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Nonparametric Test for the Form of Parametric Regression with Time Series Errors

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Abstract

We propose a new nonparametric method for testing the parametric form of a regression function in the presence of time series errors. The nonparametric test is motivated by recent advancement in the theory of ANOVA with large number of factor levels and also utilizes a new difference-based estimation method in nonparametric regression with time-series errors proposed by Hall and Van Keilegom (2003). The test statistic is asymptotically normal under the null and local alternative hypotheses. We also propose a bootstrap method to calculate the critical values and prove its consistency. In a Monte Carlo study, we demonstrate that this bootstrap procedure has good properties for moderate sample size.

Key words: bootstrap, correlated errors, lack-of-fit test, nearest-neighbor windows, nonparametric regression, residual, time-series errors, trend

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1 Introduction

We consider the problem of testing the parametric form of a trend function against an omnibus alternative in the presence of time-series errors. Data (x_{im}, Y_{im}), $i = 1, \ldots, m$, are generated from the following nonparametric regression model

$$Y_{im} = g(x_{im}) + \epsilon_{im},\ \ i = 1, ..., m, \tag{1.1}$$

where Y_{im} is the response, $g(\cdot)$ is an unknown, but smooth regression function, $x_{1m} \leq \ldots \leq x_{mm}$ are fixed design points, and $\epsilon_{1m}, \ldots, \epsilon_{mm}$ are stationary time-series errors. The m in the subscript will be omitted from the notation when no confusion is possible.

Model (1.1) has broad applications. For instance, a large number of studies have been devoted to investigating the presence of global "warming trend" due to green house effects. The standard approach is to assume that $g(x_i) = a + bx_i$ (Woodward and Gray, 1993, 1995), where Y_i is the temperature at time i. It is often proposed to test $b = 0$, the rejection of this null hypothesis generally leads to the belief of the existence of linear trend. Despite its mathematical convenience, there is no special reason to believe why a simple linear trend function would be suitable to model the complex climate system. Taking the linearity assumption as a priori can cause serious bias in both estimation and prediction. Model (1.1) is also used frequently in economics, Y_i can represent stock price, GNP growth rate, consumer price index, etc. Kim and Hart (1998) gave an example of astronomy data, where the successive maxima on the light curve of the variable star T Centaurus are studied.

In this paper, we propose a nonparametric method for testing the following null hypothesis:

$$H_0 :\ \ g(x) \in S_\Theta, \tag{1.2}$$

where $S_\Theta = \{g(\cdot, \theta), \theta \in \Theta\}$ is a parametric family of functions, Θ a subset of a Euclidean space, $g(\cdot, \theta)$ is a function on $\mathcal{R}$. This will allow us to check for example whether the linearity form of the trend function is a valid assumption. Only a very weak smoothness condition is imposed on $g(x)$, thus different from parametric tests which only detect deviations from the null hypothesis in certain directions, the nonparametric test is designed to be powerful against a large class of alternatives, thus in this sense, it is called "omnibus".

When the errors ϵ_i are independent, a large amount of literature has been devoted to testing (1.2), see the manuscript of Hart (1997) for a comprehensive review. The dependent error case, although is common in application, has been discussed relatively little. Related work in this aspect includes Brillinger (1989), González-Manteiga and Vilar-Fernández (1995), Bai (1996), Woodward, Bottone and Gray (1997), Kim and Hart (1998), Vogelsang (1998), Sun and Pantula (1999), Vilar-Fernández and González-Manteiga (2000). The aboved mentioned work, however, is restricted to the case when under the null hypothesis $g(x, \theta)$ is a linear model.

The nonparametric testing procedure proposed by us is a new one: It is motivated by recent advancement in the theory of ANOVA with large number of factor levels and also utilizes a new difference-based estimation method in nonparametric regression with time-series errors proposed by Hall and Van Keilegom (2003). The new test generalizes the work of Wang, Akritas and Van Keilegom (2002) for independent error situation. We also propose a bootstrap procedure to calculate the critical value of the new test which is suitable for moderately large sample size. The consistency of this bootstrap procedure is proved.

The rest of the paper is organized as follows. We introduce the test statistic in Section 2. In Section 3, the asymptotic normality of our test under the null and a local alternative hypothesis is presented. The bootstrap procedure is discussed in Section 4. Numerical simulations, which demonstrate the good level and power behaviors of the bootstrapped test, are reported in Section 5. In the Appendix, we provide the technical arguments.

2 Test Statistic

The disturbances ϵ_{im}, $i = 1, \ldots, m$, form a segment of a covariance-stationary autoregressive process of order p with zero mean, an assumption frequently satisfied in applications. More specifically, we assume $\epsilon_{im} = \sum_{j=1}^{p} \phi_j \epsilon_{i-j,m} + e_{im}$, where $\{e_{im}, -\infty < i < \infty\}$ (or simply e_i) are iid variables with mean 0 and variance σ^2.

Instead of using directly the original data, we will work with the following autoregressive transformation of (x_i, Y_i) under H_0 :

$$Z_i = \left(Y_{i+p} - \sum_{j=1}^{p} \widehat{\phi}_j Y_{i+p-j} \right) - \left(g(x_{i+p}, \widehat{\theta}) - \sum_{j=1}^{p} \widehat{\phi}_j g(x_{i+p-j}, \widehat{\theta}) \right), i = 1, \ldots, n \qquad (2.1)$$

with $n = m - p$, where $\widehat{\phi}_j$ is a $\sqrt{n}$-consistent estimator of ϕ_j, $j = 1, \ldots, p$, and $\widehat{\theta}$ is a $\sqrt{n}$-consistent estimator of θ under the null hypothesis, to be defined at the end of this section. We expect that the Z_i's approximately form a white noise process under H_0. This type of linear filtering has been used to estimate the parameters in the trend function $g(\cdot)$ in the time series literature. Here we propose to apply the test statistic of Wang, Akritas and Van Keilegom (2002) to this transformed sequence (2.1) to obtain an omnibus test for (1.2). This application is new and is different from the related work of González-Manteiga and Vilar-Fernández (1995), Kim and Hart (1998), and Vilar-Fernández and González-Manteiga (2000), who handled correlation directly through asymptotic derivations.

Motivated by recent developments in heteroscedastic ANOVA with large number of factor levels (Akritas and Papadatos, 2001, Wang and Akritas, 2002), Wang, Akritas and Van Keilegom (2002) proposed a new omnibus test for the constant mean null hypothesis in regression model (1.1) with independent errors. The basic idea is to consider each distinct covariate value x_i as a 'category' and construct a window W_i around each x_i consisting of the k_n nearest covariate values (for some k_n going to infinity). That is, they construct an artificial balanced one-way ANOVA with n categories, where the responses in the i-th category are the Y-values that correspond to the covariate values belonging to W_i. In what follows each window W_i will also be understood as the set containing the indices j of the covariate values that belong to the window around x_i, that is (for k_n odd)

$$W_i = \left\{ j : |\widehat{G}(x_j) - \widehat{G}(x_i)| \leq \frac{k_n - 1}{2n} \right\}, \tag{2.2}$$

where $\widehat{G}(x) = \frac{1}{n}\sum_{j=1}^{n} I\,(x_j \leq x)$. To test the null hypothesis of no effects, i.e., $g(x) = C$, where C is an unknown constant, Wang, Akritas and Van Keilegom (2002) suggested looking at $MST - MSE$, where MST is the treatment sum of squares, MSE is the error sum of squares, both computed from the hypothetical one-way ANOVA. This test statistic is thus related to the classical F-test statistic. The test has asymptotic normal law under the null hypothesis and local alternative, it is asymptotically unbiased under the null and has been demonstrated to possess accurate type I error rate and favorable power behaviors in simulations. In the next section, we show that if we apply this testing procedure to the autoregressive transformed data (2.1), the resulting test statistic is also asymptotically normal.

Below we introduce some notations. Using the procedure described above, we create a hypothetical one-way ANOVA (n cells, k_n observations per cell) from (x_i, Z_i), $i = 1, \ldots, n$. We use V_{ij}, $j = 1, \ldots, k_n$, to denote the k_n observations in the i-th category, i.e. $\{V_{i1}, \ldots, V_{ik_n}\} = \{Z_j : j \in W_i\}$. Let $\mathbf{V}$ be the $nk_n \times 1$ vector of all observations in this hypothetical one-way layout. Then our test statistic is defined as:

$$\begin{aligned} T_n &= MST - MSE \\ &= \frac{k_n}{n-1}\sum_{i=1}^{n}\left(\overline{V}_{i.} - \overline{V}_{..}\right)^2 - \frac{1}{n(k_n-1)}\sum_{i=1}^{n}\sum_{j=1}^{k_n}\left(V_{ij} - \overline{V}_{i.}\right)^2, \end{aligned} \tag{2.3}$$

where $\overline{V}_{i.} = k_n^{-1}\sum_{j=1}^{k_n} V_{ij}$ and $\overline{V}_{..} = n^{-1}\sum_{i=1}^{n}\overline{V}_{i.}$. T_n can be expressed as a quadratic form in $\mathbf{V}$: $T_n = \mathbf{V}'\mathbf{A}\mathbf{V}$, with

$$\mathbf{A} = \frac{nk_n - 1}{n(n-1)k_n(k_n-1)}\bigoplus_{i=1}^{n}\mathbf{J}_{k_n} - \frac{1}{n(n-1)k_n}\mathbf{J}_{nk_n} - \frac{1}{n(k_n-1)}\mathbf{I}_{nk_n}, \tag{2.4}$$

where $\mathbf{I}_{k_n}$ is the k_n-dimensional identity matrix, $\mathbf{J}_{k_n} = \mathbf{1}_{k_n}\mathbf{1}'_{k_n}$ with $\mathbf{1}_{k_n}$ standing for the k_n-dimensional column vector of 1's, and $\bigoplus$ is the notation for Kronecker (direct) sum.

We now discuss how to estimate θ and ϕ_j in (2.1). As for the estimation of θ, any $\sqrt{n}$-consistent estimator will work. When $g(x, \theta)$ is a linear function, the method of generalized least squares can be easily applied; in the nonlinear case, a $\sqrt{n}$-consistent estimator of θ was suggested by Gallant and Goebel (1976). Under some appropriate regularity conditions, $\sqrt{n}(\widehat{\theta} - \theta)$ is asymptotically normal.

For estimating ϕ_j, we employ the difference-based estimators newly proposed by Hall and Van Keilegom (2003), but any other $\sqrt{n}$-consistent estimator can be used as well (see e.g. Hart, 1994). The procedure of Hall and Van Keilegom (2003) has however the advantage that it does not depend on a bandwidth parameter. This leads to some computational convenience for the bootstrap version test in Section 4, since we don't need to bring in an extra smoothing parameter. Defining $\gamma(j) = \mathrm{Cov}(\epsilon_i, \epsilon_{i-j})$ and the difference operator D_j by $(D_jY)_i = Y_i - Y_{i-j}$, $\gamma(0)$ and $\gamma(j)$ are estimated by:

$$\begin{aligned} \widehat{\gamma}(0) &= \frac{1}{m_2 - m_1 + 1}\sum_{m=m_1}^{m_2}\frac{1}{2(n-m)}\sum_{i=m+1}^{n}\{(D_mY)_i\}^2, \\ \widehat{\gamma}(j) &= \widehat{\gamma}(0) - \frac{1}{2(n-j)}\sum_{i=j+1}^{n}\{(D_jY)_i\}^2, \quad for\ j \geq 1. \end{aligned} \tag{2.5}$$

The rate at which m_1 and m_2 go to infinity will be specified later. Let $\mathbf{B}$ be the $p \times p$ matrix having $\widehat{\gamma}(j_1 - j_2)$ as its (j_1, j_2)th element, using the Yule-Walker equations, the ϕ's are estimated by:

$$(\widehat{\phi}_1, \ldots, \widehat{\phi}_p)' = \mathbf{B}^{-1}(\widehat{\gamma}(1), \ldots, \widehat{\gamma}(p)). \tag{2.6}$$

Under mild smoothness condition on $g(x)$, Hall and Van Keilegom (2003) proved that:

$$\max_{1\leq j\leq p}|\widehat{\phi}_j-\phi_j|=O_p(m^{-1/2}).$$

3 Large Sample Results

The following assumptions are made to derive the asymptotic distribution of T_n:

Assumption A1. The design points $x_1, x_2, ..., x_n$ on [0,1] satisfy:

$$\int_0^{x_i} r(x)dx=\frac{i}{n},\quad i=1,...,n, \tag{3.1}$$

for some positive design density $r(x)$, which is Lipschitz continuous.

Assumption A2. $g(x,\theta)$ is twice continuously differentiable with respect to x and the components of θ, and

$$\sup_{x,\theta}\left|\frac{\partial}{\partial\theta_j}g(x,\theta)\right|<\infty,\ \sup_{x,\theta}\left|\frac{\partial^2}{\partial x^2}g(x,\theta)\right|<\infty,$$
$$\sup_{x,\theta}\left|\frac{\partial^2}{\partial x\partial\theta_j}g(x,\theta)\right|<\infty,\ \sup_{x,\theta}\left|\frac{\partial^2}{\partial\theta_j\partial\theta_k}g(x,\theta)\right|<\infty,$$

for $j,k=1,\ldots,\dim(\Theta)$.

Assumption A3. The estimator $\widehat{\theta}$ satisfies $\widehat{\theta}-\theta=O_p(n^{-1/2})$ under H_0.

Assumption A4. $\epsilon_{im}=\sum_{j=1}^{p}\phi_j\epsilon_{i-j,m}+e_{im}$, for some fixed integer p, $\{e_{im},-\infty<i<\infty\}$ are iid variables with mean 0, variance σ^2 and finite fourth moment. The constants $\phi_1,\ldots,\phi_p$ are such that the equation $1-\sum_{j=1}^{p}\phi_j z^j=0$ has no complex roots satisfying $|z|\leq 1$.

Assumption A5. $m_1\leq m_2$, $m_1/\log n\to\infty$ and $m_2=O(n^{1/2})$.

The asymptotic distribution of the test statistic of Wang, Akritas and Van Keilegom (2002) applied to the autoregression transformed data (2.1) turns out to have the same form as for the independent case :

Theorem 3.1 *Assume conditions A1-A5. If $n\to\infty$ and $k_n\to\infty$ such that $n^{-1}k_n^{3/2}\to 0$, then under H_0,*

$$\left(\frac{n}{k_n}\right)^{1/2}T_n\xrightarrow{d}N\left(0,\frac{4\sigma^4}{3}\right).$$

Remark 1. The variance σ^2 can be estimated by applying Rice's (1984) estimator to the autoregression transformed data (2.1):

$$\widehat{\sigma}^2 = \frac{1}{2(n-1)} \sum_{i=2}^{n} (Z_i - Z_{i-1})^2. \tag{3.2}$$

Originally, this estimator is proposed for independent observations. The Z_i's mimic a white noise sequence and it is easily shown that $\widehat{\sigma}^2 \xrightarrow{p} \sigma^2$, using the consistency of $\widehat{\theta}$ and $\widehat{\phi}_j$.

Remark 2. We may show that our test statistic can detect local alternatives converging to the null at rate $(nk_n)^{-1/4}$. In the following, we state the asymptotic normality for the case $g(x)$ is linear and $p = 1$, i.e., $\epsilon_i = \phi\epsilon_{i-1} + e_i$. The local alternative sequences we consider are:

$$g(x) = a + bx + (nk_n)^{-1/4} l(x), \tag{3.3}$$

where $l(x)$ is Lipschitz continuous, and satisfies $\int l(x)r(x)dx = \int xl(x)r(x)dx = 0$ and $\int l^2(x)r(x)dx < \infty$. Assume that the conditions of Theorem 3.1 hold, then under the sequence of local alternatives (3.3), we can show that

$$\left(\frac{n}{k_n}\right)^{1/2} T_n \to N\left((1-\phi)^2 \int l^2(x)r(x)dx, \frac{4}{3}\sigma^4\right).$$

The proof of the above result is analogous to that in Wang, Akritas and Van Keilegom (2002). The same proof applies when $g(x)$ has some general parametric form but the bias term will be different.

Remark 3. The above asymptotic results also hold in stochastic design setting where the design points $X_1, \ldots, X_n$ are ordered values of independent and identically distributed random variables whose density is $r(x)$. In random design case, we need the additional assumption that the X_i's and the random errors e_i's are independent.

4 A Bootstrap-based Testing Procedure

As it has been observed by several other authors (González-Manteiga and Vilar-Fernández, 1995, Woodward, Bottone and Gray, 1997, Kim and Hart, 1998, Vilar-Fernández and González-Manteiga, 2000), lack-of-fit tests in regression with time-series errors based on asymptotic distribution tend to overreject in finite sample size, especially when the autoregressive coefficients are positive and close to 1. To avoid size distortion, we propose to bootstrap the test statistic. We adopt the residual bootstrap approach for autoregressions. Consider the following estimator for the noise e_i:

$$\widehat{e}_i = \left(Y_i - \sum_{j=1}^{p} \widehat{\phi}_j Y_{i-j} \right) - \left(g(x_i, \widehat{\theta}) - \sum_{j=1}^{p} \widehat{\phi}_j g(x_{i-j}, \widehat{\theta}) \right)$$

for $i = p+1, ..., m$ (note that $\widehat{e}_i = Z_{i-p}$), where the $\widehat{\phi}$'s are the difference-based estimators of Hall and Van Keilegom (2003), and $\widehat{\theta}$ is a $\sqrt{n}$-consistent estimator of θ under the null hypothesis. Bickel and Freedman (1983) suggested to use the scaled residual $\left(\frac{m}{m-1}\right)^{1/2} \widehat{e}_i$ (an ad hoc action) because the residuals tend to be smaller than the true errors, see also Li and Maddala (1996). Denote the centered $\widehat{e}_i$'s by $\widehat{e}_i^c$ and the empirical distribution function of these $\widehat{e}_i^c$'s by

$$\widehat{F}_n(x) = \frac{1}{n} \sum_{i=p+1}^{m} I(\widehat{e}_i^c \le x). \tag{4.1}$$

From $\widehat{F}_n(x)$ we obtain a sample of independent and identically distributed random variables e_i^*, $i = p+1, \ldots, m$, and construct the error series:

$$\epsilon_i^* = \sum_{j=1}^{p} \widehat{\phi}_j \epsilon_{i-j}^* + e_i^*, \quad \text{for} \quad i = p+1, \ldots, m.$$

(We may take $\epsilon_i^* = \widehat{\epsilon}_i$ or 0 for $i \le p$, we will allow a burn-in period in this step, the effect of the initial value will be negligible.) The bootstrap sample (x_i, Y_i^*), $i = 1, \ldots, n$, is obtained by letting

$$Y_i^* = g(x_i, \widehat{\theta}) + \epsilon_i^*, \quad \text{for} \quad i = 1, \ldots, m.$$

Let $Z_i^*, \widehat{\theta}^*, \widehat{\phi}_j^*$ and T_n^* be calculated from the bootstrap sample in the same way as $Z_i, \widehat{\theta}, \widehat{\phi}_j$ and T_n were calculated from the original observations. Repeat the above steps a large number of times. The critical value of the test is obtained as the upper-α quantile of the bootstrap distribution of $n^{1/2} k_n^{-1/2} T_n^*$. The null hypothesis is rejected if $n^{1/2} k_n^{-1/2} T_n$ is greater than this critical value.

This residual-based scheme for generating bootstrap samples from a stationary time series process was first used in Freedman and Peters (1984) and then by many other authors, including Efron and Tibshirani (1986), Kreiss and Franke (1992), Hjellvik and Tjøstheim (1995), Kim and Hart (1998) and Vilar-Fernández and González-Manteiga (2000). The following theorem shows that the above bootstrap-based test is appropriate in the sense that it gives the asymptotically correct α-level under the null hypothesis.

Theorem 4.1 *Assume conditions A1-A5 and assume in addition that* $\widehat{\theta}^* - \widehat{\theta} = O_{p^*}(n^{-1/2})$. *If* $n \to \infty$ *and* $k_n \to \infty$ *such that* $n^{-1} k_n^{3/2} \to 0$, *then under* H_0,

$$\sup_t |P^*(n^{1/2} k_n^{-1/2} T_n^* \leq t) - P(n^{1/2} k_n^{-1/2} T_n \leq t)| = o_P(1),$$

where P^* *denotes the probability measure conditional on the original sample.*

We assume the order of autoregression p is known. Kreiss (1997) has shown that the above residual bootstrap also works in the case the order of the fitted autoregressive model may depend on the data, for instance, it may be estimated by a certain order selection procedure.

Note that the same bootstrap testing procedure can also be applied to the random design setting discussed in Remark 3 of the previous section. We replace in that case the x_i's by a bootstrap sample from their empirical distribution, independent of the e_i^*'s.

5 Numerical Results

In this section, we will summarize the results from a Monte Carlo study, which will shed light on the finite sample size and power behavior of our test statistic. The data (x_i, Y_i), $i = 1, \ldots, m$, are generated from model (1.1). We assume that the errors follow a stationary AR(1) process: $\epsilon_i = \phi \epsilon_{i-1} + e_i$, where e_i, $i = 1, \ldots, m$, are an independent and identically distributed sequence of normal random variables with zero mean and variance σ^2. We test the no effect null hypothesis $g(x) = C$. The design points x_i are equal-distant on [0,1].

The bootstrap test is computed as described in the previous section. We generate 500 samples of observations, for each sample 500 bootstrap samples are produced. The autoregressive coefficient is estimated using the estimators of Hall and Van Keilegom (2003). Following their advice, we take $m_1 = n^{0.1}$ and $m_2 = n^{0.5}$.

We first investigate the empirical size of our bootstrap test. To reveal the seriousness of the size distortion when the correlation is ignored, we include the results for the nonparametric test of Wang, Akritas and Van Keilegom (2002), which assumes that the data are independent. We also make a comparison with the test of Vogelsang (1998), which is based on partial sum regressions. This test does not require estimates of serial correlation coefficient and can handle unit root.

Data under the null hypothesis are generated by setting $g(x) = 0$. Table 1 gives the observed size of the bootstrap test when the significance level α equals 0.05, $\phi = -0.6, -0.4, -0.2, 0, 0.2, 0.4, 0.6$, and $\sigma = 0.5$, for sample size $n = m - 1 = 100$ or 200. For the new test and the test assuming independence, different local window sizes k_n are considered.

Put Table 1 about here

We observe that both the bootstrap test and Vogelsang's test give quite accurate type I error rates, while for the test assuming independence the size is much too conservative or much too liberal. When $n = 100$, the empirical size is a little liberal for $\phi = 0.6$ but it becomes close to the desired level when $n = 200$. We also tried $\sigma = 0.2$ and $\sigma = 1$ and found similar results (not reported). Also, note that the choice of k_n does not have a large influence on the result. For ϕ close to one (the correlation is positive and strong in that case), the observed type I error is often quite large (roughly around 0.20, 0.12 for $\phi = 0.8$, $n = 100$ and 200, respectively, in the setting of Table 1). This phenomenon has also been observed by González-Manteiga and Vilar-Fernández (1995), Woodward, Bottone and Gray (1997), Kim and Hart (1998) and Vilar-Fernández and González-Manteiga (2000). Woodward, Bottone and Gray (1997) conjectured that this is because the usual estimate of ϕ is biased towards zero. This is also the case here : when m_1 is small it is clear that $\widehat{\gamma}(0)$ defined in (2.5) will be biased, especially when ϕ is close to one, and hence $\widehat{\phi}$ is also biased. Better results for $\phi = 0.8$ can be obtained when the value of m_1 is increased.

For the power analysis, we consider two types of alternatives: linear alternative, $g(x) = 1 + 2x$ and sinusoidal alternative, $g(x) = \cos(2x)$. The simulation results are displayed in Table 2, for $\sigma = 1$, $\phi = 0, 0.2, 0.4, 0.6$ and sample size $n = 100$. Vogelsang's test is more powerful in the first case since it is designed to detect linear deviation from constant mean. As we can see, Vogelsang's test has no power to detect the oscillating cos alternative while our bootstrap based nonparametric test gives good power in this case.

Put Table 2 about here

6 Proofs

We'll give the proof of Theorem 3.1 and Theorem 4.1 for the case $\dim(\Theta) = 1$ and $p = 1$, i.e., $\epsilon_i = \phi\epsilon_{i-1} + e_i$, $i = 1, \ldots, n$, for some $|\phi| < 1$. The case $\dim(\Theta) > 1$ and $p > 1$ can be proved similarly with extra amount of notations.

Proof of Theorem 3.1. We have

$$\begin{aligned} Z_i &= (Y_{i+1} - \widehat{\phi}Y_i) - \left(g(x_{i+1}, \widehat{\theta}) - \widehat{\phi}g(x_i, \widehat{\theta})\right) \\ &= (Y_{i+1} - g(x_{i+1}, \theta)) + (g(x_{i+1}, \theta) - g(x_{i+1}, \widehat{\theta})) - \phi(Y_i - g(x_i, \theta)) \\ &\quad -\phi(g(x_i, \theta) - g(x_i, \widehat{\theta})) + (\phi - \widehat{\phi})(Y_i - g(x_i, \theta)) + (\phi - \widehat{\phi})(g(x_i, \theta) - g(x_i, \widehat{\theta})) \\ &= e_{i+1} + (g(x_{i+1}, \theta) - g(x_{i+1}, \widehat{\theta})) - \phi(g(x_i, \theta) - g(x_i, \widehat{\theta})) \\ &\quad +(\phi - \widehat{\phi})(Y_i - g(x_i, \theta)) + (\phi - \widehat{\phi})(g(x_i, \theta) - g(x_i, \widehat{\theta})), \end{aligned} \tag{6.1}$$

for $i = 1, \ldots, n$. Denote $\eta_1(x_i) = g(x_{i+1}, \theta) - g(x_{i+1}, \widehat{\theta})$, $\eta_2(x_i) = \phi(g(x_i, \theta) - g(x_i, \widehat{\theta}))$, $\eta_3(x_i) = (\phi - \widehat{\phi})(Y_i - g(x_i, \theta))$ and $\eta_4(x_i) = (\phi - \widehat{\phi})(g(x_i, \theta) - g(x_i, \widehat{\theta}))$. Let $\boldsymbol{\eta}_1$ be the $nk_n \times 1$ vector of all observations in the hypothetical one-way ANOVA constructed from $(x_i, \eta_1(x_i))$, $i = 1, \ldots, n$. Similarly, we define $\boldsymbol{\eta}_2$, $\boldsymbol{\eta}_3$, $\boldsymbol{\eta}_4$, $\mathbf{e}$ and $\mathbf{Z}$. By (6.1), $\mathbf{Z} = \mathbf{e} + \boldsymbol{\eta}_1 - \boldsymbol{\eta}_2 + \boldsymbol{\eta}_3 + \boldsymbol{\eta}_4$. Thus our test statistic T_n, which is equal to $\mathbf{Z}'\mathbf{A}\mathbf{Z}$ in this new notation, can be decomposed as:

$$\begin{aligned} T_n &= \mathbf{e}'\mathbf{A}\mathbf{e} + 2\mathbf{e}'\mathbf{A}\boldsymbol{\eta}_1 - 2\mathbf{e}'\mathbf{A}\boldsymbol{\eta}_2 + 2\mathbf{e}'\mathbf{A}\boldsymbol{\eta}_3 + 2\mathbf{e}'\mathbf{A}\boldsymbol{\eta}_4 + \boldsymbol{\eta}_1'\mathbf{A}\boldsymbol{\eta}_1 - 2\boldsymbol{\eta}_1'\mathbf{A}\boldsymbol{\eta}_2 + 2\boldsymbol{\eta}_1'\mathbf{A}\boldsymbol{\eta}_3 \\ &\quad +2\boldsymbol{\eta}_1'\mathbf{A}\boldsymbol{\eta}_4 + \boldsymbol{\eta}_2'\mathbf{A}\boldsymbol{\eta}_2 - 2\boldsymbol{\eta}_2'\mathbf{A}\boldsymbol{\eta}_3 - 2\boldsymbol{\eta}_2'\mathbf{A}\boldsymbol{\eta}_4 + \boldsymbol{\eta}_3'\mathbf{A}\boldsymbol{\eta}_3 + 2\boldsymbol{\eta}_3'\mathbf{A}\boldsymbol{\eta}_4 + \boldsymbol{\eta}_4'\mathbf{A}\boldsymbol{\eta}_4. \end{aligned} \tag{6.2}$$

Since $\{e_i\}$ are iid, by the result of Wang, Akritas and Van Keilegom (2002), $n^{1/2}{k_n}^{-1/2}\mathbf{e}'\mathbf{A}\mathbf{e} \xrightarrow{d} N\left(0, \frac{4\sigma^2}{3}\right)$, where $\sigma^2 = Var(e_i)$. We need to show that all the other 14 terms in (6.2) are asymptotically negligible. As representative examples, we'll show

$$n^{1/2}k_n^{-1/2}\mathbf{e}'\mathbf{A}\boldsymbol{\eta}_1 \xrightarrow{p} 0, \quad n^{1/2}k_n^{-1/2}\mathbf{e}'\mathbf{A}\boldsymbol{\eta}_3 \xrightarrow{p} 0, \quad n^{1/2}k_n^{-1/2}\boldsymbol{\eta}_3'\mathbf{A}\boldsymbol{\eta}_3 \xrightarrow{p} 0. \tag{6.3}$$

We will first show that $n^{1/2}k_n^{-1/2}\mathbf{e}'\mathbf{A}\boldsymbol{\eta}_1 \xrightarrow{p} 0$.

$$\begin{aligned} &\mathbf{e}'\mathbf{A}\boldsymbol{\eta}_1 \\ =\ & \frac{nk_n - 1}{n(n-1)k_n(k_n-1)} \sum_{i=1}^{n} \left[\sum_{j=1}^{n} \eta_1(x_j) I(j \in W_i)\right] \left[\sum_{k=1}^{n} e_{k+1} I(k \in W_i)\right] \\ & -\frac{1}{n(n-1)k_n} \left[k_n \sum_{i=1}^{n} \eta_1(x_i)\right] \left[k_n \sum_{k=1}^{n} e_{k+1}\right] - \frac{k_n}{n(k_n-1)} \sum_{i=1}^{n} \eta_1(x_i) e_{i+1} \end{aligned}$$

$$
\begin{aligned}
&= \frac{nk_n - 1}{n(n-1)(k_n-1)} \sum_{i=1}^n \eta_1(x_i) \left[\sum_{k=1}^n e_{k+1} I(k \in W_i) \right] \\
&\quad - \frac{k_n}{n(n-1)} \left[\sum_{i=1}^n \eta_1(x_i) \right] \left[\sum_{k=1}^n e_{k+1} \right] - \frac{k_n}{n(k_n-1)} \sum_{i=1}^n \eta_1(x_i) e_{i+1} \\
&\quad + \frac{nk_n - 1}{n(n-1)(k_n-1)} O_p(n^{-3/2} k_n) \sum_{i=1}^n \left| \sum_{k=1}^n e_{k+1} I(k \in W_i) \right|, \qquad (6.4)
\end{aligned}
$$

since

$$
\begin{aligned}
|\eta_1(x_i) - \eta_1(x_j)| &\leq \sup_{\theta,x} \left| \frac{\partial^2}{\partial\theta\partial x} g(x,\theta) \right| |\widehat{\theta} - \theta| \, |x_{i+1} - x_{j+1}| \\
&= O_p(n^{-3/2})
\end{aligned}
$$

uniformly in $1 \leq i \leq n$ and $j \in W_i$. It follows that (6.4) equals

$$
\begin{aligned}
&\frac{(nk_n-1)k_n}{n(n-1)(k_n-1)} \sum_{k=1}^n \eta_1(x_k) e_{k+1} - \frac{k_n}{n(n-1)} \left[\sum_{i=1}^n \eta_1(x_i) \right] \left[\sum_{k=1}^n e_{k+1} \right] \\
&- \frac{k_n}{n(k_n-1)} \sum_{i=1}^n \eta_1(x_i) e_{i+1} + \frac{nk_n-1}{n(n-1)(k_n-1)} O_p(n^{-3/2} k_n^2) \sum_{k=1}^n |e_{k+1}| \\
= \ & \frac{k_n}{n-1} \sum_{k=1}^n \left(\eta_1(x_k) - n^{-1} \sum_{i=1}^n \eta_1(x_i) \right) e_{k+1} + \frac{nk_n-1}{n(n-1)(k_n-1)} O_p(n^{-3/2} k_n^2) \sum_{k=1}^n |e_{k+1}| \\
= \ & \frac{k_n}{n-1} (\theta - \widehat{\theta}) \sum_{k=1}^n \left(\frac{\partial}{\partial\theta} g(x_{k+1}, \theta) - n^{-1} \sum_{i=1}^n \frac{\partial}{\partial\theta} g(x_{i+1}, \theta) \right) e_{k+1} \\
&+ \frac{k_n}{2(n-1)} (\theta - \widehat{\theta})^2 \sum_{k=1}^n \left(\frac{\partial^2}{\partial\theta^2} g(x_{k+1}, \tilde{\theta}) - n^{-1} \sum_{i=1}^n \frac{\partial^2}{\partial\theta^2} g(x_{i+1}, \tilde{\theta}) \right) e_{k+1} + O_p(n^{-3/2} k_n^2) \\
= \ & O_p(n^{-1} k_n) + O_p(n^{-3/2} k_n^2)
\end{aligned}
$$

where $\tilde{\theta}$ is between θ and $\widehat{\theta}$, and where the last step follows from the central limit theorem for triangular arrays (use e.g. Liapunov's condition and the fact that $E|e_i|^4 < \infty$). Thus $n^{1/2} k_n^{-1/2} \mathbf{e}' \mathbf{A} \boldsymbol{\eta}_1 = O_p(n^{-1/2} k_n^{1/2}) + O_p(n^{-1} k_n^{3/2}) = o_p(1)$. We now look at $\mathbf{e}' \mathbf{A} \boldsymbol{\eta}_3$:

$$
\begin{aligned}
&\mathbf{e}' \mathbf{A} \boldsymbol{\eta}_3 \\
= \ & \frac{nk_n-1}{n(n-1)k_n(k_n-1)} (\phi - \widehat{\phi}) \sum_{i=1}^n \left[\sum_{j=1}^n \epsilon_j I(j \in W_i) \right] \left[\sum_{k=1}^n e_{k+1} I(k \in W_i) \right] \\
&- \frac{k_n}{n(n-1)} (\phi - \widehat{\phi}) \left[\sum_{i=1}^n \epsilon_i \right] \left[\sum_{k=1}^n e_{k+1} \right] - \frac{k_n}{n(k_n-1)} (\phi - \widehat{\phi}) \sum_{i=1}^n \epsilon_i e_{i+1}
\end{aligned}
$$

$$= \quad D_1 - D_2 - D_3,$$

where the definition of D_i, $i = 1, 2, 3$, should be clear from the context. Evaluate D_3 first. Since $\epsilon_j = \phi\epsilon_{j-1} + e_j$ and since the e_i's $(-\infty < i < \infty)$ are independent, e_{i+1} is independent of ϵ_j for $j \le i$. Hence,

$$\begin{aligned} E\left(\sum_{i=1}^{n} \epsilon_i e_{i+1}\right)^2 &= \sum_{i=1}^{n} E(\epsilon_i^2 e_{i+1}^2) + 2E\left(\sum_{i<i'} \epsilon_i \epsilon_{i'} e_{i+1} e_{i'+1}\right) \\ &= \frac{n\sigma^4}{1-\phi^2} + 0 = O(n). \end{aligned}$$

Thus $n^{1/2}k_n^{-1/2}D_3 = n^{1/2}k_n^{-1/2}O_p(n^{-1}) = O_p(n^{-1/2}k_n^{-1/2}) = o_p(1)$. Now check D_2. Note that $\sum_{k=1}^{n} e_{k+1} = O_p(n^{1/2})$ and $\sum_{i=1}^{n} \epsilon_i = O_p(n)$, and so $D_2 = O_p(n^{-1}k_n) = o_p(n^{-1/2}k_n^{1/2})$. Finally consider D_1. We will first evaluate the order of $\sum_{j,k}\sum_{j',k'} E(\epsilon_j \epsilon_{j'} e_{k+1} e_{k'+1})$. Consider three cases for the four indices in the sum: (1) (j, j', k, k') contains exactly two pairs, the order of such terms is $O(n^2)$; (2) (j, j', k, k') contains exactly one pair, the order of such terms is $O(n^3)$; (3) there is no pair in (j, j', k, k'). For this case we have:

$$\begin{aligned} &\sum_{j,j',k,k' \text{ are distinct}}^{n} E(\epsilon_j \epsilon_{j'} e_{k+1} e_{k'+1}) \\ &= 4\sum_{j<j'} \sum_{k<k', j,j',k,k' \text{ are distinct}} E(\epsilon_j \epsilon_{j'} e_{k+1} e_{k'+1}) \\ &= 4\sum_{k<k'<j<j'} E(\epsilon_j \epsilon_{j'} e_{k+1} e_{k'+1}) + 4\sum_{k<j<k'<j'} E(\epsilon_j \epsilon_{j'} e_{k+1} e_{k'+1}) \\ &\quad + 4\sum_{j<k<k'<j'} E(\epsilon_j \epsilon_{j'} e_{k+1} e_{k'+1}) \\ &= Q_1 + Q_2 + Q_3, \end{aligned}$$

where the definition of Q_i, for $i = 1, 2, 3$, is clear from the context. Look at Q_1 first. Since for any k, $\epsilon_j = \phi^{j-k}\epsilon_k + \sum_{\ell=1}^{j-k} \phi^{j-k-\ell} e_{k+\ell}$, $\epsilon_{j'} = \phi^{j'-k}\epsilon_k + \sum_{\ell=1}^{j'-k} \phi^{j'-k-\ell} e_{k+\ell}$, the only terms in the product $\epsilon_j \epsilon_{j'} e_{k+1} e_{k'+1}$ with nonzero expectation are $\phi^{j-k-1} e_{k+1}^2 \phi^{j'-k'-1} e_{k'+1}^2$ and $\phi^{j-k'-1} e_{k'+1}^2 \phi^{j'-k-1} e_{k+1}^2$. Thus,

$$Q_1 = O\left(\sum_{k<k'<j<j'}^{n} \phi^{j+j'-k-k'-2}\right)$$

$$
\begin{aligned}
&= O\left(\sum_{j'=4}^{n}\sum_{j=3}^{j'-1}\sum_{k'=2}^{j-1}\sum_{k=1}^{k'-1}\phi^{j+j'-k-k'-2}\right) \\
&= O\left(\sum_{j'=4}^{n}\sum_{j=3}^{j'-1}\sum_{k'=2}^{j-1}\phi^{j+j'-k'-2}\frac{\phi^{-(k'-1)}-1}{1-\phi}\right) \qquad (6.5)\\
&= O\left(\sum_{j'=4}^{n}\sum_{j=3}^{j'-1}\sum_{k'=2}^{j-1}\phi^{j+j'-2k'-1}\right) \\
&= O\left(\sum_{j'=4}^{n}\sum_{j=3}^{j'-1}\phi^{j+j'-1}\frac{\phi^{-2(j-2)}-1}{\phi^2-\phi^4}\right) \\
&= O(n^2),
\end{aligned}
$$

since $|\phi^{j+j'-1}\frac{\phi^{-2(j-2)}-1}{\phi^2-\phi^4}| \leq 1$ is bounded. Q_2 and Q_3 can be evaluated similarly. It now follows that similarly as above

$$
E\left[\sum_{i=1}^{n}\sum_{i'=1}^{n}\sum_{j,k}\sum_{j',k'}\epsilon_j\epsilon_{j'}e_{k+1}e_{k'+1}I(j,k\in W_i)I(j',k'\in W_{i'})\right] = O(n^2k_n^2),
$$

hence $n^{1/2}k_n^{-1/2}D_1 = O_p(k_n^{-1/2}) = o_p(1)$. We thus obtain $n^{1/2}k_n^{-1/2}\mathbf{e}'\mathbf{A}\boldsymbol{\eta}_3 \xrightarrow{p} 0$. At last, we evaluate $\boldsymbol{\eta}_3'\mathbf{A}\boldsymbol{\eta}_3$:

$$
\begin{aligned}
&\boldsymbol{\eta}_3'\mathbf{A}\boldsymbol{\eta}_3 \\
= \;& \frac{nk_n-1}{n(n-1)k_n(k_n-1)}(\phi-\widehat{\phi})^2\sum_{i=1}^{n}\left[\sum_{j=1}^{n}\epsilon_j I(j\in W_i)\right]^2 \\
&-\frac{k_n}{n(n-1)}(\phi-\widehat{\phi})^2\left[\sum_{i=1}^{n}\epsilon_i\right]^2 - \frac{k_n}{n(k_n-1)}(\phi-\widehat{\phi})^2\sum_{i=1}^{n}\epsilon_i^2 \\
= \;& (\phi-\widehat{\phi})^2(E_1-E_2-E_3) = O_p(n^{-1})(E_1-E_2-E_3),
\end{aligned}
$$

where the definition of E_1, E_2 and E_3 is clear from the context. Look at E_3 first : $E_3 \geq 0$ and

$$
\begin{aligned}
E(E_3) &= \frac{k_n}{n(k_n-1)}\sum_{i=1}^{n}E(\epsilon_i^2) \\
&= O(n^{-1})\frac{n\sigma^2}{1-\phi^2} = O(1),
\end{aligned}
$$

thus $n^{1/2}k_n^{-1/2}O_p(n^{-1})E_3 = o_p(1)$. Next, $E_2 \geq 0$ and $E(E_2) = O(n^{-2}k_n n^2) = O(k_n)$. So $n^{1/2}k_n^{-1/2}O_p(n^{-1})E_2 = O_p(n^{-1/2}k_n^{1/2}) = o_p(1)$. Now check E_1 : $E_1 \geq 0$ and $E(E_1) =$

$O(n^{-1}k_n^{-1}nk_n^2) = O(k_n)$, and so $O_p(n^{-1})E_1 = O_p(n^{-1}k_n) = o_p(n^{-1/2}k_n^{1/2})$. The derivation for the remaining terms in (6.2) is similar or easier than the derivation for the three terms we have considered above. Hence, the result follows. □

To prove Theorem 4.1, the following two lemmas are useful:

Lemma 6.1 *Let F be the cumulative distribution function of e_i. Then, for any $k \geq 1$, for $\widehat{F}_n$ defined in (4.1), and under H_0,*

$$d_k(\widehat{F}_n, F) \to 0, \ \textit{in} \ \ \textit{probability},$$

where d_k denotes the Mallow's distance (for its definition and properties, refer to Shao and Tu (1995), pp. 73-74).

Proof. Denote by F_n the usual empirical distribution function based on the unobserved $e_2, \ldots, e_m$. By Bickel and Freedman (1981, Lemma 8.4),

$$d_k(F_n, F) \to 0, \text{as } n \to \infty, \ \text{almost} \ \ \text{everywhere}. \tag{6.6}$$

Let J be the Laplace distribution in $\{2, \ldots, m\}$, that is $J = j$ with probability $\frac{1}{n}$, $j = 2, \ldots, m$. We define random variables U_1 and V_1 with marginals F_n and $\widehat{F}_n$ respectively according to $U_1 = e_J$, $V_1 = \widehat{e}_J$. We have

$$d_k^k(F_n, \widehat{F}_n) = \inf_{U,V} E(U - V)^k \leq E(U_1 - V_1)^k = \frac{1}{n} \sum_{j=2}^{m} (e_j - \widehat{e}_j)^k,$$

where the infimum is taken over all random variables U and V that have distribution F_n and $\widehat{F}_n$ respectively. Under H_0, $e_j = (Y_j - \phi Y_{j-1}) - (g(x_j, \theta) - \phi g(x_{j-1}, \theta))$, $\widehat{e}_j = (Y_j - \widehat{\phi} Y_{j-1}) - (g(x_j, \widehat{\theta}) - \widehat{\phi} g(x_{j-1}, \widehat{\theta}))$, and so we have

$$\begin{aligned} e_j - \widehat{e}_j &= (\widehat{\phi} - \phi) Y_{j-1} + (g(x_j, \widehat{\theta}) - g(x_j, \theta)) \\ &\quad - (\widehat{\phi} - \phi) g(x_{j-1}, \widehat{\theta}) - \phi(g(x_{j-1}, \widehat{\theta}) - g(x_{j-1}, \theta)). \end{aligned}$$

From the consistency of $\widehat{\theta}$ and $\widehat{\phi}$ (see Hall and Van Keilegom (2003) for the latter), it is easy to show that $\frac{1}{n} \sum_{j=2}^{m} (e_j - \widehat{e}_j)^k = o_p(1)$ and thus $d_k(F_n, \widehat{F}_n) = o_p(1)$. Combined with (6.6), we finish the proof. □

Lemma 6.2 *Under the conditions of Theorem 4.1,*

$$\max_{1 \leq j \leq p} |\widehat{\phi}_j^* - \widehat{\phi}_j| = O_{p^*}(n^{-1/2})$$

($j = 1, \ldots, p$).

Proof. We will show that $\sup_{0\le j\le n}|\widehat{\gamma}^*(j)-\widehat{\gamma}(j)|=O_{p^*}(n^{-1/2})$, where $\widehat{\gamma}^*(j)$ is obtained by replacing Y by Y^* in the definition of $\widehat{\gamma}(j)$. The result then follows immediately. For showing this, we closely follow the proof of statement (2.11) in Hall and Van Keilegom (2003) up to the point where it is shown that $\widehat{\gamma}(j)=\gamma(j)+\Delta_j+o_p(n^{-1/2})$ and $\Delta_j=O_p(n^{-1/2})$. The proof of their equations (4.1)-(4.3) and of the rate of Δ_j follows from calculations similar to the ones in (6.5). In the bootstrap world the analogue of this derivation continues to hold provided that $\sup_{1\le j\le p}|\widehat{\phi}_j-\phi_j|=o_p(1)$ and $E^*|e_j^*|^k\xrightarrow{p}E|e_j|^k$ for $k=1,2,3,4$. The former assertion follows from Hall and Van Keilegom (2003), the latter follows from Lemma 6.1 in combination with Lemma 8.3 in Bickel and Freedman (1981). □

Proof of Theorem 4.1. We have

$$\begin{aligned}T_n^* &= (\mathbf{e}^*)'\mathbf{A}\mathbf{e}^*+2(\mathbf{e}^*)'\mathbf{A}\boldsymbol{\eta}_1^*-2(\mathbf{e}^*)'\mathbf{A}\boldsymbol{\eta}_2^*+2(\mathbf{e}^*)'\mathbf{A}\boldsymbol{\eta}_3^*+2(\mathbf{e}^*)'\mathbf{A}\boldsymbol{\eta}_4^*+(\boldsymbol{\eta}_1^*)'\mathbf{A}\boldsymbol{\eta}_1^*\\ &\quad -2(\boldsymbol{\eta}_1^*)'\mathbf{A}\boldsymbol{\eta}_2^*+2(\boldsymbol{\eta}_1^*)'\mathbf{A}\boldsymbol{\eta}_3^*+2(\boldsymbol{\eta}_1^*)'\mathbf{A}\boldsymbol{\eta}_4^*+(\boldsymbol{\eta}_2^*)'\mathbf{A}\boldsymbol{\eta}_2^*-2(\boldsymbol{\eta}_2^*)'\mathbf{A}\boldsymbol{\eta}_3^*-2(\boldsymbol{\eta}_2^*)'\mathbf{A}\boldsymbol{\eta}_4^*\\ &\quad +(\boldsymbol{\eta}_3^*)'\mathbf{A}\boldsymbol{\eta}_3^*+2(\boldsymbol{\eta}_3^*)'\mathbf{A}\boldsymbol{\eta}_4^*+(\boldsymbol{\eta}_4^*)'\mathbf{A}\boldsymbol{\eta}_4^*,\end{aligned}\tag{6.7}$$

where $\eta_1^*(x_i)=g(x_{i+1},\widehat{\theta})-g(x_{i+1},\widehat{\theta}^*)$, $\eta_2^*(x_i)=\widehat{\phi}(g(x_i,\widehat{\theta})-g(x_i,\widehat{\theta}^*))$, $\eta_3^*(x_i)=(\widehat{\phi}-\widehat{\phi}^*)(Y_i^*-g(x_i,\widehat{\theta}))$, $\eta_4^*(x_i)=(\widehat{\phi}-\widehat{\phi}^*)(g(x_i,\widehat{\theta})-g(x_i,\widehat{\theta}^*))$, and $\boldsymbol{\eta}_j^*$ $(j=1,2,3,4)$ (respectively $\mathbf{e}^*$) is the $nk_n\times 1$ vector of all observations in the hypothetical one-way ANOVA constructed from $(x_i,\eta_j^*(x_i))$ (respectively (x_i,e_i^*)), $i=1,\ldots,n$. Similarly as in the proof of Theorem 2.2 in Wang, Akritas and Van Keilegom (2002), we can show that

$$\left(\frac{n}{k_n}\right)^{1/2}(V^*)^{-1/2}(\mathbf{e}^*)'\mathbf{A}\mathbf{e}^*\xrightarrow{d^*}N\left(0,1\right),$$

where

$$V^*=\frac{2}{n(k_n-1)^2k_n}\sum_{i_1=1}^n\sum_{i_2=1}^n\sum_{j\neq l}Var^*(e_j^*)Var^*(e_l^*)I(j,l\in W_{i_1}\cap W_{i_2}).$$

By Lemma 8.3 in Bickel and Freedman (1981) and Lemma 6.1, $\mathrm{Var}^*(e_j^*)\xrightarrow{p}\sigma^2$. It is thus easy to show that $V^*\xrightarrow{p}\frac{4}{3}\sigma^4$. It remains to show that all the other terms in (6.7) are $o_{p^*}(1)$. First note that the consistency of $\widehat{\theta}^*$ and $\widehat{\phi}_j^*$ follows from the statement of the theorem and Lemma 6.2 respectively. The proof is now similar to that of Theorem 3.1 (for the verification of Liapunov's condition, use the fact that $E^*|e_j^*|^4\xrightarrow{p}E|e_j|^4<\infty$, and for derivation (6.5) make use of $\widehat{\phi}-\phi\xrightarrow{p}0$ and $\mathrm{Var}^*(e_j^*)\xrightarrow{p}\sigma^2$). This finishes the proof. □

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Table 1: Empirical level of the new test, the test of Wang, Akritas and Van Keilegom (2002) for independent data and the test of Vogelsang (1998) ($\alpha = 0.05$).

		sample size $n = 100$						
		ϕ						
Test	k_n	-0.6	-0.4	-0.2	0.0	0.2	0.4	0.6
New	5	0.066	0.048	0.054	0.062	0.050	0.064	0.088
	7	0.058	0.050	0.046	0.068	0.046	0.072	0.084
	9	0.056	0.048	0.044	0.066	0.054	0.070	0.092
Indep.	5	0.000	0.000	0.002	0.065	0.537	0.954	0.996
	7	0.000	0.000	0.002	0.050	0.437	0.892	0.992
	9	0.000	0.000	0.003	0.048	0.377	0.828	0.986
Vogelsang		0.064						

		sample size $n = 200$						
		ϕ						
Test	k_n	-0.6	-0.4	-0.2	0.0	0.2	0.4	0.6
New	7	0.056	0.062	0.040	0.050	0.062	0.050	0.066
	9	0.056	0.064	0.048	0.040	0.060	0.048	0.058
	11	0.054	0.066	0.060	0.050	0.060	0.048	0.066
Indep.	7	0.000	0.000	0.002	0.070	0.644	0.992	1.000
	9	0.000	0.000	0.002	0.060	0.560	0.972	1.000
	11	0.000	0.000	0.002	0.056	0.514	0.950	1.000
Vogelsang		0.047						

Table 2: Power of the new test and the test of Vogelsang (1998) for $n = 100$ and $\alpha = 0.05$.

$g(x) = 1 + 2x$

Test	k_n	ϕ = 0.0	0.2	0.4	0.6
New	5	0.975	0.858	0.618	0.498
	7	0.980	0.840	0.605	0.438
	9	0.983	0.843	0.590	0.433
Vogelsang		0.812			

$g(x) = \cos(2x)$

Test	k_n	ϕ = 0.0	0.2	0.4	0.6
New	5	0.863	0.690	0.455	0.303
	7	0.863	0.668	0.435	0.275
	9	0.860	0.680	0.438	0.260
Vogelsang		0.000			