Andrea Vaona

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Via G. Buffi, 13 CH-6900 Lugano

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Andrea Vaona[1]

Institute for Economic Research, University of Lugano, and Kiel Institute for the World Economy, Via Maderno 24, CH-6904 Lugano, Switzerland. E-mail: andrea.vaona@lu.unisi.ch.

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Abstract

We show that some nonparametric specification tests can be robust to disturbance autocorrelation. This robustness can be affected by the specification of the true model and by the sample size. Once applied to the prediction of changes in the Euro Repo rate by means of an index based on ECB wording, we find that the least sensitive nonparametric tests can have a comparable performance to a RESET test with robust standard errors.

1 Introduction

Ramsey's regression specification test (RESET) was showed to have serious size problems in presence of disturbance serial correlation. In particular, though Thursby (1979) originally found that the version of the RESET test due to Thursby and Schmidt (1977) - TS-RESET - could produce the expected number of Type I errors even with autocorrelated residuals, Porter and Kashyap (1984) showed that this might not hold true in presence of serial correlation of the regressors and recommended to rely on robust estimators for diagnostic equations. Leung and Yu (2001) argued that this lack of robustness is due to spurious correlation arising among highly serially correlated series even when they are independent, which leads to reject a model specification even when it is correct.

Various nonparametric specification tests were offered in the literature (Gozalo, 1993; Härdle and Mammen, 1993; Zheng, 1996; Ellison and Ellison, 2000; Horowitz and Härdle, 1994; Bierens, 1990 and Stute, 1997). Miles and Mora (2003) provided a comprehensive framework for all the test statistics above together with Monte Carlo evidence on their power. Only the statistics proposed by Ellison and Ellison (2000) and Härdle and Mammen (1993) survived their set of simulation exercises[2].

As a matter of consequence the present study focuses on these last two test statistics and on further, more recent statistics not considered in Miles and Mora (2003), such as those by Horowitz and Spokoiny (2001),

[2] Among them, the former was recommended as there is no need to resort to bootstrapping to assess its significance: the critical values of the standard normal distribution suffice. See also Dacuycuy (2005, 2006) for further Monte Carlo studies on the Zheng (1996) and the Ellison and Ellison (2000) statistics.

Hall and Yatchew (2005), Tripathi and Kitamura (2003), Guerre and Lavergne (2005) and Escanciano (2006).

Our research question is if they are robust to serial correlation, allowing to assess under a new perspective the strongest results of a considerable research effort of the recent econometric literature. Furthermore, differently to most of the literature on the sensitivity of the RESET test to disturbance autocorrelation, we do not limit our analysis to simulations, rather we apply our results to a dataset recently proposed in the central bank communication literature (Rosa and Verga, 2007).

In order to make our results comparable to those of the literature concerning the sensitivity of the RESET test to disturbance autocorrelation we do not consider nonparametric tests for dynamic models (Guay and Guerre, 2006), non-linear models (Guerre and Lavergne, 2002), models with binary discrete variables (Hsiao et al., 2007) and the varying coefficient model (Fan and Zhang, 2004). We also do not focus on tests based on series expansions (Sun and Li, 2006) or on piecewise constant functions and trigonometric polynomials (Baraud et al., 2003). Finally, endogeneity issues are beyond the scope of this paper (Horowitz, 2006).

The next section is devoted to introduce the test statistics under analysis, how our Monte Carlo experiments are structured and the results of our simulation exercises. Section 3 shows the results of our empirical application. The last section concludes.

2 Nonparametric Specification Tests, Monte Carlo Experiments and Simulation Evidence

Consider a set of $1+d$ economic variables (Y,X). Suppose we are interested in studying the conditional mean $m(x)=E(Y|X=x)$, where x is some fixed value of X and $m(x)$ is supposed to be well defined. The target of nonparametric specification tests is to check that the parametric specification of $m(\cdot)$, $m(X,\theta)$, is not rejected by the data.

The Härdle and Mammen (1993) statistic compares the Nadaraya–Watson estimator and a kernel-smoothed parametric estimator, on the basis of their weighted squared difference:

$$S_T^{(HM)} = Th^{d/2}\int \left[\hat{m}_h(x)-\hat{m}_{h,\hat{\theta}}(x)\right]^2 \pi(x)\,dx$$

where T is the number of observations, h is the smoothing bandwidth, $\hat{m}_h(x)$ is the Nadaraya–Watson estimator, $\pi(x) : \mathbf{R}^d \to \mathbf{R}$ is a weight function. The integral is taken over the support of X. $\hat{m}_{h,\hat{\theta}}(x) = \frac{\sum_{t=0}^{T} K\left(\frac{x-x_t}{h}\right) m(x_t,\hat{\theta})}{\sum_{t=0}^{T} K\left(\frac{x-x_t}{h}\right)}$ where $K(\cdot)$ is the kernel smoothing function. The asymptotic null distribution of $S_T^{(HM)}$ is normal, but for finite samples a bootstrap procedure, usually referred to as *wild bootstrap*, is recommended and it is illustrated below.

Tripathi and Kitamura (2003) proposed a test for conditional moment restrictions based on a smoothed empirical likelihood ratio (SELR) which is defined as follows:

$$SELR = 2\sum_{t=1}^{T} I\left(x_t \in \Omega_X^*\right)\sum_{s=1}^{T} w_{ts,h}\log\left(1+\frac{\Lambda\hat{e}_\theta}{T}\right)$$

where $I(\cdot)$ is an index function, Ω_X^* is a certain fixed subset of the support of X, Ω_X, $\hat{e}_\theta = y - m(x,\hat{\theta})$, Λ is a diagonal matrix with $\Lambda_{tt} = T\hat{V}^{-1}\left(x_t,\hat{\theta}\right)\sum_{s=1}^{T} w_{ts,h}\hat{e}_{\theta s}$, $\hat{V}^{-1}\left(x_t,\hat{\theta}\right) = \sum_{s=1}^{T} w_{ts,h}\hat{e}_{\theta s}^2$ and $w_{ts,h} = \frac{K\left(\frac{x_t - x_s}{h}\right)}{\sum\limits_{k=1}^{T} K\left(\frac{x_t - x_k}{h}\right)}$.[3]

One of the possible applications of the proposed test statistic, $S_T^{(TK)}$, is to test a parametric regression function against a nonparametric alternative. For $d \leq 3$, this test statistic assumes the following form:

$$S_T^{(TK)} = \frac{h^{d/2}SELR - h^{-d/2}\frac{\left[\sum_{t=1}^{T} I\left(x_t \in \Omega_X^*\right)\right]}{T}\int_{-1}^{1} K^2(x)\,dx}{2\frac{\left[\sum_{t=1}^{T} I\left(x_t \in \Omega_X^*\right)\right]}{T}\int_{-2}^{2}\left[\int_{-1}^{1} K(v)\,K(x-v)\,dv\right]^2 dx}$$

One further test statistic was proposed by Hall and Yatchew (2005), $S_T^{(HY)}$, which has the advantage of not depending on the choice of a bandwidth. To obtain $S_T^{(HY)}$ one first partitions the dataset in three subsamples, D_1, D_2 and D_3. Then she defines

$$\psi_1 = \frac{\sum_{t=1}^{T} \hat{e}_{\theta t} I\left(\{y_t, x_t\} \in D_1\right)}{T}$$
$$\psi_2 = \frac{\sum_{t=1}^{T} \hat{e}_{\theta t} I\left(\{y_t, x_t\} \in D_2\right)}{T}$$
$$\psi_3 = \frac{\sum_{t=1}^{T} \hat{e}_{\theta t} I\left(\{y_t, x_t\} \in D_3\right)}{T}$$

Finally it is possible to obtain:

[3] Λ_{tt} is a first step approximation of the Lagrangian multiplier associated with the constraint that the smoothed sum of the parametric residuals is equal to zero in the maximization of the expected loglikelihood function.

$$S_T^{(HY)} = \sqrt{T\frac{(\psi_1)^2 + (\psi_2)^2 + (\psi_3)^2}{3}}$$

The limiting distribution of $S_T^{(HY)}$ is "a complicated function of a Gaussian process the properties of which depend on unknowns" (Hall and Yatchew, 2005). As a matter of consequence, bootstrapping is recommended and, similarly to $S_T^{(HM)}$, $S_T^{(TK)}$, Hall and Yatchew (2005) propose to use *wild bootstrapping.*

Similarly to Hall and Yatchew (2005), the statistic proposed by Escanciano (2006), $S_T^{(E)}$, does not rely on a smoothing parameter and it is as follows

$$S_T^{(E)} = \sum_{s=1}^{T}\sum_{t=1}^{T}\sum_{r=1}^{T} \hat{e}_{\theta s}\hat{e}_{\theta t}A_{str}$$

with

$$A_{str} = \frac{\pi^{\frac{d}{2}-1}}{\Gamma\left(\frac{d}{2}+1\right)}\left|\pi - ar\cos\left(\frac{(x_s - x_r)'(x_t - x_r)}{\|(x_s - x_r)\|\,\|(x_t - x_r)\|}\right)\right|$$

where $\Gamma(\cdot)$ is the gamma function.

Also the critical values of $S_T^{(E)}$ are approximated by means of *wild bootstrapping*, which works as follows. One first generates independent $\{e_t^*\}_{t=0}^T$ from a discrete distribution with:

$$\Pr\left\{e_t^* = \frac{1+\sqrt{5}}{2}e_t\right\} = \frac{5-\sqrt{5}}{10} \text{ and } \Pr\left\{e_t^* = \frac{1-\sqrt{5}}{2}e_t\right\} = \frac{5+\sqrt{5}}{10}$$

where $e_t = y_t - \hat{m}_h(x_t)$ and y_t is a realization of Y. For $S_T^{(HY)}$, e_t is centred before proceeding further. Afterwards, one computes the bootstrap data $\{x_t^*; y_t^*\}$ where $x_t^* = x_t$ and $y_t^* = m(x_t, \hat{\theta}) + e_t^*$ and the corresponding bootstrap statistic $S_T^{*(HM)}$. The process is repeated B times to obtain $\left\{S_{t,j}^*\right\}_{j=1}^B$. H_0 is rejected if $S_t > S_{T,(1-\alpha)}^*$where $S_{T,(1-\alpha)}^*$ is the $(1-\alpha)$ quantile of $\left\{S_{t,j}^*\right\}_{j=1}^B$ and α is the significance level.

The test proposed by Horowitz and Spokoiny (2001) is based on a centred, Studentized version of the sum of the squared differences between the nonparametric and the smoothed parametric estimator:

$$S_{h,T} = \sum_{t=1}^{T} \left[\hat{m}_h(x_t) - \hat{m}_{h,\hat{\theta}}(x_t)\right]^2$$

The statistic assumes the following form:

$$S_T^{(HS)} = \frac{S_{h,T} - \hat{\Sigma}_{h,T}}{\hat{V}(S_{h,T})}$$

where $\hat{\Sigma}_{h,T} = \sum_{t=1}^{T} a_{tt,h}\sigma_T^2(X_t)$ and $\hat{V}^2(S_{h,T}) = 2\sum_{t=1}^{T}\sum_{s=1}^{T} a_{ts,h}\sigma_T^2(X_t)\sigma_T^2(X_s)$, with $a_{ts,h}$ is the $ts-th$ element of $A = W_h'W_h$ where the $ts-th$ element of W_h is $w_{ts,h}$.

$\sigma_T^2(X_t)$ assumes different forms depending on the assumptions concerning the error of the model under analysis and the number of regressors. We assume that the researcher starts with the assumption of homoskedasticity. If $d = 1$, let $X_{(1)} < X_{(2)} < ... < X_{(s)} < ... < X_{(T)}$ be an ordered sequence of design points, then

$$\hat{\sigma}_T^2(X_t) = \hat{\sigma}_T^2 = \frac{1}{2(T-1)}\sum_{s=1}^{T-1}\left(Y_{(s+1)} - Y_{(s)}\right)^2$$

If $1 < d \leq 4$, instead

$$\hat{\sigma}_T^2(X_t) = \hat{\sigma}_T^2 = \frac{1}{2T}\sum_{s=1}^{T-1}\left(Y_s - Y_{t(s)}\right)^2$$

where $t(s)$ is a set of indices that is defined recursively as follows:

$$t(1) = \arg\min_{t=2,...,T} \|X_t - X_1\|$$

and

$$t(s) = \arg\min_{t\neq s, t(1),...,t(s-1)} \|X_t - X_s\| \text{ for } s = 2, ..., T$$

with $\|\cdot\| = \sum_t^T (\cdot_t)^2$. In words, $t(s)$ is "the index of the design point that is nearest to X_t among those whose indices are not $t(1), ..., t(s-1)$" (Horowitz and Spokoiny, 2001, p. 608). Finally, $S_T^{(HS)}$ is computed over a finite set of bandwidths and only its maximum value is considered.

Also the test proposed by Guerre and Lavergne (2006) makes direct use of $\hat{e}_\theta$ as follows. First

$$S_{Th}^{(GL)} = \sum_{1\leq t,s\leq T} \hat{e}_{\theta t}\sqrt{a_{ts,h}}\hat{e}_{\theta s}$$

is computed over a finite set of bandwidths, H_T, which gets larger the larger is the sample size and whose order, J_T, is equal to the first integer smaller than $\log T$. The criterion to select the value of $S_{Th}^{(GL)}$ among those computed differs with respect of Horowitz and Spokoiny (2001). Indeed, one does not pick the bandwidth corresponding to the maximum value over those calculated but:

$$\tilde{h} = \arg\max_{h \in H_T} \left\{ S_{T,h}^{(GL)} - \gamma_T \hat{v}_{h,h_0} \right\}$$

where $\gamma_T = c\sqrt{2\log J_T}$, $c = 1.5$, $\hat{v}_{h,h_0} = 2\sum_{t=1}^{T}\sum_{s=1}^{T}\left(a_{ts,h} - a_{ts,h_0}\right)\hat{\sigma}_T^2\left(X_t\right)\hat{\sigma}_T^2\left(X_s\right)$ and h_0 is the bandwidth corresponding to the $S_{T,h}^{(GL)}$ having the smallest variance. The null of no misspecification is rejected when $\frac{S_{T\tilde{h}}^{(GL)}}{\hat{v}_{h_0}} \geq z_\alpha$, where $\hat{v}_{h_0} = 2\sum_{t=1}^{T}\sum_{s=1}^{T}\left(a_{ts,h_0}\right)\hat{\sigma}_T^2\left(X_t\right)\hat{\sigma}_T^2\left(X_s\right)$ and z_α is a bootstrapped critical value. $\hat{\sigma}_T^2\left(X_t\right)$ is computed as in Horowitz and Spokoiny (2001).

With difference to the other tests illustrated above, inference regarding $S_T^{(HS)}$ and $S_{T\tilde{h}}^{(GL)}$ relies on bootstrapping, but not on its "wild" version, as $\{e_t^*\}_{t=0}^{T}$ are not generated from a discrete distribution, but they are sampled randomly from a normal distribution with variance $\hat{\sigma}_T^2\left(X_t\right)$.

The idea behind the statistic proposed by Ellison and Ellison (2000) is that either the omission of a relevant variable or the choice of an incorrect functional form for $m(\cdot)$ produces a spatial structure in the residuals, exploiting which it is possible to detect model misspecification. The Ellison and Ellison (2000) test has the following form:

$$S_T^{(EE)} = \frac{\hat{e}_\theta' W_T \hat{e}_\theta}{\sqrt{2}s\left(EW_TE\right)} + FSC_T$$

where E is a $T \times T$ diagonal matrix with $tt-th$ element $\hat{e}_{\theta t}$, and $s(\cdot) = \left(\sum_{t,s} \cdot_{ts}^2\right)^{\frac{1}{2}}$. W_T is a weight matrix, whose elements are as follows:

$$w_{tsT} = \begin{cases} \dfrac{K(\frac{x_t - x_s}{h})}{\sum\limits_{k \neq t} K(\frac{x_t - x_k}{h})} \text{ if } t \neq s \text{ and } \sum\limits_{k \neq t} K(\frac{x_t - x_k}{h}) > 0 \\ 0 \quad \text{otherwise} \end{cases}$$

FSC_T is a finite sample correction with the following form:

$$FSC_T = \frac{\sum\limits_{\ell=0}^{d} \hat{\gamma}_\ell}{\sqrt{2}s\,(W_T)}$$

"where $\hat{\gamma}_\ell$ is the coefficient on $X_{\cdot\ell}$, the $\ell - th$ explanatory variable in the null model, in a regression of $W_T X_{\cdot\ell}$ on X and a constant and $\hat{\gamma}_0$ is the constant term from a regression of $W_T \mathbf{1}_T$ on X, where $\mathbf{1}_T$ is a $T \times 1$ vector of ones" (Ellison and Ellison, 2000). $S_T^{(EE)} \to N(0, 1)$.

All the tests considered are consistent against any alternative hypothesis.

In order to make our results comparable to those of the previous literature on the issue, we stick to the designs for Monte Carlo experiments adopted by Porter and Kashyap (1984) and Leung and Yu (2001), illustrated in Table 1. We consider sample sizes of 50 and 200. All the tests are conducted at the 5 percent level of significance. Parameters θ_0 and θ_1 are always estimated by least squares.

Again for sake of comparability we specify the nonparametric tests as in Miles and Mora (2003). In all univariate kernel estimations we use the quartic kernel:

$$K(\cdot) = \frac{15}{16}\left(1 - \cdot^2\right)^2 I\,(|\cdot| \leq 1)$$

h was set equal to $\lambda S_X n^{-\frac{1}{5}}$ where S_X is the sample standard deviation of the regressor and $\lambda = 3.5$[4]. In the multivariate estimations we use the product of quartic kernels and each regressor is previously divided by its sample standard deviation. h was set equal to $\lambda n^{-\frac{1}{(d+4)}}$ with two exceptions. For $S_T^{(HS)}$, the bandwidth was chosen in the set $\left\{2.5n^{-\frac{1}{(d+4)}}, 3n^{-\frac{1}{(d+4)}}, 3.5n^{-\frac{1}{(d+4)}}, 4n^{-\frac{1}{(d+4)}}, 4.5n^{-\frac{1}{(d+4)}}\right\}$ to maximize the value of the statistic as recommended by Horowitz and Spokoiny (2001). For $S_{Th}^{(GL)}$ we set the bandwidth equal to $\lambda n^{-\frac{1}{(d+4)}}$ with $\lambda = 3.5 + (\frac{(J_T+1)}{4}) - (i * 0.5)$ with $i = 1, ..., J_T + 1$.

The statistic $S_T^{(HM)}$ is computed with $\pi(x^s) = I(x^s \in [-1,96;1,96])$ for design 1 and $\pi(x^s) = I(x^s \in [-1,8;1,8] \times \ [-1,8;1,8])$ for design 2, where x^s is the matrix of the standardized regressors including the time trend. In a similar fashion Ω_X^* is chosen to build $S_T^{(TK)}$. In all cases, the integral in $S_T^{(HM)}$ is approximated numerically. When a bootstrap procedure is required we perform $B = 300$ bootstrap replications.

For $S_T^{(HY)}$ we partition the dataset according to y_t being smaller than the 33rd percentile of y, included between its 33rd and 66th percentile or greater than its 66th percentile.

Our results are based on 2000 replications of the data-generating process and they are set out in Tables 2 and 3.

Similarly to RESET and TS-RESET, nonparametric specification tests are sensitive to disturbance autocorrelation. However, different tests display different degrees of sensitiveness, depending on the design of the experiments and on sample sizes.

[4]Miles and Mora (2003) showed that the Ellison and Ellison (2000) and the Härdle and Mammen (1993) tests are robust to changes in λ in a neighborhood of 3.5.

The least robust tests are those proposed by Tripathi and Kitamura (2003), Horowitz and Spokoiny (2001), Guerre and Lavergne (2006) and Ellison and Ellison (2000). Regarding design 1, the Härdle and Mammen (1993), the Hall and Yatchew (2005) and the Escanciano (2006) tests are not robust to serial correlation in large samples, but they display more robustness in small samples. The Escanciano (2006) test seems to be the most performing one, however, it would also appear possible to rely on Härdle and Mammen (1993) even with a very high disturbance autocorrelation if the independent variable has a small or moderate serial correlation. On the other hand, with high serial correlation in both the independent variable and the disturbance, Hall and Yatchew (2005) could be used.

For design 2, a similar pattern emerges. Moreover, the tests by Härdle and Mammen (1993) and by Hall and Yatchew (2005) display some robustness to serial correlation in large samples too. The performance of the Escanciano (2006) test is not harmed by serial correlation in the disturbance and in the independent variable both in small and in large samples.

3 An application to central bank communication

In the literature on the sensitiveness of RESET tests to disturbance autocorrelation, it is rather difficult to find empirical applications because Monte Carlo experiments are usually conducted for naïve designs as those in Table 1. However, a recent literature, reviewed in Rosa and Verga (2007), has focused on assessing the consistency of the communication of the European Central Bank (ECB). In this context, Rosa and

Verga (2007) proposed the following model:

$$R_{t+m} - R_t = \beta_0 + \beta_1 \left(r_{t,t+m,1} - R_t\right) + \beta_2 Index_t + \epsilon_t \tag{1}$$

where R_{t+m} is the monthly Repo rate in force within m months from t, β_0,β_1,β_2 are parameters, ϵ_t is a stochastic error, $r_{t,t+m,1}$ is the (implicit) Euribor rate quoted on day t for an interbank loan for 1 month starting at day $t+m$ months and $Index_t$ is the Rosa and Verga (2007) wording indicator.

This indicator translates into an ordered scale the qualitative information contained in the introductory statement of the ECB President in his monthly press conference held on Governing Council meeting days. Compared to the other indicators in the literature, $Index_t$ has the advantage to be elaborated on the basis of the rules of "hermeneutic theory"[5]. The purpose of estimating (1) is to check if central bankers' words provide complementary information with respect to those already in possess of financial markets and mirrored by Euribor rates.

Two hypotheses underlie (1) : the monetary authority should not be severely time-inconsistent and the public (including researchers) should understand the language of the monetary authority. The rejection of (1) by a model misspecification test would imply that something is missing, in terms of either functional form or omitted variables.

The dataset spans from January 1999 to December 2004. The sources of the data are indicated in the appendix of Rosa and Verga (2007). Estimates are performed for $m = 1, ..., 6$[6]. (1) is a forecasting regression,

[5] See Rosa and Verga (2007) note 10 and pp. 149 and 150.

[6] In the results presented by Rosa and Verga (2007), $Index_t$ performs rather well,

therefore overlapping data imply that the larger is m and the greater is disturbance serial correlation (Harri and Brorsen, 2002). Our results are contained in Table 4. The first two lines of the table show an estimate of the autocorrelation present in the residuals and in $(r_{t,t+m,1} - R_t)$. Also $Index_t$ turns out to display a considerable level of autocorrelation once regressed on its first lag:

$$Index_t = \underset{(0.44)}{0.07} + \underset{(0.00)}{0.83}\ Index_{t-1} + v_t$$

where p-values are reported in parentheses and v_t is a stochastic error.

We use as benchmark the TS-RESET test robust to autocorrelation, which supports the model for all the values of m. In this application a TS-RESET test not robust to autocorrelation performs better than the RESET test: though its p-values are always smaller than the TS-RESET test robust to autocorrelation, it never rejects the data. Regarding nonparametric tests for model misspecification, the potential risks arising from disturbance autocorrelation clearly appears in Table 4. The Ellison and Ellison (2000) test returns inconsistent results with those of our benchmark statistic. The other tests perform better, but, not surprisingly, only those proposed by Härdle and Mammen (1993), Hall and Yatchew (2005) and Escanciano (2006) never reject the model.

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but we did not manage to replicate them. Our results are contained in Table A in the Appendix. $Index_t$ would appear to have poor information content at very short and very long time horizons. In other terms, central bank communications would have a lagged impact on Repo rates, vanishing after 4 months. This is not strictly relevant for the main focus of this paper, as we would like to assess the potential pitfalls arising when using nonparametric misspecification tests in presence of serial correlation, not the explanatory power of the index proposed by Rosa and Verga (2007).

4 Conclusions

In this paper we have showed that nonparametric model specification tests can be sensitive to disturbance autocorrelation. However, the Escanciano (2006) test performs rather well and it is possible to combine the Härdle and Mammen (1993) and the Hall and Yatchew (2005) tests to have reliable results. This strategy holds true in large samples only for models with a time trend. The results of the proposed empirical application are in line with those of the Monte Carlo simulations.

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Table 1 – Simulation designs

Design	True Specification	Null Specification	Data Generating Process
[1]	$y_t = 0 + 0.5x_t + u_t$	$y_t = \theta_0 + \theta_1 x_t + u_t$	$x_t = \rho_1 x_{t-1} + v_t$
			$\rho_1 = 0,0.25,0.5,0.75,0.9,0.95$
			$u_t = \rho_2 u_{t-1} + \varepsilon_t$
			$\rho_2 = 0,0.25,0.5,0.75,0.9,0.95$
[2]	$y_t = t + 0.5x_t + u_t$	$y_t = \theta_0 t + \theta_1 x_t + u_t$	$x_0 \sim N(5/(1-\rho_1),10/(1-\rho_1^{\,2}))$
			$u_0 \sim N(0,1/(1-\rho_2^{\,2}))$
			$v_t \sim i.i.d.N(0,50)$
			$\varepsilon_t \sim i.i.d.N(0,1)$
			$t = 1,2,...,T$

Table 2 – Rejection frequencies of nonparametric specification tests for design [1]

	ρ_1	T=50						T=200					
		ρ_2											
		0.00	**0.25**	**0.50**	**0.70**	**0.90**	**0.95**	**0.00**	**0.25**	**0.50**	**0.70**	**0.90**	**0.95**
Härdle and Mammen (1993)	**0.00**	0.35	0.40	0.85	1.55	2.85	4.50	0.60	1.50	1.75	2.90	4.10	4.65
	0.25	0.35	0.20	0.45	1.90	3.40	4.80	1.10	1.30	2.10	2.90	5.20	4.60
	0.50	0.15	0.05	0.60	1.60	4.95	5.15	0.60	1.00	1.95	4.10	7.30	8.70
	0.70	0.00	0.20	0.35	2.80	8.15	10.95	0.10	0.75	1.50	7.00	16.70	21.10
	0.90	0.00	0.00	0.05	2.00	8.50	13.65	0.00	0.00	0.55	10.10	30.55	44.80
	0.95	0.00	0.00	0.00	0.75	5.20	10.05	0.00	0.00	0.15	5.30	32.75	53.25
Tripathi and Kitamura (2003)	**0.00**	1.50	0.80	1.40	2.35	3.65	3.40	2.60	2.80	3.20	4.70	5.10	4.30
	0.25	1.55	1.55	1.85	1.95	3.00	4.55	2.25	3.25	2.95	4.15	4.50	4.60
	0.50	0.95	1.10	2.05	3.60	4.95	5.10	2.85	2.95	3.10	5.20	6.35	8.05
	0.70	0.60	1.30	2.90	5.65	10.65	13.05	1.95	3.45	6.35	10.50	15.75	19.40
	0.90	0.40	1.75	4.85	13.70	25.85	29.10	1.00	2.50	8.00	23.60	42.75	46.65
	0.95	0.50	1.80	6.95	19.70	37.35	43.90	0.50	2.25	11.85	35.10	62.05	70.80
Hall and Yatchew (2005)	**0.00**	1.25	1.95	2.80	5.30	11.00	17.15	6.10	7.55	11.40	23.05	49.65	67.80
	0.25	1.75	2.45	2.55	4.95	11.00	17.05	5.85	6.05	10.10	23.20	48.85	64.45
	0.50	1.05	1.65	2.10	4.70	9.75	14.20	4.30	3.75	6.95	16.40	39.80	58.55
	0.70	0.75	0.90	1.10	3.95	7.60	12.00	1.75	1.80	4.65	10.90	27.50	44.40
	0.90	0.05	0.15	0.35	1.70	5.95	9.10	0.40	0.75	2.20	7.00	21.30	30.95
	0.95	0.00	0.00	0.15	1.30	5.00	6.70	0.00	0.00	0.20	2.60	16.65	25.90
Horowitz and Spokoiny (2001)	**0.00**	2.50	2.45	2.20	2.05	1.90	1.95	1.85	2.00	2.10	1.90	1.75	1.95
	0.25	2.85	2.90	2.50	2.40	2.50	2.60	2.15	2.05	1.90	1.85	1.85	2.35
	0.50	3.10	2.25	3.40	4.20	3.25	3.85	1.95	2.60	2.60	3.15	3.40	3.35
	0.70	2.20	4.05	7.40	10.60	12.35	13.65	1.50	3.60	6.20	10.50	15.20	16.80
	0.90	0.80	4.10	10.10	23.55	34.50	37.05	2.30	6.60	17.05	31.55	44.45	47.25
	0.95	0.65	2.20	9.05	28.35	47.25	55.10	1.30	7.00	26.10	51.10	69.65	74.90
Guerre and Lavergne (2006)	**0.00**	5.40	4.05	5.00	5.85	5.20	4.45	4.75	5.50	5.55	4.35	5.15	4.60
	0.25	4.90	4.65	4.50	6.00	5.20	6.40	4.35	4.85	6.50	9.20	13.80	16.35
	0.50	4.70	6.15	6.70	8.40	10.10	8.80	5.65	6.00	7.75	14.75	28.60	37.05
	0.70	5.80	7.50	11.80	17.90	21.00	21.80	4.90	6.20	9.80	22.35	45.85	62.25
	0.90	5.15	10.85	19.85	35.30	43.45	48.90	5.30	5.55	10.65	27.75	57.20	76.55
	0.95	5.70	12.35	24.00	45.05	58.35	62.70	4.65	6.25	11.05	27.55	64.35	81.45
Escanciano (2006)	**0.00**	0.25	0.35	0.60	1.40	2.95	3.30	1.05	1.7	1.95	2.85	3.25	4.60
	0.25	0.20	0.15	0.50	0.95	2.15	3.20	0.75	1.00	1.65	2.70	5.15	4.95
	0.50	0.15	0.10	0.55	1.35	3.50	4.80	0.75	0.70	1.45	4.90	7.95	8.20
	0.70	0.00	0.05	0.30	1.60	6.35	9.35	0.10	0.45	1.85	7.80	15.90	23.20
	0.90	0.00	0.00	0.00	0.95	7.10	10.65	0.00	0.05	0.40	9.40	30.65	41.60
	0.95	0.00	0.00	0.00	0.35	4.10	8.95	0.00	0.00	0.00	3.40	30.80	49.65
Ellison and Ellison (2000)	**0.00**	0.60	0.35	0.45	0.55	0.65	0.45	1.15	1.55	1.25	1.60	1.05	0.80
	0.25	0.35	0.40	0.80	0.50	0.40	0.70	0.90	1.15	1.95	1.50	1.60	1.45
	0.50	0.30	0.60	1.05	1.40	1.40	0.90	1.85	1.95	1.60	3.05	4.05	3.80
	0.70	0.65	1.15	2.50	4.55	6.35	6.80	0.45	3.25	5.95	11.30	14.35	18.05
	0.90	0.60	2.85	8.35	16.30	23.35	23.95	1.70	6.25	18.30	36.60	46.30	48.95
	0.95	1.05	3.20	12.30	27.30	36.45	42.10	2.15	10.70	32.85	60.40	73.65	78.30

ρ_1: is the exogenous variable autocorrelation parameter; ρ_2: is the error autocorrelation parameter.

Table 3 – Rejection frequencies of nonparametric specification tests for design [2]

	ρ_1	**T=50**						**T=200**					
		ρ_2											
		0.00	**0.25**	**0.50**	**0.70**	**0.90**	**0.95**	**0.00**	**0.25**	**0.50**	**0.70**	**0.90**	**0.95**
Härdle and Mammen (1993)	**0.00**	0.00	0.00	0.00	0.00	1.10	3.25	0.00	0.00	0.00	0.00	0.00	0.65
	0.25	0.00	0.00	0.00	0.05	1.55	4.45	0.00	0.00	0.00	0.00	0.00	0.55
	0.50	0.00	0.00	0.00	0.05	1.80	4.40	0.00	0.00	0.00	0.00	0.00	0.95
	0.70	0.00	0.00	0.00	0.25	4.35	7.90	0.00	0.00	0.00	0.00	0.00	1.50
	0.90	0.00	0.00	0.05	0.60	6.20	10.90	0.00	0.00	0.00	0.00	0.05	4.45
	0.95	0.00	0.00	0.00	1.40	8.75	13.70	0.00	0.00	0.00	0.00	0.45	7.50
Tripathi and Kitamura (2003)	**0.00**	1.25	1.85	3.55	7.00	12.85	16.15	1.30	1.00	2.85	5.35	11.75	21.10
	0.25	2.00	2.55	3.20	7.55	14.30	17.85	0.95	1.40	1.55	4.90	12.05	18.35
	0.50	3.10	3.70	5.85	9.05	16.00	21.60	1.45	2.10	3.20	5.35	11.65	19.30
	0.70	5.40	8.30	12.50	16.30	22.70	27.45	5.05	6.65	9.85	11.80	19.50	26.20
	0.90	4.00	6.25	8.95	15.70	23.55	26.45	8.95	12.10	14.65	23.00	30.10	36.95
	0.95	2.80	6.05	8.30	15.65	24.75	29.55	2.50	4.10	7.40	12.80	20.25	26.75
Hall and Yatchew (2005)	**0.00**	0.00	0.00	0.00	2.20	13.05	18.70	0.00	0.00	0.00	0.00	5.50	25.00
	0.25	0.00	0.00	0.05	2.20	13.25	17.05	0.00	0.00	0.00	0.00	6.80	25.05
	0.50	0.00	0.00	0.00	1.15	9.85	16.30	0.00	0.00	0.00	0.00	4.80	21.90
	0.70	0.00	0.00	0.00	0.65	5.15	9.10	0.00	0.00	0.00	0.00	3.15	17.00
	0.90	0.00	0.00	0.00	0.20	2.00	2.60	0.00	0.00	0.00	0.00	1.95	11.00
	0.95	0.00	0.00	0.00	0.00	0.55	1.05	0.00	0.00	0.00	0.00	0.60	5.15
Horowitz and Spokoiny (2001)	**0.00**	0.00	0.00	0.05	2.45	20.75	31.05	0.00	0.00	0.00	4.40	64.90	88.15
	0.25	0.00	0.00	0.05	3.15	23.55	33.45	0.00	0.00	0.00	5.10	66.75	90.70
	0.50	0.00	0.00	0.05	4.75	25.45	34.40	0.00	0.00	0.00	7.65	71.95	93.95
	0.70	0.00	0.00	0.20	8.05	31.60	40.80	0.00	0.00	0.15	15.55	83.50	95.65
	0.90	0.00	0.00	0.90	12.80	37.75	46.05	0.00	0.00	0.75	34.25	91.45	98.05
	0.95	0.00	0.00	1.45	17.40	43.15	52.70	0.00	0.00	0.90	40.95	92.60	98.85
Guerre and Lavergne (2006)	**0.00**	22.35	25.35	31.60	41.60	57.50	63.20	4.45	7.80	16.15	37.30	67.55	80.85
	0.25	20.70	26.10	30.40	44.65	62.05	67.50	5.05	7.95	19.75	42.40	70.80	83.05
	0.50	19.55	26.70	36.75	52.60	67.75	73.90	4.05	11.55	25.30	52.45	78.00	87.80
	0.70	20.85	30.80	46.00	65.20	75.85	81.75	4.85	16.20	38.50	70.10	89.65	94.25
	0.90	28.75	38.05	51.75	69.35	82.40	87.10	6.25	20.45	50.55	81.95	96.35	97.85
	0.95	31.95	41.55	54.20	70.40	81.80	85.95	10.40	26.30	56.10	83.65	95.25	98.70
Escanciano (2006)	**0.00**	0.00	0.00	0.00	0.00	2.90	7.40	0.00	0.00	0.00	0.00	0.05	6.70
	0.25	0.00	0.00	0.00	0.00	3.10	8.00	0.00	0.00	0.00	0.00	0.00	6.20
	0.50	0.00	0.00	0.00	0.00	2.90	6.15	0.00	0.00	0.00	0.00	0.10	5.70
	0.70	0.00	0.00	0.00	0.00	2.25	6.45	0.00	0.00	0.00	0.00	0.15	5.20
	0.90	0.00	0.00	0.00	0.00	1.50	3.60	0.00	0.00	0.00	0.00	0.10	5.25
	0.95	0.00	0.00	0.00	0.05	2.40	5.90	0.00	0.00	0.00	0.00	0.05	4.55
Ellison and Ellison (2000)	**0.00**	1.35	7.35	18.05	46.00	72.45	79.25	2.40	8.55	23.35	57.90	87.45	94.70
	0.25	2.15	7.45	19.70	48.85	75.15	82.75	2.50	8.40	26.90	59.35	88.75	95.20
	0.50	1.40	6.90	23.45	52.05	77.30	83.60	2.75	9.60	31.30	68.10	92.15	96.90
	0.70	1.90	8.75	26.60	59.85	81.10	87.00	1.95	15.20	43.10	79.70	95.85	98.75
	0.90	1.30	7.75	24.15	58.60	81.70	87.40	2.20	16.95	54.30	90.65	98.35	99.65
	0.95	1.25	5.90	22.85	55.80	76.35	82.65	2.40	17.25	53.85	91.75	98.90	99.75

ρ_1: is the exogenous variable autocorrelation parameter; ρ_2: is the error autocorrelation parameter.

Table 4 – Misspecification tests for a model predicting Repo changes using ECB words (Model 1)

	m=1	m=2	m=3	m=4	m=5	m=6
Residual Autocorrelation§§	-0.15	0.48*	0.54*	0.59*	0.66*	0.63*
($r_{t,t+m,1}$-R_t) autocorrelation§§	0.52*	0.68*	0.77*	0.81*	0.84*	0.83*
TS-RESET robust to autocorrelation°	0.10	0.23	0.72	0.78	0.94	0.63
RESET°	0.13	<u>0.01</u>	<u>0.00</u>	<u>0.00</u>	<u>0.00</u>	<u>0.00</u>
TS-RESET°	0.07	0.22	0.41	0.58	0.40	0.22
Härdle and Mammen (1993)§	0.68	0.19	0.22	0.38	0.32	0.38
Tripathi and Kitamura (2003)§	0.33	<u>0.03</u>	0.53	0.26	0.33	0.12
Hall and Yatchew (2005)§	1.00	1.00	0.47	0.96	0.66	0.42
Horowitz and Spokoiny (2001)§	0.90	0.57	0.07	0.18	<u>0.04</u>	0.51
Guerre and Lavergne (2006)§	0.90	0.57	0.07	0.18	<u>0.04</u>	0.51
Escanciano (2006)§	0.80	0.15	0.13	0.21	0.18	0.12
Ellison and Ellison (2000)°	<u>0.01</u>	<u>0.00</u>	<u>0.00</u>	<u>0.00</u>	<u>0.00</u>	<u>0.00</u>
Observations	62	62	62	62	62	62

m: time horizon of Repo changes. §§: coefficient of the regression of the concerned variable on its first lag. The model includes a constant. °: p-values. §: bootstrapped p-values. *: significant at a 5 % level. The null of all the tests is that the model is well specified.

Appendix

Table A - Prediction of the Repo change using both ECB words and financial market's information set

	m=1	m=2	m=3	m=4	m=5	m=6
Constant	-0.07	-0.13	-0.18	-0.24	-0.29	-0.34
	(0.00)	(0.00)	(0.00)	(0.00)	(0.13)	(0.36)
$(r_{t,t+m,1}-R_t)$	0.38	0.55	0.74	1.09	1.31	1.49
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
$Index_t$	0.03	0.09	0.12	0.10	0.08	0.06
	(0.13)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)

Note: monthly observations on days of ECB Governing Council meetings, January 1999–December 2004. The econometric method is Ordinary Least Squares. Newey-West standard errors in brackets. In this overlapping data case, the forecast error is not realized until m-months in the future, so it will follow a MA(*m-1*) time series process. Therefore we set the maximum lag length of the disturbance process to *m-1*.

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