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EMPIRICAL PROCESS OF THE SQUARED RESIDUALS OF AN ARCH SEQUENCE[1]

BY LAJOS HORVÁTH, PIOTR KOKOSZKA AND GILLES TEYSSIÈRE

University of Utah, University of Liverpool, European Commission and GREQAM

We derive the asymptotic distribution of the sequential empirical process of the squared residuals of an ARCH(p) sequence. Unlike the residuals of an ARMA process, these residuals do not behave in this context like asymptotically independent random variables, and the asymptotic distribution involves a term depending on the parameters of the model. We show that in certain applications, including the detection of changes in the distribution of the unobservable innovations, our result leads to asymptotically distribution free statistics.

1. Introduction and results. Procedures based on the empirical distribution function of independent identically distributed observations occupy a central place in statistical inference; see Shorack and Wellner (1986). For time series data, residuals must be considered, and since these necessarily depend on parameter estimates, the asymptotic theory for the empirical distribution function is more complex in such cases. Nevertheless, inference based on residuals, especially model goodness-of-fit tests and various diagnostic checks, is a fundamental tool in the statistical analysis of *linear* time series models; see Brockwell and Davis (1991). By contrast, large sample theory for the residuals of nonlinear time series models is much less developed. Li and Mak (1994) and Horváth and Kokoszka (2001) study squared residual autocorrelations of ARCH sequences, whose importance in various specification tests was demonstrated by Lundbergh and Teräsvirta (1998). Tjøstheim (1999) considers nonparametric tests based on squared residuals.

In this paper we consider the ARCH(p) model defined by the equations

$$y_t = \sigma_t \varepsilon_t, \qquad \sigma_t^2 = b_0 + \sum_{j=1}^{p} b_j y_{t-j}^2,$$

where $\{\varepsilon_t, -\infty < i < \infty\}$ are independent identically distributed random variables with

$$E\varepsilon_0 = 0 \quad \text{and} \quad E\varepsilon_0^2 = 1.$$

We assume that $\mathbf{b} = (b_0, b_1, \dots, b_p)$ is the parameter vector satisfying

$$b_0 > 0, \qquad b_i \geq 0, \qquad 1 \leq i \leq p.$$

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The distribution function of ε_0^2 will be denoted by F and we assume that

$$f(t) = F'(t) \text{ exists and is continuous on } (0, \infty), \tag{1.1}$$

$$\lim_{t\to 0} tf(t) = 0 \quad \text{and} \quad \lim_{t\to\infty} tf(t) = 0. \tag{1.2}$$

We assume that the parameter vector $\mathbf{b}$ is estimated from a sample $y_1, y_2, \ldots, y_n$ by an estimator $\hat{\mathbf{b}}_n = (\hat{b}_0, \hat{b}_1, \ldots, \hat{b}_p)$ which admits the representation

$$\hat{b}_i - b_i = \frac{1}{n} \sum_{1\le j\le n} l_i(\varepsilon_j^2) f_i(\varepsilon_{j-1}, \varepsilon_{j-2}, \ldots) + o_P(n^{-1/2}), \qquad 0 \le i \le p. \tag{1.3}$$

The functions l_i and f_i are regular in the sense that

$$El_i(\varepsilon_0^2) = 0, \qquad 0 \le i \le p, \tag{1.4}$$

$$E\big[l_i(\varepsilon_0^2)\big]^2 < \infty, \qquad 0 \le i \le p \tag{1.5}$$

and

$$E[f_i(\varepsilon_0, \varepsilon_{-1}, \ldots)]^2 < \infty, \qquad 0 \le i \le p. \tag{1.6}$$

We show in Section 2 that commonly used estimators [see, e.g., Chapter 4 of Gouriéroux (1997)] admit representation (1.3).

The squared residuals are defined as

$$\hat{\varepsilon}_k^2 = \frac{y_k^2}{\hat{\sigma}_k^2}, \qquad p < k \le n,$$

where

$$\hat{\sigma}_k^2 = \hat{b}_0 + \sum_{1\le j\le p} \hat{b}_j y_{k-j}^2, \qquad p < k \le n.$$

In this paper we study the weak convergence of the sequential (or two-time parameter) empirical process of the squared residuals

$$\hat{e}_n(t, s) = n^{1/2} s\big(\widehat{F}_n(t, s) - F(t)\big),$$

where

$$\widehat{F}_n(t, s) = \frac{1}{ns} \sum_{p<k\le ns} \mathbf{I}\{\hat{\varepsilon}_k^2 \le t\}, \qquad p/n < s \le 1$$

and $\widehat{F}_n(t, s) = 0$ if $0 \le s \le p/n$. We note that $\hat{e}_n(t, 1)$ is the empirical process of the $\hat{\varepsilon}_k^2$, $p < k \le n$.

Following Giraitis, Kokoszka and Leipus (2000) we also assume that

$$E\varepsilon_0^4 < \infty \tag{1.7}$$

and

$$\big(E\varepsilon_0^4\big)^{1/2} \sum_{1\le j\le p} b_j < 1. \tag{1.8}$$

If condition (1.8) is satisfied, then the ARCH equations above have a unique strictly stationary solution such that $Ey_k^4 < \infty$ and the squares y_k^2 have a Volterra representation

$$y_k^2 = \sum_{l=0}^{\infty} \sum_{j_1,\dots,j_l=1}^{p} b_{j_1} \cdots b_{j_l} \varepsilon_k^2 \varepsilon_{k-j_1}^2 \cdots \varepsilon_{k-j_1-\cdots-j_l}^2.$$

Thus y_k^2 is a function of $\varepsilon_k, \varepsilon_{k-1}, \dots$, and so it follows from the standard theory [see, e.g., Stout (1974), pages 181 and 182] that the sequence $\{y_k^2\}$ is ergodic.

We note that condition (1.8) is not necessary for the covariance stationarity of the process $\{y_k^2\}$, but it is easy to verify. Necessary and sufficient conditions are more complex and difficult to state in a closed form for $p > 2$ see Section 3.4 of Teräsvirta (1999). It is also well known that ARCH(p) and more general sequences from the GARCH family are not only ergodic but also mixing with geometric rate; we refer to Lu and Cheng [(1997), Remark 4.2] for further references. The results of this paper remain valid if condition (1.8) is replaced by any assumption guaranteeing that the process $\{y_k\}$ is strictly stationary with $Ey_k^4 < \infty$ and ergodic. The theory developed here may not be valid if the assuption $Ey_k^4 < \infty$ is dropped as the results of Davis and Mikosch (1998) suggest. These authors consider, however, functions of the observations y_k rather than estimated residuals.

In order to state our main result we need further notation:

$$\alpha_i = Ef_i(\varepsilon_{-1}, \varepsilon_{-2}, \dots), \qquad 0 \le i \le p,$$

$$g_i(t) = \alpha_i \int_0^t l_i(u) f(u)\, du, \qquad 0 \le i \le p,$$

$$\beta_0 = E\left[\frac{1}{\sigma_0^2}\right], \qquad \beta_i = E\left[\frac{y_{-i}^2}{\sigma_0^2}\right], 1 \le i \le p,$$

$$\gamma_{ij} = E\big[l_i(\varepsilon_0^2) l_j(\varepsilon_0^2)\big] E[f_i(\varepsilon_{-1}, \varepsilon_{-2}, \dots) f_j(\varepsilon_{-1}, \varepsilon_{-2}, \dots)], \qquad 0 \le i, j \le p,$$

$$\begin{aligned} r(t, t', s, s') = {} & (s \wedge s')(F(t \wedge t') - F(t)F(t')) \\ & + tf(t)ss' \sum_{0 \le i \le p} \beta_i g_i(t') + t' f(t') ss' \sum_{0 \le i \le p} \beta_i g_i(t) \\ & + tf(t) t' f(t') ss' \sum_{0 \le i, j \le p} \beta_i \gamma_{ij} \beta_j, \end{aligned}$$

where $s \wedge t = \min(s, t)$.

THEOREM 1.1. *If conditions* (1.1)–(1.8) *hold, then*

$$\hat{e}_n(t, s) \to \Gamma(t, s),$$

where the convergence is in the Skorokhod space $\mathscr{D}([0, \infty] \times [0, 1])$ *and* Γ *is a Gaussian process with*

$$E\Gamma(t, s) = 0 \quad \text{and} \quad E[\Gamma(t, s)\Gamma(t', s')] = r(t, t', s, s').$$

The convergence in Theorem 1.1 is equivalent to the convergence $\hat{e}_n \times (F^{-1}(x), s) \to \Gamma(F^{-1}(x), s)$, $0 \le x, s \le 1$, in $\mathscr{D}([0,1]\times[0,1])$. [$F^{-1}(0) \ge -\infty$ and $F^{-1}(1) \le \infty$ denote the endpoints of the support of F.]

Before presenting detailed proofs, we outline the argument. Set

$$\delta_i^2 = \frac{\hat{\sigma}_i^2}{\sigma_i^2}$$

and observe that

$$\sup_{0\le t<\infty}\sup_{0\le s\le 1}\left|\hat{e}_n(t,s) - \left(\hat{e}_{n,1}(t,s) + \hat{e}_{n,2}(t,s)\right)\right| = O_P\left(n^{-1/2}\right),$$

where

$$\hat{e}_{n,1}(t,s) = n^{-1/2}\sum_{p<i\le ns}\left(\mathbf{I}\left\{\varepsilon_i^2 \le t\delta_i^2\right\} - F\left(t\delta_i^2\right)\right)$$

and

$$\hat{e}_{n,2}(t,s) = n^{-1/2}\sum_{p<i\le ns}\left(F\left(t\delta_i^2\right) - F(t)\right).$$

We will show in the proofs that δ_i^2 is so close to 1 that the difference between $\hat{e}_{n,1}(t,s)$ and

$$e_n(t,s) = n^{-1/2}\sum_{p<i\le ns}\left(\mathbf{I}\left\{\varepsilon_i^2 \le t\right\} - F(t)\right) \tag{1.9}$$

is negligible. As for the second term, $\hat{e}_{n,2}(t,s)$ will be approximated by

$$h_n(t,s) = tf(t)n^{-1/2}\sum_{p<i\le ns}\left(\delta_i^2 - 1\right). \tag{1.10}$$

Observe that

$$\begin{aligned}
\sum_{p<i\le ns}\left(\delta_i^2 - 1\right) &= \sum_{p<i\le ns}\frac{\hat{\sigma}_i^2 - \sigma_i^2}{\sigma_i^2} \\
&= (\hat{b}_0 - b_0)\sum_{p<i\le ns}\frac{1}{\sigma_i^2} + (\hat{b}_1 - b_1)\sum_{p<i\le ns}\frac{y_{i-1}^2}{\sigma_i^2} \\
&\quad + \cdots + (\hat{b}_p - b_p)\sum_{p<i\le ns}\frac{y_{i-p}^2}{\sigma_i^2} \\
&= ns\big\{(\hat{b}_0 - b_0)\beta_0 + (\hat{b}_1 - b_1)\beta_1 \\
&\qquad + \cdots + (\hat{b}_p - b_p)\beta_p + o_P(n^{-1/2})\big\}.
\end{aligned}$$

Hence

$$\sup_{0\le t<\infty}\sup_{0\le s\le 1}\left|\hat{e}_n(t,s) - \left(e_n(t,s) + tf(t)s\sum_{0\le i\le p} n^{1/2}(\hat{b}_i - b_i)\beta_i\right)\right| = o_P(1) \tag{1.11}$$

and therefore the joint convergence of $e_n(t, s)$ and $\sqrt{n}(\hat{\mathbf{b}}_n - \mathbf{b})$ will imply the result in Theorem (1.1).

Relation (1.11) explains the structure of $\hat{e}_n(t, s)$ and also the formula for the covariance function of the limiting Gaussian process. It is interesting to compare (1.11) with Theorem 1 of Bai (1994) which shows that for the residuals in an ARMA model $\hat{e}_n(t, s) - e_n(t, s)$ is uniformly $o_P(1)$. Boldin (1998) showed that an additional term, analogous to $tf(t)s\sum_{0\le i\le p} n^{1/2}(\hat{b}_i - b_i)\beta_i$, makes an asymptotic contribution to the empirical process of residuals (not squared residuals) of an ARCH(1) model.

The covariance function $r(\cdot)$ depends on several unknown parameters and functions. In Section 2 we obtain explicit formulas for l_i, f_i, $0 \le i \le p$, in the case of most commonly used estimators. Observe that α_i, $g_i(t)$ and β_i are expected values and can be consistently estimated by the corresponding averages. Note also that $\{\gamma_{ij},\ 0 \le i, j \le p\}$ is the asymptotic covariance matrix of $n^{1/2}(\hat{\mathbf{b}}_n - \mathbf{b})$ and its estimation is discussed in Gouriéroux (1997). In some applications (see Section 3), it is not necessary to estimate the parameters in $r(\cdot)$.

We will show in the proof of Theorem (1.1) that

$$\Big\{e_n(t, s),\ n^{1/2} \sum_{0\le i\le p} (\hat{b}_i - b_i)\beta_i,\ 0 \le t < \infty, 0 \le s \le 1\Big\}$$

converges weakly to a Gaussian vector valued process $\{K(F(t), s), \xi, 0 \le t < \infty, 0 \le s \le 1\}$ with

$$E\big[K(F(t), s)\big] = 0, \qquad E\xi = 0, \tag{1.12}$$

$$E\big[K(F(t), s)K(F(t'), s')\big] = (s \wedge s')(F(t \wedge t') - F(t)F(t')), \tag{1.13}$$

$$E\xi^2 = \sum_{0\le i,\ j\le p} \beta_i \gamma_{ij} \beta_j, \tag{1.14}$$

$$E\big[K(F(t), s)\xi\big] = s \sum_{0\le i\le p} \beta_i g_i(t). \tag{1.15}$$

Observe that $\{K(x, s), 0 \le x, s \le 1\}$ is a Kiefer process. Also, the limit process $\Gamma(t, s)$ in Theorem 1.1 can be written as

$$\begin{aligned}\big\{\Gamma(t, s),\ 0 \le t < \infty, 0 \le s \le 1\big\} \stackrel{d}{=} \big\{&K(F(t), s) + tf(t)s\xi,\\ &0 \le t < \infty, 0 \le s \le 1\big\}.\end{aligned} \tag{1.16}$$

In particular, the process $\Gamma(\cdot, 1)$ admits the representation

$$\big\{\Gamma(t, 1), 0 \le t < \infty\big\} \stackrel{d}{=} \big\{B(F(t)) + tf(t)\xi, 0 \le t < \infty\big\}, \tag{1.17}$$

where $\{B(x), 0 \le x \le 1\}$ is a Brownian bridge.

The results in (1.11), (1.16) and (1.17) are similar to main theorems on parameter estimated processes. Durbin (1973a, b) was the first who considered the weak convergence of the empirical process when parameters are estimated. He mainly studied the case when the parameters are estimated by the maximum likelihood method. Burke, Csörgő, Csörgő and Révész (1979) and Csörgő and Révész (1981) considered the general case when it is assumed only that the difference between the estimator and the estimated parameter is approximately given by an integral with respect to the empirical process of the observations. The limit in their case has a representation like (1.17), but in the iid case, ξ is a stochastic integral of a deterministic function with respect to $B(F(\cdot))$. Our case is somewhat different.

In the next section we consider three examples when condition (1.3) is satisfied. Section 3 discusses some applications of Theorem 1.1, whose proof is postponed until Section 4.

2. Asymptotic linearity of estimators. In this section we consider several examples of estimators satisfying (1.3). We would like to point out that, in general, asymptotic linearity like (1.3) is usually not difficult to establish whenever asymptotic normality holds.

In the remainder of this section we use the notation $\mathbf{s} = (s_0, s_1, \dots, s_p) \in [0, \infty)^{p+1}$ and

$$\sigma_i^2(\mathbf{s}) = s_0 + s_1 y_{i-1}^2 + \cdots + s_p y_{i-p}^2.$$

2.1. *Pseudomaximum likelihood estimation.* Let

$$\mathscr{L}_n(\mathbf{s}) = -\frac{1}{2} \sum_{1 \le i \le n} \log \sigma_i^2(\mathbf{s}) - \frac{1}{2} \sum_{1 \le i \le n} \frac{y_i^2}{\sigma_i^2(\mathbf{s})}$$

denote the log of the pseudolikelihood function. The estimator is the solution of the equation $\mathscr{L}_n'(\hat{\mathbf{b}}_n) = \mathbf{0}$, where

$$\mathscr{L}_n'(\mathbf{s}) = \frac{1}{2} \sum_{1 \le i \le n} \left(\frac{y_i^2}{\sigma_i^2(\mathbf{s})} - 1 \right) \frac{1}{\sigma_i^2(\mathbf{s})} \frac{\partial \sigma_i^2(\mathbf{s})}{\partial(\mathbf{s})}.$$

As pointed out in Section 4.1 of Gouriéroux (1997), $\hat{\mathbf{b}}_n$ is asymptotically normal under standard regularity conditions even if the ε_i are not standard normal; that is, conditionally on the past, the observations are not necessarily normal. This holds true for time series models much more general than the ARCH(p) considered here. In the following we assume only that

$$\|\hat{\mathbf{b}}_n - \mathbf{b}\| = O_P(n^{-1/2}), \tag{2.1}$$

where $\mathbf{b}$ is the true value of the parameter vector. The second derivative of $\mathscr{L}_n(\mathbf{s})$ is the matrix $\mathscr{L}_n''(\mathbf{s})$. By the ergodic theorem,

$$\left| \frac{1}{n} \mathscr{L}_n''(\mathbf{b}) - \mathscr{J}(\mathbf{b}) \right| \xrightarrow{P} 0, \tag{2.2}$$

where $\mathscr{J}(\mathbf{b})$ is a deterministic matrix. Since $\mathscr{L}_n'(\hat{\mathbf{b}}_n) - \mathscr{L}_n'(\mathbf{b}) = -\mathscr{L}_n'(\mathbf{b})$, by the mean value theorem, we have

$$\mathscr{L}_n''(\beta)(\hat{\mathbf{b}}_n - \mathbf{b}) = -\mathscr{L}_n'(\mathbf{b}), \tag{2.3}$$

where β is a point between $\hat{\mathbf{b}}_n$ and $\mathbf{b}$. The matrix $\mathscr{J}(\mathbf{b})$ is invertible [see, e.g., Gouriéroux (1997) pages 50 and 51], so (2.3) yields

$$\hat{\mathbf{b}}_n - \mathbf{b} = -\frac{1}{n}\mathscr{J}^{-1}(\mathbf{b})\mathscr{L}_n'(\mathbf{b}) - \mathscr{J}^{-1}(\mathbf{b})\left(\frac{1}{n}\mathscr{L}_n''(\beta) - \mathscr{J}(\mathbf{b})\right)(\hat{\mathbf{b}}_n - \mathbf{b}). \tag{2.4}$$

To see that the second term on the right-hand side of (2.4) is $o_P(n^{-1/2})$, use (2.1) and the decomposition

$$\frac{1}{n}\mathscr{L}_n''(\beta) - \mathscr{J}(\mathbf{b}) = \frac{1}{n}\big(\mathscr{L}_n''(\beta) - \mathscr{L}_n''(\mathbf{b})\big) + \left(\frac{1}{n}\mathscr{L}_n''(\mathbf{b}) - \mathscr{J}(\mathbf{b})\right). \tag{2.5}$$

The first term on the right-hand side of (2.5) is $o_P(1/n)$ because $\beta \to^P \mathbf{b}$, whereas the second is $o_P(1)$ by (2.2). Representation (1.3) now follows since

$$\mathscr{L}_n'(\mathbf{b}) = \frac{1}{2}\sum_{1\le i\le n}\big(\varepsilon_i^2 - 1\big)\frac{1}{\sigma_i^2(\mathbf{b})}\left[\frac{\partial\sigma_i^2(\mathbf{s})}{\partial\mathbf{s}}\right]_{\mathbf{s}=\mathbf{b}}.$$

2.2. *Conditional least squares.* By definition $E(y_k^2|\mathscr{F}_{k-1}) = \sigma_i^2(\mathbf{s})$, so the conditional sum of squares is

$$Q_n(\mathbf{s}) = \sum_{p<i\le n}\big(y_i^2 - \sigma_i^2(\mathbf{s})\big)^2,$$

The conditional least squares estimators of the b_k, $k = 0, 1, \dots, p$, are the solutions of the equations

$$\frac{\partial Q_n}{\partial s_k} = -2\sum_{p<i\le n}\big(y_i^2 - \sigma_i^2(\mathbf{s})\big)\frac{\partial\sigma_i^2(\mathbf{s})}{\partial s_k} = 0, \qquad k = 0, 1, \dots, p.$$

In order to establish (1.3) we proceed similarly as in Section 2.1. Direct verification shows that $Q_n''(\mathbf{s})(= Q_n''(\mathbf{b}))$ does not depend on $\mathbf{s}$ and by the ergodic theorem.

$$n^{-1}Q_n''(\mathbf{b}) \xrightarrow{P} \Lambda(\mathbf{b}). \tag{2.6}$$

Since $Q_n'(\hat{\mathbf{b}}_n) = \mathbf{0}$,

$$-Q_n'(\mathbf{b}) = Q_n''(\mathbf{b})(\hat{\mathbf{b}}_n - \mathbf{b}). \tag{2.7}$$

Relations (2.7) and (2.6) in conjunction with a central limit theorem for the squares of an ARCH process imply that $n^{1/2}(\hat{\mathbf{b}}_n - \mathbf{b})$ is asymptotically normal. Consequently, arguing as in Section 2.1,

$$\hat{\mathbf{b}}_n - \mathbf{b} = -\frac{1}{n}\Lambda^{-1}Q_n'(\mathbf{b}) + o_P(n^{-1/2}).$$

Representation (1.3) now follows on observing that

$$\frac{\partial Q_n}{\partial s_0}\Big|_{\mathbf{s}=\mathbf{b}} = -2\sum_{p<i\le n}\left(\varepsilon_i^2-1\right)\sigma_i^2(\mathbf{b})$$

and for $1 \le k \le p$,

$$\frac{\partial Q_n}{\partial s_k}\Big|_{\mathbf{s}=\mathbf{b}} = -2\sum_{p<i\le n}\left(\varepsilon_i^2-1\right)\sigma_i^2(\mathbf{b})y_{i-k}^2.$$

2.3. *Conditional likelihood.* Let h denote the density function of ε_0. The conditional density of y_i given $\mathscr{F}_{i-1}$ is $h(t/\sigma_i)/\sigma_i$. Hence the conditional log likelihood function is

$$\mathscr{L}_n^*(\mathbf{s}) = -\frac{1}{2}\sum_{1\le i\le n}\log\sigma_i^2(\mathbf{s}) + \sum_{1\le i\le n}\log h(y_i/\sigma_i(\mathbf{s})).$$

If ε_0 is normal, then $\mathscr{L}_n^*$ coincides with $\mathscr{L}_n$ in Section (2.1). The conditional likelihood estimator $\hat{\mathbf{b}}_n^*$ is the solution to $\mathscr{L}_n^{*\prime}(\hat{\mathbf{b}}_n^*) = \mathbf{0}$, where

$$\mathscr{L}_n^{*\prime}(\mathbf{s}) = -\frac{1}{2}\sum_{1\le i\le n}\left\{\frac{y_i}{\sigma_i(\mathbf{s})}\frac{h'(y_i/\sigma_i(\mathbf{s}))}{h(y_i/\sigma_i(\mathbf{s}))}+1\right\}\frac{1}{\sigma_i^2(\mathbf{s})}\frac{\partial\sigma_i^2(\mathbf{s})}{\partial\mathbf{s}}.$$

Assuming that $n^{-1}E\mathscr{L}_n^{*\prime\prime}(\mathbf{s}) \to \mathscr{L}^{*\prime\prime}(\mathbf{s})$ and $\mathscr{L}^{*\prime\prime}(\mathbf{b})$ is a positive definite matrix we obtain that $\mathscr{L}^*(\mathbf{b}) > \mathscr{L}^*(\mathbf{s})$ in a neighborhood of $\mathbf{b}$ where $\mathscr{L}^*(\mathbf{s}) = \lim_{n\to\infty}\mathscr{L}_n^*(\mathbf{s})/n$. This implies that

$$\left\|\hat{\mathbf{b}}_n^* - \mathbf{b}\right\| = o_P(1) \tag{2.8}$$

[cf. Amemiya (1985)]. Standard arguments show that (2.8) implies that (2.1) also holds for $\hat{\mathbf{b}}_n^*$ and therefore the arguments used in Section 2.1 give

$$\hat{\mathbf{b}}_n^* - \mathbf{b} = -\left[\mathscr{L}^{*\prime\prime}(\mathbf{b})\right]^{-1}\frac{1}{n}\mathscr{L}_n^{*\prime}(\mathbf{b}) + o_P\left(n^{-1/2}\right).$$

Since

$$\mathscr{L}_n^{*\prime}(\mathbf{b}) = -\frac{1}{2}\sum_{1\le i\le n}\left\{\varepsilon_i\frac{h'(\varepsilon_i)}{h(\varepsilon_i)}+1\right\}\frac{1}{\sigma_i^2(\mathbf{b})}\left(1, y_{i-1}^2, \ldots, y_{i-p}^2\right),$$

we showed that (1.3) holds.

3. Some applications. Similarly to the parameter estimated process, the limit of $\hat{e}_n(t,s)$ depends on the unknown parameter $\mathbf{b}$ and the correlation between $\mathbf{I}\{\varepsilon_i^2 \le t\}$ and $\mathbf{b}$. The martingale approach of Khmaladze (1981) was used to transform the parameter estimated process so that the limit is a Brownian motion and therefore distribution free. The transformation is based on a martingale representation of the empirical process and in case of parameter estimation the transformation must be constructed from the data. Khamaladze's martingale approach was extended by Koul and Stute (1999) to regression and autoregression. These authors use an m-dependent version of (1.3), corresponding to $f_i(\cdot) \equiv 1$, $l_i = l_i(\varepsilon_j^2, \dots, \varepsilon_{j-m}^2)$, $1 \le i \le p$. It is not immediately clear how the martingale approach can be adapted to ARCH sequences. Csörgő (1983) and Horváth (1985) suggested kernel-transforms of the parameter estimated empirical process to get a distribution (parameter) free limit. Their approach uses the covariance function (or its estimator), so, in principle, it might be used to transform $\hat{e}_n(t, s)$. However, the required calculation of the eigenfunctions and eigenvalues of the estimated covariance function is computationally unreasonable.

Next we consider some simple and immediate consequences of Theorem 1.1.

EXAMPLE 3.1. The detection of possible changes in the structure of observation has been extensively studied in various settings [see, e.g, Csörgő and Horváth (1997)]. Chu (1995) and Mikosch and Stărică (1999) investigated changes in GARCH models. Theorem 1.1 can be used to test for a possible change in the distribution of the innovations in an ARCH model. Following the method described in detail in Section 2.6 of Csörgő and Horváth (1997), the process

$$w_n(t, s) = \begin{cases} 0, & 0 \le s \le p/n, \\ \dfrac{[ns](n - [ns])}{n^{3/2}} \Big(\hat{F}_n(t, s) - \hat{F}_n^*(t, s) \Big), & p/n < s \le (n-1)/n, \\ 0, & (n-1)/n < s \le 1, \end{cases}$$

with

$$\hat{F}_n^*(t, s) = \frac{1}{n - ns} \sum_{ns < i \le n} \mathbf{I}\{\hat{\varepsilon}_i^2 \le t\}$$

compares the empirical distribution function of $\hat{\varepsilon}_{p+1}^2, \dots, \hat{\varepsilon}_{[ns]}^2$ to that of $\hat{\varepsilon}_{[ns]+1}^2, \dots, \hat{\varepsilon}_n^2$. Observe that $w_n(t, s)$ has the same limit in $\mathscr{D}([0, \infty] \times [0, 1])$ as $\hat{e}_n(t, s) - s\hat{e}_n(t, 1)$ which by (1.11) has the same limit as $e_n(t, s) - se_n(t, 1)$ [the contributions from the function h_n defined in (1.10) cancel]. Thus Theorem 1.1 implies that $w_n(t, s)$ converges in $\mathscr{D}([0, \infty] \times [0, 1])$ to $K^*(F(t), s)$, where $K^*(u, s)$ is a tied-down Kiefer process, that is, a Gaussian process with $E[K^*(u, s)] = 0$ and

$$E[K^*(u, s)K^*(u', s')] = (u \wedge u' - uu')(s \wedge s' - ss').$$

Since F is continuous,

$$\int_0^1 \int_0^\infty w_n^2(t, s)\, d\hat{F}_n(t, 1)\, ds \xrightarrow{d} \int_0^1 \int_0^1 [K^*(u, s)]^2\, du ds$$

and

$$\sup_{0 \le t < \infty,\, 0 \le s \le 1} \Big| w_n(t, s) \Big| \xrightarrow{d} \sup_{0 \le u,\, s \le 1} \Big| K^*(u, s) \Big|.$$

Blum, Kiefer and Rossenblatt (1961), Cotteril and Csörgő (1985) and Martynov (1992) studied the distribution function of $\int_0^1 \int_0^1 [K^*(t, s)]^2\, dt\, ds$. They computed its characteristic function and inverting the characteristic function one can compute the distribution function at points of interest. Tables for the distribution function of $\int_0^1 \int_0^1 [K^*(t,s)]^2 \, dt ds$ can be found in Blum, Kiefer, and Rossenblatt (1961) and Cotteril and Csörgŏ (1985). The distribution of $\sup_{0 \le u, s \le 1} |K^*(u,s)|$ is tabulated in Picard (1985).

EXAMPLE 3.2. In this example we consider sums of functions of the squared residuals. Let $\psi(\cdot)$ be a function on $[0,\infty)$ with finite variation on $[0,\infty)$. Theorem (1.1) and integration by parts give

$$\int_0^\infty \psi(t) \, d\hat{e}_n(t,1) \xrightarrow{d} \int_0^\infty \psi(t) \, d\Gamma(t,1).$$

If

$$\int_0^\infty \psi(t) d(tf(t)) = 0, \tag{3.1}$$

then by (1.17) we have

$$\frac{n^{1/2}(\frac{1}{n}\sum_{p<i\le n} \psi(\hat{\varepsilon}_i^2) - \int_0^\infty \psi(t) \, dF(t))}{(\operatorname{Var} \psi(\varepsilon_0^2))^{1/2}} \xrightarrow{d} N(0,1), \tag{3.2}$$

where $N(0,1)$ stands for the standard normal random variable. We note that neither condition (3.1) nor the norming constants in (3.2) depend on the unknown $\mathbf{b}$.

EXAMPLE 3.3. In this example a very simple χ^2-test is suggested. It can be used to check if the distribution function of $\hat{\varepsilon}_1^2, \dots, \hat{\varepsilon}_{[ns]}^2$ is F, for any $0 < s \le 1$. Let $M \ge 1$ be an integer and $0 < s < 1$. Define $0 \le t_0 < t_1 < t_2 < \cdots < t_M \le \infty$. By Theorem 1.1, the vector $\boldsymbol{Z}_M = (Z_{1,M}, \dots, Z_{M,M})$,

$$\begin{aligned} Z_{i,M} = n^{1/2} \Bigg[& \frac{1}{ns} \sum_{p \le j \le ns} \Big(\mathbf{I}\Big\{ t_{i-1} \le \hat{\varepsilon}_j^2 < t_i \Big\} - (F(t_i) - F(t_{i-1})) \Big) \\ & - \frac{1}{n(1-s)} \sum_{ns < j \le n} \Big(\mathbf{I}\{ t_{i-1} \le \hat{\varepsilon}_j^2 < t_i \} - (F(t_i) - F(t_{i-1})) \Big) \Bigg], \end{aligned}$$

is asymptotically normal with zero mean and covariance matrix $[s(1-s)]^{-1}\boldsymbol{D}$, $\boldsymbol{D} = \{d_{ij}, \ 1 \le i, j \le M\}$, where

$$d_{ij} = \begin{cases} (F(t_i) - F(t_{i-1})) - (F(t_i) - F(t_{i-1}))^2, & i = j, \\ -(F(t_i) - F(t_{i-1}))(F(t_j) - F(t_{j-1})), & i \ne j. \end{cases}$$

It is well-known [just observe that $\boldsymbol{D}$ is the covariance matrix of $e_n(t_{i-1},1) - e_n(t_i,1)$, $1 \le i \le M$] that the rank of $\boldsymbol{D}$ is $M-1$. If $\boldsymbol{D}^{-1}$ denotes a generalized inverse of $\boldsymbol{D}$ [cf. Seber (1977), page 76], then $s(1-s)\boldsymbol{Z}_M \boldsymbol{D}^{-1} \boldsymbol{Z}_M^T$ is asymptotically $\chi^2(M-1)$ [cf. Theorem 2.8 in Seber (1977)].

Since the asymptotic distribution of $\hat{e}_n(t,1)$ depends on unknown parameters, classical goodness-of-fit tests, like the Kolmogorov–Smirnov and Cramér–von Mises tests, are not directly applicable. However, Theorem 1.1 shows that appropriate functionals of $\hat{e}_n(t, 1)$ have an asymptotic distribution, and it may be hoped that bootstrap goodness-of-fit tests may be developed. In Horváth, Kokoszka and Teyssiére (2000) we proposed and examined by means of a simulation study such tests for ARCH(p) and more general GARCH models. Bootstrap tests based on $\hat{e}_n(t, 1)$ can, in principle, be expected to detect any departure from the null hypothesis of independent, identically distributed squared innovations ε_i^2 with specified distribution function F. In practice, however, certain alternatives, like, for example, the change-point alternative discussed in Example (3.1), are not reliably detected and it may be expected that the asymptotic test described in Example 3.1 will have higher power. Other alternatives, like, for example, distribution function F even slightly different from the postulated one, are detected with probability around 0.8 even for series of length 200 (financial time series based on intradaily trading have often lengths of several thousand). The tests have close to perfect size even for series of length 50. We refer to Horváth, Kokoszka and Teyssiére (2000) for further details.

It appears often to be the case that for GARCH and related models many important statistics are not asymptotically pivotal and alternative approaches like bootstrap and response surface analysis must be used in such situations [see Frances and Vandijk (2000)] for an analysis of a related problem of outlier detection.

4. Proof of Theorem 1.1. Set

$$\zeta_n = n^{-1/2} \sum_{0\le k\le p} \beta_k \sum_{1\le j\le n} l_k(\varepsilon_j^2) f_k(\varepsilon_{j-1}, \varepsilon_{j-2}, \dots) \tag{4.1}$$

and

$$u_n(t, s) = tf(t)s\zeta_n. \tag{4.2}$$

In light of (1.16), Theorem 1.1 follows from Lemmas 4.6 and 4.7 below. The proof of Lemma 4.7 is fairly standard. In order to establish Lemma 4.6 we need a number of auxiliary lemmas.

In the following, $\|\cdot\|$ stands for the maximum norm of a vector.

LEMMA 4.1. *If conditions* (1.3)–(1.8) *hold*, *then*

$$\sqrt{n}\|\hat{\mathbf{b}}_n - \mathbf{b}\| = O_P(1). \tag{4.3}$$

PROOF. Define $X_{j,i} = l_i(\varepsilon_j^2) f_i(\varepsilon_{j-1}, \varepsilon_{j-2}, \dots)$ and observe that for any $0 \le i \le p$, the variables $X_{j,i}$ are zero mean, uncorrelated and form a stationary sequence. Therefore $\mathrm{Var}[\sum_{1\le j\le n} X_{j,i}] = O(n)$ and so (4.3) follows from (1.3) and the Chebyshev inequality. □

Recall now the definition of the function $h_n(t, s)$ given in (1.10).

LEMMA 4.2. *If* (1.1)–(1.8) *hold, then*

$$\sup_{0\le t<\infty}\sup_{0\le s\le 1}|\hat{e}_{n,2}(t,s)-h_n(t,s)|=o_P(1).$$

PROOF. By the mean value theorem,

$$\begin{aligned} F(t\delta_i^2)-F(t)&=f(\vartheta_i)t(\delta_i^2-1)\\ &=tf(t)(\delta_i^2-1)+t(f(\vartheta_i)-f(t))(\delta_i^2-1), \end{aligned} \tag{4.4}$$

where ϑ_i is between t and $t\delta_i^2$. Hence

$$\begin{aligned} \hat{e}_{n,2}(t,s)-h_n(t,s)&=n^{-1/2}\sum_{p<i\le ns}t(f(\vartheta_i)-f(t))(\delta_i^2-1)\\ &=n^{-1/2}(\hat{b}_0-b_0)\sum_{p<i\le ns}\frac{1}{\sigma_i^2}t(f(t)-f(\vartheta_i))\\ &\quad+n^{-1/2}\sum_{1\le k\le p}(\hat{b}_k-b_k)\sum_{p<i\le ns}\frac{y_{i-k}^2}{\sigma_i^2}t(f(t)-f(\vartheta_i)). \end{aligned}$$

We will verify below that

$$\max_{p<i\le n}\sup_{0\le t<\infty}t|f(t)-f(\vartheta_i)|=o_P(1), \tag{4.5}$$

which together with Lemma 4.1 will imply that

$$\sup_{0\le s\le 1}\sup_{0\le t<\infty}|\hat{e}_{n,2}(t,s)-h_n(t,s)|=o_P(1)\frac{1}{n}\sum_{p<i\le n}\frac{1}{\sigma_i^2}\left[1+\sum_{1\le k\le p}y_{i-k}^2\right],$$

and so the claim will follow from the ergodic theorem.

In order to verify (4.5), we first show that

$$\max_{p<i\le n}\sup_{0\le t<\infty}\frac{|\vartheta_i-t|}{t}=o_P(1). \tag{4.6}$$

Observe that

$$|\vartheta_i-t|\le\frac{t}{b_0}\left\{|\hat{b}_0-b_0|+|\hat{b}_1-b_1|y_{i-1}^2+\cdots+|\hat{b}_p-b_p|y_{i-p}^2\right\}. \tag{4.7}$$

Since $\{y_i^2\}$ is a stationary sequence with $Ey_i^4<\infty$,

$$\max_{1\le i\le n}y_i^2=o_P(n^{1/2}). \tag{4.8}$$

Putting together (4.7), Lemma 4.1 and (4.8) we obtain (4.6).

Now to verify (4.5) fix $\varepsilon>0$ and $0<T<\infty$ so that $\sup_{T/2\le t<\infty}tf(t)\le\varepsilon$ [cf. (1.2)]. By continuity of f it is not difficult to see that (4.6) together with (1.1) and the first relation in (1.2) yield

$$\max_{p<i\le n}\sup_{0\le t\le T}t|f(t)-f(\vartheta_i)|=o_P(1).$$

Let

$$\widetilde{\Omega}_n = \left\{\omega : \frac{\vartheta_i}{t} \geq \frac{1}{2} \text{ for all } 0 \leq t < \infty,\ p < i \leq n\right\}.$$

By (4.6), we have $\lim_{n\to\infty} P\{\widetilde{\Omega}_n\} = 1$. On the event $\widetilde{\Omega}_n$ we have

$$\begin{aligned}\max_{p<i\leq n} \sup_{T\leq t<\infty} t|f(t) - f(\vartheta_i)| &\leq \varepsilon + \max_{p<i\leq n} \sup_{T\leq t<\infty} \frac{t}{\vartheta_i} \max_{p<i\leq n} \sup_{T\leq t<\infty} \vartheta_i f(\vartheta_i)\\ &\leq \varepsilon + 2 \sup_{T/2\leq t<\infty} tf(t) \leq 3\varepsilon,\end{aligned}$$

and so (4.5) is proved. □

For any $\mathbf{a} = (a_0, a_1, \dots, a_p) \in R^{p+1}$ define

$$\gamma_i\mathbf{a} = a_0\frac{1}{\sigma_i^2} + a_1\frac{y_{i-1}^2}{\sigma_i^2} + \cdots + a_p\frac{y_{i-p}^2}{\sigma_i^2},$$

so that $\delta_i^2 = 1 + \gamma_i(\hat{\mathbf{b}}_n - \mathbf{b})$. Let

$$e_n^*(t, s, \mathbf{a}) = n^{-1/2} \sum_{p<i\leq ns} \Big(\mathbf{I}\big\{\varepsilon_i^2 \leq t + tn^{-1/2}\gamma_i(\mathbf{a})\big\} - F(t + tn^{-1/2}\gamma_i(\mathbf{a}))\Big).$$

Observe that

$$e_n^*(t, s, n^{1/2}(\hat{\mathbf{b}}_n - \mathbf{b})) = \hat{e}_{n,1}(t, s). \tag{4.9}$$

LEMMA 4.3. *If conditions* (1.1)–(1.8) *hold, then for any* $x > 0$,

$$P\Big\{\sup_{0\leq s\leq 1} |e_n^*(t, s, \mathbf{a}) - e_n(t, s)| \geq x\Big\} \leq \frac{1}{n}\frac{C}{x^4}(\|\mathbf{a}\|^2 + 1),$$

if $n \geq 16\|\mathbf{a}\|^2/b_0^2$, *with some constant* C (*which depends on* f, p, b_0, Ey_0^4 *but not on* t).

PROOF. Let

$$\xi_i(t, \mathbf{a}) = \mathbf{I}\{\varepsilon_i^2 \leq t + tn^{-1/2}\gamma_i(\mathbf{a})\} - \mathbf{I}\{\varepsilon_i^2 \leq t\} - (F(t + tn^{-1/2}\gamma_i(\mathbf{a})) - F(t)),$$

so that

$$e_n^*(t, s, \mathbf{a}) - e_n(t, s) = n^{-1/2} \sum_{p<i\leq ns} \xi_i(t, \mathbf{a}).$$

Since t and $\mathbf{a}$ are fixed, we write in the following ξ_i and γ_i instead of $\xi_i(t, \mathbf{a})$ and $\gamma_i\mathbf{a}$. The σ-algebra generated by $\{\varepsilon_j,\ -\infty < j \leq i\}$ will be denoted by $\mathscr{F}_i$. It is not difficult to verify that $E\{\xi_i^2|\mathscr{F}_{i-1}\} \leq |F(t + tn^{-1/2}\gamma_i) - F(t)\Big|$ and so we have

$$\sum_{1\leq i\leq n} E\big\{\xi_i^2|\mathscr{F}_{i-1}\big\} \leq \sum_{1\leq i\leq n} |F(t + tn^{-1/2}\gamma_i) - F(t)|. \tag{4.10}$$

There is ξ_i^* between t and $t + tn^{-1/2}\gamma_i$ such that $F(t + tn^{-1/2}\gamma_i) - F(t) = f(\xi_i^*)tn^{-1/2}\gamma_i$ and so, by conditions (1.1) and (1.2),

$$|F(t+tn^{-1/2}\gamma_i) - F(t)\Big| = \xi_i^* f(\xi_i^*)(t/\xi_i^*)n^{-1/2}|\gamma_i| \le c_1(t/\xi_i^*)n^{-1/2}|\gamma_i|,$$

where $c_1 = \sup_{0\le u\le\infty} uf(u)$. If $n^{-1/2}\gamma_i \ge -1/2$, then $t/2 \le \min(t, t + n^{-1/2}\gamma_i)$, so $t/\xi_i^* \le 2$ and therefore we have

$$|F(t+tn^{-1/2}\gamma_i) - F(t)| \le 2c_1 n^{-1/2}|\gamma_i|\mathbf{I}\{n^{-1/2}\gamma_i \ge -1/2\} + \mathbf{I}\{n^{-1/2}\gamma_i < -1/2\}. \tag{4.11}$$

By (4.10) and (4.11),

$$\begin{aligned}\Bigg[\sum_{1\le i\le n} E\{\xi_i^2|\mathscr{F}_{i-1}\}\Bigg]^2 \le{}& \frac{4c_1^2}{n}\sum_{1\le i,\,j\le n}|\gamma_i||\gamma_j| \\ &+4c_1 n^{-1/2}\sum_{1\le i,\,j\le n}|\gamma_i|\mathbf{I}\{n^{-1/2}\gamma_j < -1/2\} \\ &+\sum_{1\le i,\,j\le n}\mathbf{I}\{n^{-1/2}\gamma_i < -1/2\}\mathbf{I}\{n^{-1/2}\gamma_j < -1/2\}.\end{aligned} \tag{4.12}$$

Note that (recall the definition of σ_i^2)

$$|\gamma_i| \le \|\mathbf{a}\| b_0^{-1}(1 + y_{i-1}^2 + \cdots + y_{i-p}^2) \tag{4.13}$$

and therefore $E\gamma_i^2 \le c_2\|\mathbf{a}\|^2$, where c_2 depends on p, b_0 and Ey_0^4. By the Cauchy–Schwarz inequality we also have that $E[|\gamma_i||\gamma_j|] \le c_2\|\mathbf{a}\|^2$. Using again (4.13) and the existence of Ey_0^4 we get that

$$\begin{aligned}P\{\gamma_i < -n^{1/2}/2\} &\le P\{|\gamma_i| \ge n^{1/2}/2\} \\ &\le P\left\{1 + y_{i-1}^2 + \cdots + y_{i-p}^2 \ge \frac{n^{1/2}b_0}{2\|\mathbf{a}\|}\right\} \\ &\le P\left\{y_{i-1}^2 + \cdots + y_{i-p}^2 \ge \frac{n^{1/2}b_0}{4\|\mathbf{a}\|}\right\} \\ &\le \frac{16p^2\|\mathbf{a}\|^2 Ey_0^4}{nb_0^2},\end{aligned} \tag{4.14}$$

provided $n \ge 16\|\mathbf{a}\|^2/b_0^2$. Hence

$$E\big[|\gamma_i|\mathbf{I}\{n^{-1/2}\gamma_i < -1/2\}\big] \le \big(E\gamma_i^2\big)^{1/2}\big(P\{n^{-1/2}\gamma_i < -1/2\}\big)^{1/2} \le c_3\|\mathbf{a}\|^2 n^{-1/2}$$

and

$$\begin{aligned}&E\big[\mathbf{I}\{n^{-1/2}\gamma_i < -1/2\}\mathbf{I}\{n^{-1/2}\gamma_j < -1/2\}\big] \\ &\quad\le \big(P\{n^{-1/2}\gamma_i < -1/2\}P\{n^{-1/2}\gamma_j < -1/2\}\big)^{1/2} \le \frac{16p^2\|\mathbf{a}\|^2 Ey_0^4}{nb_0^2}.\end{aligned}$$

Therefore, by (4),

$$E\left[\sum_{1\le i\le n} E\{\xi_i^2|\mathscr{F}_{i-1}\}\right]^2 \le c_4\|\mathbf{a}\|^2 n.$$

Since $|\xi_i| \le 2$, the last inequality and Theorem 2.11 on page 23 of Hall and Heyde (1980) give

$$E\left[\max_{p<k\le n}\left|\sum_{p<i\le k}\xi_i\right|^4\right] \le c_5\left[c_4\|\mathbf{a}\|^2 n + 2^4\right] \le c_6\left(\|\mathbf{a}\|^2+1\right)n$$

and therefore by the Markov inequality,

$$P\left\{n^{-1/2}\max_{p<k\le n}\left|\sum_{p<i\le k}\xi_i\right| \ge x\right\} = P\left\{\max_{p<k\le n}\left|\sum_{p<i\le k}\xi_i\right|^4 \ge x^4 n^2\right\} \le \frac{c_6(\|\mathbf{a}\|^2+1)}{x^4 n}.$$

This completes the proof of Lemma 4.3. □

LEMMA 4.4. *If conditions* (1.1)–(1.8) *hold, then for any* $\mathbf{a} \in R^{p+1}$ *we have*

$$\sup_{0\le s\le 1}\sup_{0\le t<\infty}\left|e_n^*(t,s,\mathbf{a}) - e_n(t,s)\right| = o_P(1).$$

PROOF. Following the notation introduced at the beginning of the proof of Lemma 4.3, we write $\xi_i(t)$ and γ_i instead of $\xi_i(t,\mathbf{a})$ and $\gamma_i(\mathbf{a})$ ($\mathbf{a}$ is fixed). To prove the claim, we will verify that for any $x > 0$,

$$\lim_{T\to 0}\limsup_{n\to\infty} P\left\{\sup_{0\le s\le 1}\sup_{0\le t\le T}\left|e_n^*(t,s,\mathbf{a}) - e_n(t,s)\right| \ge x\right\} = 0, \tag{4.15}$$

$$\lim_{\delta\to 0}\limsup_{n\to\infty} P\left\{\sup_{0\le s\le 1}\sup_{T\le t\le \delta n^{1/2}}\left|e_n^*(t,s,\mathbf{a}) - e_n(t,s)\right| \ge x\right\} = 0 \tag{4.16}$$

for any $T > 0$, and

$$\sup_{0\le s\le 1}\sup_{\delta n^{1/2}\le t<\infty}\left|e_n^*(t,s,\mathbf{a}) - e_n(t,s)\right| = o_P(1) \tag{4.17}$$

for any $\delta > 0$.

Verification of (4.15). Fix $\epsilon > 0$. By condition (1.2) there is $T^* \le 1$ such that $\sup_{0\le t\le T^*} tf(t) \le \epsilon$. Define $0 = t_0 < t_1 < \cdots < t_N < T^*/2 \le t_{N+1}$ satisfying

$$F(t_{i+1}) - F(t_i) = \epsilon n^{-1/2}, \qquad 0 \le i \le N.$$

It is clear that $N \le n^{1/2}\epsilon^{-1} + 1$ and so by Lemma 4.3 we have

$$P\left\{n^{-1/2}\max_{1\le j\le N+1}\max_{p<k\le n}\left|\sum_{p<i\le k}\xi_i(t_j)\right| \ge x\right\} \le c_1 n^{-1/2}. \tag{4.18}$$

It thus remains to show that

$$\limsup_{n\to\infty} P\left\{ n^{-1/2} \max_{0\le j\le N} \sup_{t_j\le t\le t_{j+1}} \max_{p<k\le n} \left| \sum_{p<i\le k} \xi_i(t) - \sum_{p<i\le k} \xi_i(t_j) \right| \ge x \right\} = 0. \tag{4.19}$$

For any $t_j \le t \le t_{j+1}$ we have

$$\begin{aligned}
&\max_{p<k\le n} \sup_{t_j\le t\le t_{j+1}} \Bigg| \sum_{p<i\le k} \Big[\big(\mathbf{I}\{\varepsilon_i^2 \le t + tn^{-1/2}\gamma_i\} - F(t+tn^{-1/2}\gamma_i) \big) \\
&\qquad\qquad - \big(\mathbf{I}\{\varepsilon_i^2 \le t_j + t_j n^{-1/2}\gamma_i\} - F(t_j + t_j n^{-1/2}\gamma_i) \big) \Big] \Bigg| \\
&\le \max_{p<k\le n} \sup_{t_j\le t\le t_{j+1}} \Bigg| \sum_{p<i\le k} \big(\mathbf{I}\{\varepsilon_i^2 \le t + tn^{-1/2}\gamma_i\} - \mathbf{I}\{\varepsilon_i^2 \le t_j + t_j n^{-1/2}\gamma_i\} \big) \Bigg| \\
&\quad + \max_{p<k\le n} \sup_{t_j\le t\le t_{j+1}} \Bigg| \sum_{p<i\le k} \big(F(t+tn^{-1/2}\gamma_i) - F(t_j + t_j n^{-1/2}\gamma_i) \big) \Bigg| \\
&\le \max_{p<k\le n} \sum_{p<i\le k} \big(\mathbf{I}\{\varepsilon_i^2 \le t_{j+1} + t_{j+1} n^{-1/2}\gamma_i\} - \mathbf{I}\{\varepsilon_i^2 \le t_j + t_j n^{-1/2}\gamma_i\} \big) \\
&\quad \times \mathbf{I}\{n^{-1/2}\gamma_i \ge -1/2\} + \max_{p<k\le n} \sum_{p<i\le k} \big(F(t_{j+1} + t_{j+1} n^{-1/2}\gamma_i) \\
&\quad - F(t_j + t_j n^{-1/2}\gamma_i) \big) \mathbf{I}\{n^{-1/2}\gamma_i \ge -1/2\} + 2 \sum_{1\le i\le n} \mathbf{I}\{n^{-1/2}\gamma_i < -1/2\} \\
&\le \max_{p<k\le n} \Bigg| \sum_{p<i\le k} \Big[\big(\mathbf{I}\{\varepsilon_i^2 \le t_{j+1} + t_{j+1} n^{-1/2}\gamma_i\} - F(t_{j+1} + t_{j+1} n^{-1/2}\gamma_i) \big) \\
&\quad - \big(\mathbf{I}\{\varepsilon_i^2 \le t_j + t_j n^{-1/2}\gamma_i\} - F(t_j + t_j n^{-1/2}\gamma_i) \big) \Big] \mathbf{I}\{n^{-1/2}\gamma_i \ge -1/2\} \Bigg| \\
&\quad + 2 \sum_{p<i\le n} \big(F(t_{j+1} + t_{j+1} n^{-1/2}\gamma_i) - F(t_j + t_j n^{-1/2}\gamma_i) \big) \mathbf{I}\{n^{-1/2}\gamma_i \ge -1/2\} \\
&\quad + 2 \sum_{1\le i\le n} \mathbf{I}\{n^{-1/2}\gamma_i < -1/2\} \\
&\le \max_{p<k\le n} \Bigg| \sum_{p<i\le k} \big(\mathbf{I}\{\varepsilon_i^2 \le t_{j+1} + t_{j+1} n^{-1/2}\gamma_i\} - F(t_{j+1} + t_{j+1} n^{-1/2}\gamma_i) \big) \\
&\qquad\qquad - \sum_{p<i\le k} \big(\mathbf{I}\{\varepsilon_i^2 \le t_j + t_j n^{-1/2}\gamma_i\} - F(t_j + t_j n^{-1/2}\gamma_i) \big) \Bigg| \\
&\quad + 2 \sum_{p<i\le n} \big(F(t_{j+1} + t_{j+1} n^{-1/2}\gamma_i) - F(t_j + t_j n^{-1/2}\gamma_i) \big) \mathbf{I}\{n^{-1/2}\gamma_i \ge -1/2\} \\
&\quad + 4 \sum_{1\le i\le n} \mathbf{I}\{n^{-1/2}\gamma_i < -1/2\}
\end{aligned}$$

$$\leq \max_{p<k\leq n}\left[\left|\sum_{p<i\leq k}\xi_i(t_j)\right|+\left|\sum_{p<i\leq k}\xi_i(t_{j+1})\right|\right]$$
$$+\max_{p<k\leq n}|V_k(j)|+2W(j)+4\sum_{1\leq i\leq n}\mathbf{I}\{n^{-1/2}\gamma_i<-1/2\},$$

where

$$V_k(j)=\sum_{p<i\leq k}\left[\left(\mathbf{I}\{\varepsilon_i^2\leq t_{j=1}\}-F(t_{j+1})\right)-\left(\mathbf{I}\{\varepsilon_i^2\leq t_j\}-F(t_j)\right)\right]$$

and

$$W(j)=\sum_{p<i\leq n}\left(F(t_{j+1}+t_{j+1}n^{-1/2}\gamma_i)-F(t_j+t_jn^{-1/2}\gamma_i)\right)\mathbf{I}\{n^{-1/2}\gamma_i\geq-1/2\}.$$

A similar, in fact, simpler argument shows that

$$\max_{p<k\leq n}\sup_{t_j\leq t\leq t_{j+1}}\left|\sum_{p<i\leq k}\left[\left(\mathbf{I}\{\varepsilon_i^2\leq t\}-F(t)\right)-\left(\mathbf{I}\{\varepsilon_i^2\leq t_j\}-F(t_j)\right)\right]\right|$$
$$\leq\max_{p<k\leq n}|V_k(j)|+2n(F(t_{j+1})-F(t_j)).$$

Thus we get

$$\max_{p<k\leq n}\sup_{t_j\leq t\leq t_{j+1}}\left|\sum_{p<i\leq k}\xi_i(t)-\sum_{p<i\leq k}\xi_i(t_j)\right|$$
$$\leq\max_{p<k\leq n}\left[\left|\sum_{p<i\leq k}\xi_i(t_j)\right|+\left|\sum_{p<i\leq k}\xi_i(t_{j+1})\right|\right]+2\max_{p<k\leq n}|V_k(j)|+2W(j) \tag{4.20}$$
$$+4\sum_{1\leq i\leq n}\mathbf{I}\{n^{-1/2}\gamma_i<-1/2]+2n(F(t_{j+1})-F(t_j)).$$

Now we establish appropriate upper bounds for the terms on the right-hand side of (4.20) which (when divided by $n^{1/2}$) will imply (4.19). The first two terms are asymptotically negligible by (4.18). Focusing on the $V_k(j)$, we see that by the modulus of continuity of the empirical process [and (4.22) below],

$$\max_{0\leq j\leq N}\sup_{0\leq s\leq 1}|e_n(t_{j+1},s)-e_n(t_j,s)|=o_P(1) \tag{4.21}$$

[cf., e.g., Shorack and Wellner (1986), page 542]. By the definition of the t_j, $0\leq j\leq N+1$, we have

$$\max_{0\leq j\leq N}n(F(t_{j+1})-F(t_j))\leq\varepsilon n^{1/2}. \tag{4.22}$$

If $n^{-1/2}\gamma_i \geq -1/2$, then we have $t/(t \wedge (t+tn^{-1/2}\gamma_i)) \leq 2$. Hence by the mean-value theorem and the choice of T^* we have

$$\begin{aligned}\max_{0\leq j\leq N} W(j) &\leq 2 \max_{0\leq j\leq N+1} \sum_{p<i\leq n} \left|F(t_j+t_j n^{-1/2}\gamma_i)-F(t_j)\right| \mathbf{I}\{n^{-1/2}\gamma_i \geq -1/2\} \\ &\quad + \max_{0\leq j\leq N} n(F(t_{j+1})-F(t_j)) \\ &\leq 2 \max_{1\leq j\leq N+1} \sum_{p<i\leq n} f(\xi_{i,j}) t_j n^{-1/2}|\gamma_i| \mathbf{I}\{n^{-1/2}\gamma_i \geq -1/2\} + \varepsilon n^{1/2} \\ &\leq 2 \max_{0\leq j\leq N+1} \sum_{p<i\leq n} \xi_{i,j} f(\xi_{i,j})(t_j/\xi_{i,j}) n^{-1/2}|\gamma_i| \mathbf{I}\{n^{-1/2}\gamma_i \geq -1/2\} + \varepsilon n^{1/2} \\ &\leq 4\varepsilon n^{-1/2} \sum_{p<i\leq n} |\gamma_i| + \varepsilon n^{1/2},\end{aligned} \tag{4.23}$$

where $\xi_{i,j}$ is between t_j and $t_j+t_j n^{-1/2}\gamma_i$. Using the definitions of γ_i and σ_i and the ergodic theorem we have

$$\frac{1}{n} \sum_{p<i\leq n} |\gamma_i| \leq \frac{\|\mathbf{a}\|+1}{b_0} \frac{p+1}{n} \sum_{p<i\leq n} y_i^2 = O_P(1). \tag{4.24}$$

Applying (4.14) we conclude that

$$E \sum_{p<i\leq n} \mathbf{I}\{n^{-1/2}\gamma_i < -1/2\} = O(1). \tag{4.25}$$

Verification of (4.16). We will use again Lemma 4.3 and (4.20). Let $t_0 = T$ and $t_i = T + i\delta^{1/6} n^{-1/2}, 1 \leq i \leq M-1, t_{M-1} \leq \delta n^{1/2} \leq t_M$. We note that $M \leq n\delta^{5/6} + 1$. By Lemma 4.3 we have

$$P\left\{\max_{0\leq j\leq M} \max_{p<k\leq n} n^{-1/2} \left|\sum_{p<i\leq k} \xi_i(t_j)\right| \geq x\right\} \leq \frac{c_2}{x^4} \frac{M+1}{n} \leq \frac{c_3}{x^4}\left(\delta^{5/6} + \frac{1}{n}\right). \tag{4.26}$$

Using again the modulus of continuity of $e_n(t,s)$, we get

$$\max_{0\leq j\leq M} \max_{p<k\leq n} |V_k(j)| = o_P(n^{1/2}). \tag{4.27}$$

Since $f(t)$ is bounded on $[T,\infty)$, say, by c_4, we obtain

$$\max_{0\leq j\leq M} n(F(t_{j+1})-F(t_j)) \leq c_4 \delta^{1/6} n^{1/2}, \tag{4.28}$$

and by the mean-value theorem

$$\begin{aligned}\max_{0\leq j\leq M} W(j) &\leq c_4 \max_{0\leq j\leq M} \sum_{p<i\leq n} |t_{j+1}-t_j|(1+n^{-1/2}|\gamma_i|) \mathbf{I}\{n^{-1/2}\gamma_i \geq -1/2\} \\ &\leq c_4 \delta^{1/6} n^{-1/2}\left(n + n^{-1/2} \sum_{p<i\leq n} |\gamma_i|\right) \leq \delta^{1/6} n^{1/2} O_P(1)\end{aligned} \tag{4.29}$$

on account of (4.24). Relation (4.16) follows from (4.20), (4.26)–(4.29) and (4.25).

Verification of (4.17). We use Lemma 4.3 and the following bound:

$$
\begin{aligned}
&\max_{p<k\le n}\ \sup_{\delta n^{1/2}\le t<\infty}\left|\sum_{p<i\le k}\xi_i(t)\right|\\
&\quad\le \max_{p<k\le n}\ \sup_{\delta n^{1/2}\le t<\infty}\sum_{p<i\le k}\left(1-\mathbf{I}\{\varepsilon_i^2\le t+tn^{-1/2}\gamma_i\}\right)\\
&\qquad+\max_{p<k\le n}\ \sup_{\delta n^{1/2}\le t<\infty}\sum_{p<i\le k}\left(1-F(t+tn^{-1/2}\gamma_i)\right)\\
&\qquad+\sum_{1\le i\le n}\left(1-\mathbf{I}\{\varepsilon_i^2\le\delta n^{1/2}\}\right)+n\left(1-F(\delta n^{1/2})\right)\\
&\quad\le \sum_{p<i\le n}\left(1-\mathbf{I}\{\varepsilon_i^2\le\delta n^{1/2}+\delta\gamma_i\}\right)\mathbf{I}\{n^{-1/2}\gamma_i\ge -1/2\}\\
&\qquad+\sum_{p<i\le n}\left(1-F(\varepsilon_i^2\le\delta n^{1/2}+\delta\gamma_i)\right)\mathbf{I}\{n^{-1/2}\gamma_i\ge -1/2\}\\
&\qquad+2\sum_{p<i\le n}\mathbf{I}\{n^{1/2}\gamma_i<-1/2\}+\left|\sum_{1\le i\le n}\left(\mathbf{I}\{\varepsilon_i^2\le\delta n^{1/2}\}-F(\delta n^{1/2})\right)\right|\\
&\qquad+2n\left(1-F(\delta n^{1/2})\right)\\
&\quad\le\left|\sum_{p<i\le n}\xi_i(\delta n^{1/2})\right|+2\left|\sum_{1\le i\le n}\left(\mathbf{I}\{\varepsilon_i^2\le\delta n^{1/2}\}-F(\delta n^{1/2})\right)\right|\\
&\qquad+2\sum_{p<i\le n}\left(1-F(\delta n^{1/2}+\delta\gamma_i)\right)\mathbf{I}\{n^{-1/2}\gamma_i\ge -1/2\}\\
&\qquad+4\sum_{p<i\le n}\mathbf{I}\{n^{-1/2}\gamma_i<-1/2\}+4n\left(1-F(\delta n^{1/2})\right).
\end{aligned}
\tag{4.30}
$$

By Lemma 4.3,

$$
\left|\sum_{p<i\le n}\xi_i(\delta n^{1/2})\right|=o_P(n^{1/2}).
\tag{4.31}
$$

Observing that $E(\sum_{1\le i\le n}(I\{\epsilon_i^2\le\delta n^{1/2}\}-F(\delta n^{1/2})))^2=nF(\delta n^{1/2})(1-F(\delta n^{1/2}))=o(n)$, by the Chebyshev inequality we have

$$
\left|\sum_{1\le i\le n}\left(\mathbf{I}\{\varepsilon_i^2\le\delta n^{1/2}\}-F(\delta n^{1/2})\right)\right|=o_P(n^{1/2}).
\tag{4.32}
$$

Also

$$
\begin{aligned}
&\sum_{p<i\le n}\left(1-F(\delta n^{1/2}+\delta\gamma_i)\right)\mathbf{I}\{n^{-1/2}\gamma_i\ge -1/2\}\\
&\qquad\le n\left(1-F\left(\delta n^{1/2}/2\right)\right)=o(1),
\end{aligned}
\tag{4.33}
$$

by assumption (1.7) which implies that $nP(\varepsilon_0^4>n)\to 0$. Relations (4.30)–(4.33) and (4.25) imply (4.17). □

For any $A>0$ define $\mathbf{A}=[-A,A]^{p+1}$.

LEMMA 4.5. *If conditions* (1.1)–(1.8) *hold, then for any* $A>0$,

$$\sup_{\mathbf{a}\in\mathbf{A}}\sup_{0\le s\le 1}\sup_{0\le t<\infty}|e_n^*(t,s,\mathbf{a})-e_n(t,s)|=o_P(1).$$

PROOF. Denote

$$X_n(\mathbf{a})=\sup_{0\le s\le 1}\sup_{0\le t<\infty}\left|n^{-1/2}\sum_{p<i\le ns}\xi_i(t,\mathbf{a})\right|.$$

In Lemma 4.4. we showed that $X_n(\mathbf{a})\to^P 0$ for every $\mathbf{a}$. Here we show that $\sup_{\mathbf{a}\in\mathbf{A}} X_n(\mathbf{a})\to^P 0$.

Consider $\varepsilon>0$ such that $2A/\epsilon$ is an integer, which will be specified at the end of the proof. Define $a(k)=-A+k\varepsilon, 1\le k\le K\equiv 2A/\epsilon$. Set $\mathbf{k}=(k_0,\ k_1,\dots,k_p)$, $1\le k_0,\ k_1,\dots,k_p\le K$ and consider the grid of $(p+1)K$ points $\mathbf{a}(\mathbf{k})=(a(k_0),\ a(k_1),\dots,a(k_p))$. Consider also the corresponding $(p+1)K$ cells,

$$\mathbf{A}(\mathbf{k})=\{(a_0,a_1,\dots,a_p)\in\mathbf{A}\colon a(k_i)-\epsilon\le a_i\le a(k_i)\}$$

and the points $\mathbf{a}^*(\mathbf{k})=(a(k_0)-\epsilon,a(k_1)-\epsilon,\dots,a(k_p)-\epsilon)$. Observe that

$$\begin{aligned}&F\big(t+tn^{-1/2}\gamma_i(a_0,\dots,a_{j-1},a(k_j),a_{j+1},\dots,a_p)\big)\\&\qquad-F\big(t+tn^{-1/2}\gamma_i(a_0,\dots,a_{j-1},a(k_j)-\epsilon,a_{j+1},\dots,a_p)\big)\\&\qquad=\begin{cases}f(\xi_{i0}^*)tn^{-1/2}\epsilon/\sigma_i^2, & \text{if } j=0,\\ f(\xi_{ij}^*)tn^{-1/2}\epsilon y_{i-j}^2/\sigma_i^2, & \text{if } 1\le j\le p,\end{cases}\end{aligned}$$

where ξ_{ij}^* is a point between the arguments of F above. Consider the set

$$\Omega_n^*=\left\{\omega:\max_{p<i\le n}\sup_{\mathbf{a}\in\mathbf{A}}n^{-1/2}|\gamma_i(\mathbf{a})|>1/2\right\}.$$

and notice that by (4.8), $\lim_{n\to\infty}P\{\Omega_n^*\}=0$. In the remainder of the proof we work on the complement of Ω_n^*. Since $t/\xi_{ij}^*\le 2$,

$$\begin{aligned}&F\big(t+tn^{-1/2}\gamma_i(a_0,\dots,a_{j-1},a(k_j),a_{j+1},\dots,a_p)\big)\\&\qquad-F\big(t+tn^{-1/2}\gamma_i(a_0,\dots,a_{j-1},a(k_j)-\epsilon,a_{j+1},\dots,a_p)\big)\\&\qquad\le\begin{cases}2c_1tn^{-1/2}\epsilon/\sigma_i^2, & \text{if } j=0,\\ 2c_1tn^{-1/2}\epsilon y_{i-j}^2/\sigma_i^2, & \text{if } 1\le j\le p,\end{cases}\end{aligned}\tag{4.34}$$

where $c_1=\sup_{0\le t<\infty}tf(t)$. Applying (4.34) consecutively to each coordinate, we obtain for any $t>0$ and any $\mathbf{k}$,

$$\begin{aligned}&\sup_{\mathbf{a}\in\mathbf{A}(\mathbf{k})}\left|\sum_{p<i\le k}\big[F\big(t+tn^{-1/2}\gamma_i\mathbf{a}\big)-F\big(t+tn^{-1/2}\gamma_i(\mathbf{a}(\mathbf{k}))\big)\big]\right|\\&\qquad\le 2c_1\epsilon n^{1/2}b_0^{-1}\left[1+n^{-1}\sum_{p<i\le n}\sum_{1\le j\le p}y_{i-j}^2\right],\end{aligned}\tag{4.35}$$

and so

$$
\begin{aligned}
&\sup_{\mathbf{a}\in\mathbf{A}(\mathbf{k})}\left|\sum_{p<i\le k}[\xi_i(t,\mathbf{a})-\xi_i(t,\mathbf{a}(\mathbf{k}))]\right| \\
&\qquad\le 2c_1\epsilon n^{1/2}b_0^{-1}\left[1+n^{-1}\sum_{p<i\le n}\sum_{1\le j\le p}y_{i-j}^2\right] \\
&\qquad\quad +2\sum_{p<i\le k}\left[\mathbf{I}\{\varepsilon_i^2\le t+tn^{-1/2}\gamma_i(\mathbf{a}(\mathbf{k}))\}-\mathbf{I}\{\varepsilon_i^2\le t+tn^{-1/2}\gamma_i(\mathbf{a}^*(\mathbf{k}))\}\right] \qquad (4.36)\\
&\qquad\le 4c_1\epsilon n^{1/2}b_0^{-1}\left[1+n^{-1}\sum_{p<i\le n}\sum_{1\le j\le p}y_{i-j}^2\right] \\
&\qquad\quad +\left|\sum_{p<i\le k}\xi_i(t,\mathbf{a}(\mathbf{k}))\right|+\left|\sum_{p<i\le k}\xi_i(t,\mathbf{a}^*(\mathbf{k}))\right|.
\end{aligned}
$$

Observe that

$$
\begin{aligned}
\sup_{a\in\mathbf{A}}X_n(\mathbf{a}) \le{}& \sup_{\mathbf{k}}\sup_{\mathbf{a}\in\mathbf{A}(\mathbf{k})}\sup_{0\le s\le 1}\sup_{0\le t<\infty}\left|n^{-1/2}\sum_{p<i\le k}[\xi_i(t,\mathbf{a})-\xi_i(t,\mathbf{a}(\mathbf{k}))]\right| \\
&+\sup_{\mathbf{k}}X_n(\mathbf{a}(\mathbf{k})).
\end{aligned}
$$

Hence, by (4.36),

$$
\begin{aligned}
\sup_{\mathbf{a}\in\mathbf{A}}X_n(\mathbf{a}) \le{}& c_2\epsilon\left[1+n^{-1}\sum_{p<i\le n}\sum_{1\le j\le p}y_{i-j}^2\right] \\
&+2\sup_{\mathbf{k}}X_n(\mathbf{a}(\mathbf{k}))+\sup_{\mathbf{k}}X_n(\mathbf{a}^*(\mathbf{k}))=:\epsilon Y_n+Z_n.
\end{aligned}
$$

Fix $r>0$. We must show that $\lim_{n\to\infty}P\{\sup_{\mathbf{a}\in\mathbf{A}}X_n(\mathbf{a})>r\}=0$. By the ergodic theorem, $Y_n:=c_2[1+n^{-1}\sum_{p<i\le n}\sum_{1\le j\le p}y_{i-j}^2]$ tends in probability to a constant, say, c_3. We now fix ε so small that $\varepsilon c_3<r/2$. With ε fixed, $Z_n:=2\sup_{\mathbf{k}}X_n(\mathbf{a}(\mathbf{k}))+\sup_{\mathbf{k}}X_n(\mathbf{a}^*(\mathbf{k}))$ is a maximum over a finite number of points and so $P\{Z_n>r/2\}\to 0$ by Lemma 4.4 and $P\{\epsilon Y_n>r/2\}\to 0$ by the choice of ε. □

LEMMA 4.6. *If* (1.1)–(1.8) *hold, then*

$$
\sup_{0\le s\le 1}\sup_{0\le t<\infty}|\hat{e}_n(t,s)-(e_n(t,s)+u_n(t,s))|=o_P(1),
$$

where $u_n(t,s)$ *is defined by* (4.2) *and* (4.1).

PROOF. We use the inequality

$$
\begin{aligned}
|\hat{e}_n(t,s)-(e_n(t,s)+u_n(t,s))| \le{}& |\hat{e}_{n,1}(t,s)-e_n(t,s)|+|\hat{e}_{n,2}(t,s)-h_n(t,s)| \\
&+|h_n(t,s)-u_n(t,s)|,
\end{aligned}
$$

where h_n is given by (1.10).

Lemmas 4.1 and 4.5 yield [cf. (4.9)]

$$\sup_{0\le s\le 1}\sup_{0\le t<\infty}|\hat{e}_{n,1}(t,s)-e_n(t,s)|=o_P(1). \tag{4.37}$$

Indeed, denoting, as in the proof of Lemma 4.5, $X_n(\mathbf{a})=\sup_{0\le s\le 1}\ \sup_{0\le t<\infty}|e_n^*(t,s,\mathbf{a})-e_n(t,s)|$ and $\hat{\mathbf{a}}_n=n^{1/2}(\hat{\mathbf{b}}_n-\mathbf{b})$, we have

$$P\{X_n(\hat{\mathbf{a}}_n)>r\}\le P\left\{\sup_{\mathbf{a}\in\mathbf{A}}X_n(\mathbf{a})>r\right\}+P\{\hat{\mathbf{a}}_n\notin\mathbf{A}\}.$$

By Lemma 4.5 this implies $\limsup_{n\to\infty}P\{X_n(\hat{\mathbf{a}}_n)>r\}\le P\{\hat{\mathbf{a}}_n\notin\mathbf{A}\}$ and so (4.37) follows because by Lemma 4.1, letting $A\to\infty$, $P\{\hat{\mathbf{a}}_n\notin\mathbf{A}\}$ can be made arbitrarily small.

By Lemma 4.2,

$$\sup_{0\le s\le 1}\sup_{0\le t<\infty}|e_{n,2}(t,s)-h_n(t,s)|=o_P(1).$$

To verify that

$$\sup_{0\le s\le 1}\sup_{0\le t<\infty}|u_n(t,s)-h_n(t,s)|=o_P(1), \tag{4.38}$$

observe that by (1.3),

$$\begin{aligned}
&h_n(t,s)-u_n(t,s)\\
&\quad=\left[\frac{1}{n}\sum_{p<i\le ns}\frac{1}{\sigma_i^2}-s\beta_0\right]tf(t)n^{-1/2}\sum_{1\le j\le n}l_0(\varepsilon_j^2)f_0(\varepsilon_{j-1},\varepsilon_{j-2},\dots)\\
&\qquad+\sum_{1\le k\le p}\left[\frac{1}{n}\sum_{p<i\le ns}\frac{y_{i-k}^2}{\sigma_i^2}-s\beta_k\right]tf(t)n^{-1/2}\sum_{1\le j\le n}l_k(\varepsilon_j^2)f_k(\varepsilon_{j-1},\varepsilon_{j-2},\dots)\\
&\qquad+o_P(1)tf(t)\frac{1}{n}\sum_{p<i\le ns}\frac{1}{\sigma_i^2}\left(1+\sum_{1\le k\le p}y_{i-k}^2\right).
\end{aligned}$$

In view of (1.1) and (1.2) and (1.4)–(1.6), it remains to verify that for each $1\le k\le p$,

$$\sup_{0\le s\le 1}\left[\frac{1}{n}\sum_{p<i\le ns}\frac{y_{i-k}^2}{\sigma_i^2}-s\beta_k\right]=o_P(1), \tag{4.39}$$

with an obvious modification for $k=0$. Since for each fixed k, the random variables $y_{i-k}^2/\sigma_i^2-\beta_k$ form a stationary geometrically mixing sequence [see Guegan and Diebolt (1994)] with zero mean and finite variance, their normalized partial sum process converges weakly to a Brownian motion, and so the left-hand side of (4.39) is in fact $O_P(n^{-1/2})$. $\square$

LEMMA 4.7. *Let* $e_n(t,s)$ *and* ζ_n *be defined by* (1.9) *and* (4.1), *respectively. If* (1.1)–(1.8) *hold, then* $\{(e_n(t,s),\zeta_n),\ 0\le t<\infty, 0\le s\le 1\}$ *converges weakly in* $\mathscr{D}([0,\infty]\times[0,1])\times R$ *to the Gaussian process* $\{(K(F(t),s),\zeta), 0\le t<\infty, 0\le s\le 1\}$ *defined by* (1.12)–(1.15).

PROOF. We will verify tightness and the convergence of the finite-dimensional distributions. Since $e_n(t,s)$ is a sequential empirical process of independent identically distributed random variables, its tightness is well known; see, for example, Csörgő and Révész (1981). If we show that ζ_n converges in distribution to a normal random variable, the tightness will be proved because the vector valued process $(e_n(t,s),\zeta_n)$ is tight if the coordinates are tight.

Fix an integer M, real numbers $\lambda_1,\lambda_2\dots,\lambda_M,\lambda_{M+1}$, $0\le t_1,t_2,\dots,t_M<\infty$, $0=s_0\le s_1\le s_2\le\cdots\le s_M\le 1$ and define

$$Z(n)=\sum_{1\le m\le M}\lambda_m e_n(t_m,s_m)+\lambda_{M+1}\zeta_n.$$

It is easy to see that

$$n^{1/2}Z(n)=\sum_{1\le j\le n}\tau_j,$$

where, setting $f_k^-(j)=f_k(\varepsilon_{j-1},\varepsilon_{j-2},\dots)$,

$$\tau_j=\sum_{m(j)\le m\le M}\lambda_m\big(\mathbf{I}\{\varepsilon_j^2\le t_m\}-F(t_m)\big)+\lambda_{M+1}\sum_{0\le k\le p}\beta_k l_k(\varepsilon_j^2)f_k^-(j), \tag{4.40}$$

where $m(j)=i$ if $(ns_{i-1})\vee p<j\le ns_i, 1\le i\le M$ and

$$\tau_j=\lambda_{M+1}\sum_{0\le k\le p}\beta_k l_k(\varepsilon_j^2)f_k^-(j)\quad \text{if } 1\le j\le p \text{ or } \ ns_M<j\le n.$$

Recall that $\mathscr{F}_k$ is the σ-algebra generated by $\varepsilon_k,\varepsilon_{k-1},\dots$. We will use Theorem 3.5 of Hall and Heyde (1980) to show that $Z(n)$ is asymptotically normal. To do so, we will verify that

$$\max_{1\le j\le n}E\{\tau_j^2|\mathscr{F}_{j-1}\}=o_P(n^{1/2}), \tag{4.41}$$

$$E\left|\sum_{1\le i\le n}\big(E\{\tau_j^2|\mathscr{F}_{j-1}\}-E\tau_j^2\big)\right|=o(n) \tag{4.42}$$

and

$$E\left|\sum_{1\le i\le n}\big(\tau_j^2-E\tau_j^2\big)\right|=o(n). \tag{4.43}$$

Focusing on the more complex case (4.40), we write

$$\begin{aligned}\tau_j^2&=\Big[\sum_{m(j)\le m\le M}\lambda_m\big(\mathbf{I}\{\varepsilon_j^2\le t_m\}-F(t_m)\big)\Big]^2\\&\quad+\lambda_{M+1}^2\sum_{0\le k,k'\le p}\beta_k\beta_{k'}l_k(\varepsilon_j^2)l_{k'}(\varepsilon_j^2)f_k^-(j)f_{k'}^-(j)\\&\quad+2\lambda_{M+1}\sum_{0\le k\le p}\beta_k f_k^-(j)l_k(\varepsilon_j^2)\sum_{m(j)\le m\le M}\lambda_m\big(\mathbf{I}\{\varepsilon_j^2\le t_m\}-F(t_m)\big)\end{aligned} \tag{4.44}$$

and therefore

$$
\begin{aligned}
E\{\tau_j^2|\mathscr{F}_{j-1}\} = &\sum_{m(j)\le m\le M} \lambda_m^2\big[F(t_m)-F^2(t_m)\big] \\
&+\lambda_{M+1}^2 \sum_{0\le k,k'\le p} \beta_k\beta_{k'} E[l_k(\varepsilon_j^2)l_{k'}(\varepsilon_j^2)]f_k^-(j)f_{k'}^-(j) \\
&+2\lambda_{M+1}\sum_{0\le k\le p}\beta_k f_k^-(j)\sum_{m(j)\le m\le M}\lambda_m\int_0^{t_m} l_k(x)f(x)\,dx
\end{aligned}
\tag{4.45}
$$

and

$$
\begin{aligned}
E\tau_j^2 = &\sum_{m(j)\le m\le M}\lambda_m^2[F(t_m)-F^2(t_m)]+\lambda_{M+1}^2\sum_{0\le k,k'\le p}\beta_k\gamma_{kk'}\beta_{k'} \\
&+2\lambda_{M+1}\sum_{0\le k\le p}\beta_k\sum_{m(j)\le m\le M}\lambda_m g_k(t_m).
\end{aligned}
\tag{4.46}
$$

Since for each $k, f_k^-(j)\equiv f_i(\varepsilon_{j-1},\varepsilon_{j-2},\dots)$, $1\le j<\infty$ is a stationary sequence satisfying (1.6), $\max_{1\le j\le n}|f_i^-(j)|=o_P(n^{1/2})$, and so (4.41) follows from (4.45). Using expressions (4.44)–(4.46), we see that (4.41) and (4.43) follow from the mean ergodic theorem.

By computing $EZ^2(n)$, we obtain the desired covariance structure. □

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L. HORVÁTH
DEPARTMENT OF MATHEMATICS
UNIVERSITY OF UTAH
155 SOUTH 1440 EAST
SALT LAKE CITY
UTAH 84112-0090
EMAIL: horvath@math.utah.edu

P. KOKOSZKA
DEPARTMENT OF MATHEMATICS AND STATISTICS
UTAH STATE UNIVERSITY
3900 OLD MAIN HILL
SALT LAKE CITY, UTAH 84322-3900
EMAIL: piotr@math.usu.edu

G. TEYSSIÈRE
JOINT RESEARCH CENTRE
THE EUROPEAN COMMISSION
ISIS. TP361. VIA E. FERMI 1
I-21020 ISPRA (VA)
ITALY
EMAIL: gilles.teyssiere@jrc.it