of $\alpha \notin (0,1)$ or fails to converge altogether due to singularity of $\hat{\mathbf{S}}$.

We now focus on the time evolution of the estimates and report outcomes for the three different periods we consider. We use several sets of instruments: additionally to the original set of $D = 3$ instruments, we resort to two sets of $D = 2$ instruments: constant and lagged realization, as well as constant and lagged forecast error. A nice side effect of using $D = 2$ instruments is that singularity issues seem to be alleviated; the main reason for conducting the estimation with $D = 2$ is that we may argue (a) that the estimates resorting to a constant and the lagged realization (which tends to be persistent) lead to a robust t -statistic, while (b) the estimates building on a constant and the lagged forecast error likely fulfill stationarity assumptions and therefore deliver interpretable results for the J -statistics. Also, to ensure numerical convergence of the estimator, we use a different starting value for iterative calculation of the matrix $\hat{\mathbf{S}}$: Christodoulakis and Mamatzakis (2009), following Elliott, Komunjer, and Timmermann (2005), start their iterations by choosing the identity matrix; we calculate the starting value by plugging in $\alpha = 0.5$, which improves the convergence behavior of the estimation procedure.

⁴Obtained from http://ec.europa.eu/economy_finance/db_indicators/statistical_annex/index_en.htm.

⁵The data from the original study, as well as the codes, were kindly provided by the authors through the JAE Data Archive, <http://qed.econ.queensu.ca/jae/2009-v24.4/christodoulakis-mamatzakis/>.

TABLE 1 Instrument persistence: stylized facts, 1970–2017

		Inflation		Unemployment		Gov. balance		Investment		Current account	
		Current	Ahead	Current	Ahead	Current	Ahead	Current	Ahead	Current	Ahead
Belgium	Σ	0.781	0.793	0.778	0.747	0.284	0.795	-1.437	-0.833	0.564	0.621
	r	0.292	0.771	0.336	0.183	0.573	0.782	0.237	0.561	0.585	0.364
Denmark	Σ	1.038	1.017	0.680	0.555	0.654	0.665	-1.026	-0.605	1.065	1.182
	r	0.339	0.596	0.368	0.407	0.545	0.501	0.558	0.683	0.563	0.537
Germany	Σ	0.811	0.986	0.888	0.921	0.775	0.673	-0.092	-0.120	0.901	0.924
	r	0.434	0.880	0.380	0.203	0.365	0.682	0.589	0.919	0.570	0.801
Greece	Σ	1.037	1.009	0.761	0.926	0.848	0.621	-0.323	0.435	0.774	0.381
	r	0.358	0.437	0.212	0.538	0.423	0.597	0.350	0.519	0.596	0.547
Spain	Σ	2.968	3.411	0.835	0.778	—	0.736	0.361	-4.075	0.790	0.677
	r	0.331	0.039	0.087	0.555	0.688	0.294	-0.130	0.147	0.309	0.668
France	Σ	0.989	0.948	0.910	0.915	0.240	0.724	-0.416	0.043	0.686	0.824
	r	0.339	0.796	0.592	0.461	0.700	0.847	0.466	0.741	0.720	0.792
Ireland	Σ	1.044	0.964	0.893	0.909	0.793	0.729	-1.283	0.465	1.000	0.725
	r	0.471	0.807	0.533	0.559	0.486	0.804	0.575	0.623	0.403	0.590
Italy	Σ	0.979	0.974	0.835	0.882	0.780	0.779	—	—	-0.061	0.440
	r	0.332	0.662	0.534	0.505	-0.015	0.015	0.600	0.851	0.575	0.450
Luxembourg	Σ	0.754	0.890	0.987	0.950	0.666	0.652	—	—	0.783	0.842
	r	0.307	0.597	0.728	0.711	0.402	0.709	0.802	0.934	0.193	0.142
Netherlands	Σ	0.969	1.068	0.574	0.659	0.075	0.242	-1.195	—	0.979	0.847
	r	0.392	0.474	0.417	0.266	0.346	0.656	0.585	0.760	0.554	0.787
Portugal	Σ	0.976	1.206	0.735	0.686	0.578	0.100	-0.691	-0.999	-0.063	0.703
	r	0.367	0.335	0.151	0.335	0.072	0.594	0.119	0.150	0.415	0.380
UK	Σ	0.975	0.942	0.792	0.831	0.720	0.729	—	0.301	0.805	0.701
	r	0.701	0.406	0.353	0.496	0.605	0.563	0.676	0.926	0.080	0.619
EU	Σ	1.008	1.005	0.911	0.894	0.726	0.610	-1.101	-0.043	0.578	0.539
	r	0.577	0.808	0.464	0.397	0.715	0.863	0.287	0.874	0.511	0.662

Note: Σ denotes the sums of OLS AR(p) coefficient estimates; r denotes the contemporaneous correlation between the AR(p) residuals and the forecast errors. The order p was selected via AIC. Missing values indicate $p = 0$.

Figures 1 and 2 show point estimates $\hat{\alpha}$ together with confidence intervals based on standard normality of the t -ratio for the current-year forecasts. The results for the J -statistic $\mathcal{J}_{\hat{\alpha}}$ testing for rationality under possible loss asymmetry are only given in Figure 2, which builds on stationary instruments and thus allow for χ^2 inference following Elliott, Komunjer, and Timmermann (2005). Nicely confirming Corollary 3, the results for the case $D = 2$ with the non-stationary instrument only are almost identical to those for $D = 2$ using the stationary instrument, so we only report them in the Supplement (Section C). Although the results are mostly similar for the three sets, we do notice that there are some distinctions between Figure 1 and Figures 2 and S7. In particular for the current account, the estimator with $D = 3$ delivers estimates that may be seen as a bit too extreme. This can be explained by lack of robustness for $D = 3$. The analogous findings for the year-ahead forecasts are given in Figures 3 and 4.

To interpret the results, we follow Christodoulakis and Mamatzakis (2009). They observe optimistic tendencies in year-ahead forecasts of the EU Commission and more prudent ones for the current-year forecasts. This is reflected in the significant differences from 0.5 of the estimated α . We resort to t -statistics rather than J -statistics to assess significance, as the J -statistics are not robust to persistence. Generally, a value above 0.5 represents a higher preference of the forecaster for overprediction, since negative forecast errors are deemed more costly. Whether it can be interpreted as a prudent or an optimistic attitude depends on the variable in question. For instance, a pessimistic forecaster would tend

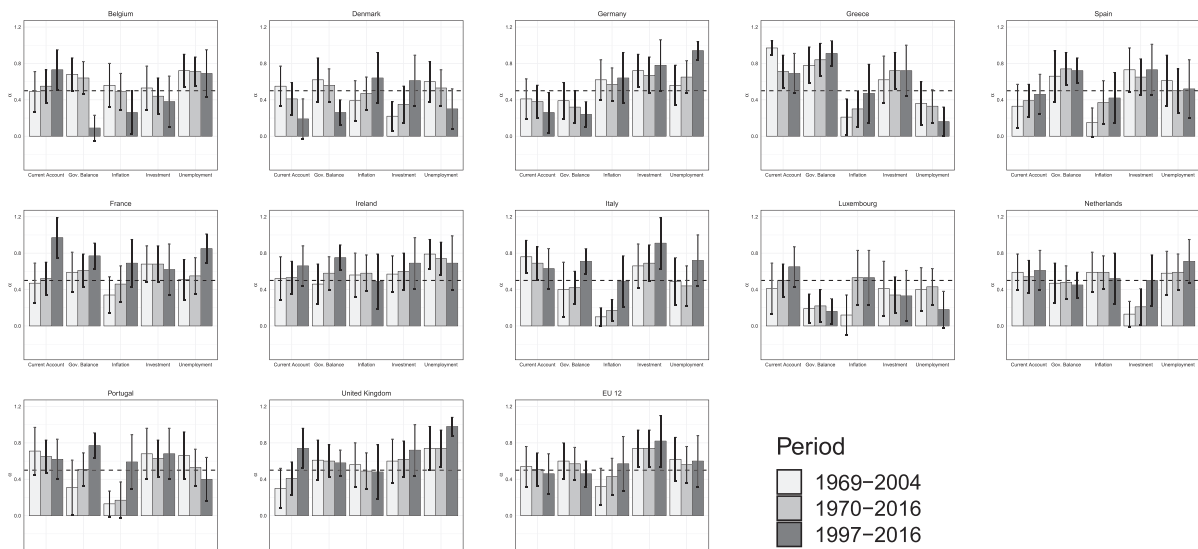


FIGURE 1 Asymmetry estimates for the EU-12 Member States over different time spans, linear asymmetric loss and $D = 3$ instruments (current-year forecasts) *Note:* The figure shows α estimates with error bars (± 2 standard errors). The estimation was conducted using three instruments: a constant, a lagged forecast error, and a lagged realization; $p = 1$

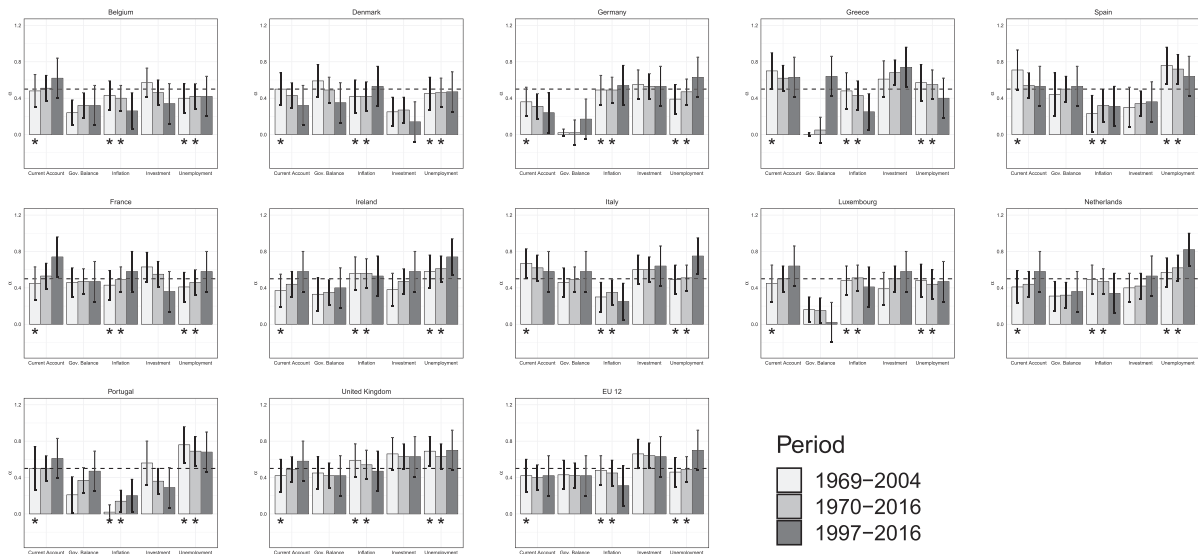


FIGURE 2 Asymmetry estimates for the EU-12 Member States over different time spans, linear asymmetric loss and $D = 2$ instruments (current-year forecasts) *Note:* The figure shows α estimates with error bars (± 2 standard errors). The estimation was conducted using two instruments: a constant and a lagged forecast error; $p = 1$. Asterisks represent significance at 5% level of $\mathcal{J}_{\hat{\alpha}}$ for rationality

to overforecast variables such as inflation and unemployment. The same holds for government balance, which yields negative values in case of a deficit. Here, underestimating the deficit is more costly in case of a prudent strategy. For investment and current account, the picture is reversed. Optimism corresponds to higher values of α , i.e., overprediction. Rationality is discussed only in the case where the set of instruments contains a constant and a lagged forecast error (see Figures 2 and 4); in other cases, we cannot rely on the $\mathcal{J}_{\hat{\alpha}}$ statistic to distinguish between asymmetry and irrationality.

Forecasts appear to have different properties in time. For both year-ahead and current-year inflation forecasts, the hypothesis of rationality can be rejected in the first two samples (original and augmented). We observe no significant deviations from symmetry for year-ahead forecasts in the last 20 years, though. For the current year, we detect optimistic preferences for Belgium, Greece, Italy, and Portugal. Similarly, the hypothesis of rationality can be rejected for

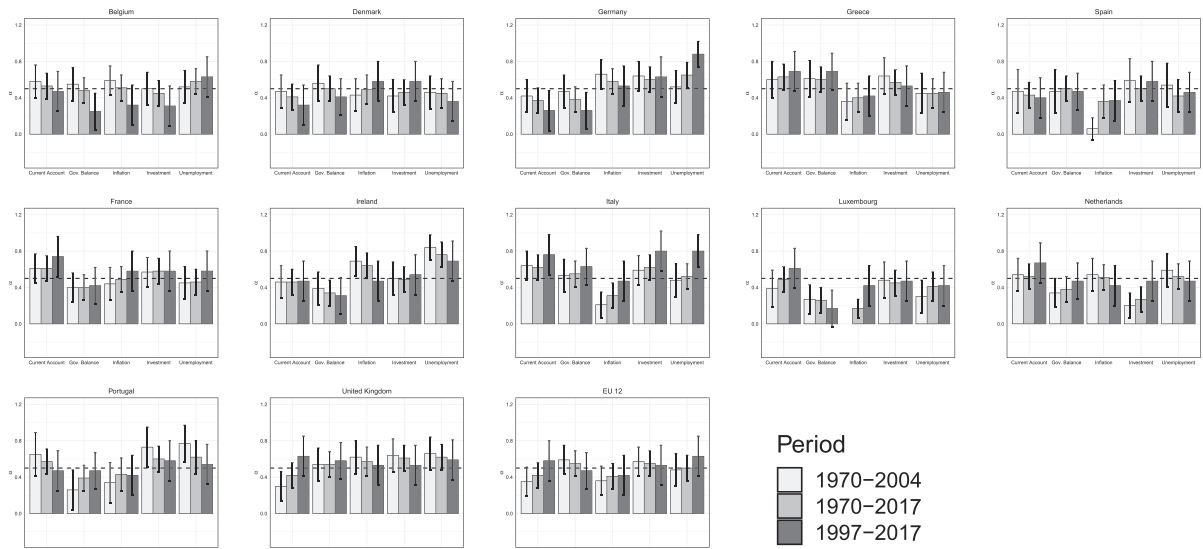


FIGURE 3 Asymmetry estimates for the EU-12 Member States over different time spans, linear asymmetric loss and $D = 3$ instruments (year-ahead forecasts) *Note:* The figure shows α estimates with error bars (± 2 standard errors). The estimation was conducted using three instruments: a constant, a lagged forecast error and a lagged realization; $p = 1$. Missing bars indicate failed convergence of the estimation algorithm

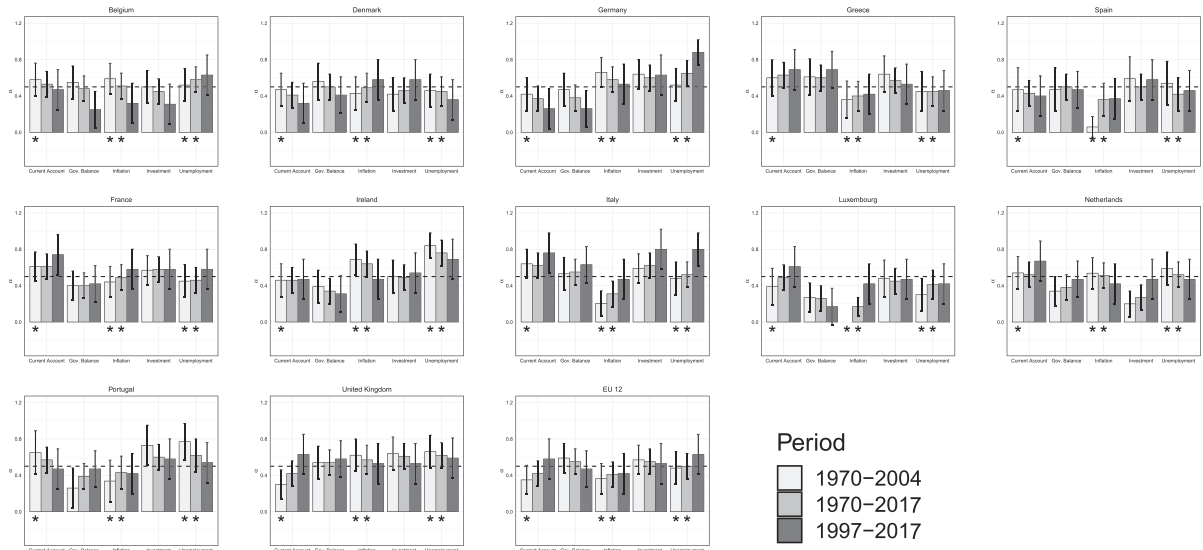


FIGURE 4 Asymmetry estimates for the EU-12 Member States over different time spans, linear asymmetric loss and $D = 2$ instruments (year-ahead forecasts) *Note:* The figure shows α estimates with error bars (± 2 standard errors). The estimation was conducted using two instruments: a constant and a lagged forecast error; $p = 1$. Asterisks represent significance at 5% level of $\mathcal{J}_{\hat{\alpha}}$ for rationality. Missing bars indicate failed convergence of the estimation algorithm

current-year as well as year-ahead unemployment forecasts in the first two samples. As for the last sample, we only find evidence of prudence for Ireland, Italy, and the Netherlands in the current year, and Germany and Italy in the year ahead (Figures 2 and 4). For government balance, we observe no deviations from rationality. For the current year, there appears to be an optimistic tendency for Germany and Luxembourg. Forecasts for Belgium, Greece, the Netherlands, and Portugal are significantly optimistic in the beginning, with a tendency towards symmetry in the last 20 years. For the year-ahead forecasts, preferences for Luxembourg remain overall optimistic, but for Belgium, Ireland, and Germany they shift to optimism in the last sample. For Portugal, however, there is a shift from optimism to symmetry over time. Investment was highlighted by Christodoulakis and Mamatzakis (2009) as being forecast overly optimistically. Here we

do not see any deviations from rationality, but results for some countries exhibit strong asymmetric preferences. In particular for the current-year forecasts, the cases of Greece, UK, Portugal, and the EU average show optimistic tendencies, while the Netherlands and Denmark deliver evidence for a more prudent forecast strategy. The year-ahead results reveal optimistic preferences for Italy and Portugal, and pessimistic ones for the Netherlands. For the current account, we see rejection of rationality in the original sample for both forecast horizons. Otherwise the evidence from the augmented sample and the last 20 years supports rational and symmetric preferences of the Commission with few minor exceptions. The notable ones are France and Italy, for which the forecasts are more optimistic, and Germany, with a tendency towards prudence. The reduced evidence against rationality in the more recent subsample could partly be due to the smaller sample size and reduced power.

Finally, the results for $p = 2$ are qualitatively very similar (see Section C of the Supplement, Figures S9 and S10), and we do not discuss them in detail. (We only provide results for $D = 2$ with stationary instruments and therefore robust J -statistics to save space.) Figure S11 in Section C of the Supplement provides rolling-window estimates of the asymmetry parameters and illustrates the overall decrease in asymmetry over time for most countries (except perhaps Germany and Belgium).

4 | CONCLUDING REMARKS

This note replicates and extends the study of Christodoulakis and Mamatzakis (2009) on the loss preferences of EU Commission forecasts for several Member States and key economic indicators.

To this end, we investigate the distribution of the estimator of the asymmetry parameter in the GMM framework of Elliott, Komunjer, and Timmermann (2005). We address here the empirically relevant case when some of the instruments are highly persistent. In theoretical derivations and Monte Carlo experiments, we find that J -statistics tend to overreject their respective null hypotheses whenever at least one instrument is highly persistent. For the t -statistic, we provide theoretical arguments that robustness to instrument persistence is given in several practically relevant cases, most prominently when $p = 1$ for all types of instruments, or, for other values of p , when one uses highly persistent or stationary instruments separately. The use of J -statistics for inference is however only recommended after persistent instruments have been eliminated from the set of instruments.

We then find that the original conclusions of Christodoulakis and Mamatzakis (2009) are largely confirmed for the extended data. However, the departures from symmetry appear to have somewhat decreased compared with the original (shorter) dataset, and rationality is rejected less often, such that EU Commission forecasts could be seen as increasingly reliable.

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SUPPORTING INFORMATION

Additional supporting information may be found online in the Supporting Information section at the end of this article.

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