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# DETECTING AT-MOST- $m$ CHANGES IN LINEAR REGRESSION MODELS

LAJOS HORVÁTH, WILLIAM POULIOT, AND SHIXUAN WANG

ABSTRACT. In this paper we provide a new procedure to test for at-most- $m$  changes in the time-dependent regression model  $y_t = \mathbf{x}_t^\top \boldsymbol{\beta}_t + e_t, 1 \leq t \leq T$ , i.e.  $\boldsymbol{\beta}_1 = \boldsymbol{\beta}_2 = \dots = \boldsymbol{\beta}_T$  under the no change null hypothesis against the alternative  $y_t = \mathbf{x}_t^\top \boldsymbol{\beta}^{(i)} + e_t$ , if  $k_{i-1}^* < t \leq k_i^*, 1 \leq i \leq m+1$  and  $\boldsymbol{\beta}^{(j)} \neq \boldsymbol{\beta}^{(\ell)}$  for some  $1 \leq j, \ell \leq m+1$  with  $k_0^* = 0, 1 < k_1^* < k_2^* < \dots < k_m^* < T, k_{m+1}^* = T$ . Our procedure is based on weighted sums of the residuals, incorporating the possibility of  $m$  changes. The weak limit of the proposed test statistic is the sum of two double exponential random variables. A small Monte Carlo simulation illustrates the applicability of the limit results in case of small and moderate sample sizes. We compare the new method to the CUSUM and standardized (weighted) CUSUM procedures and obtain the power curves of the test statistics under the alternative. We apply our method to find changes in the unconditional four factor CAPM.

## 1. INTRODUCTION

In the paper we are interested in the time-dependent regression model

$$y_t = \mathbf{x}_t^\top \boldsymbol{\beta}_t + e_t, \quad 1 \leq t \leq T. \quad (1.1)$$

We wish to test the null hypothesis of constant  $\boldsymbol{\beta}_t$ 's

$$H_0 : \boldsymbol{\beta}_1 = \boldsymbol{\beta}_2 = \dots = \boldsymbol{\beta}_T$$

against the at-most- $m$  change points alternative. With the notations  $k_0^* = 0$  and  $k_{m+1}^* = T$ , the case of at-most- $m$  changes alternative can be formulated as

$$H_A : \quad y_t = \mathbf{x}_t^\top \boldsymbol{\beta}^{(i)} + e_t, \quad \text{if } k_{i-1}^* < t \leq k_i^*, \quad 1 \leq i \leq m+1 \text{ and } \boldsymbol{\beta}^{(j)} \neq \boldsymbol{\beta}^{(\ell)}$$

for some  $1 \leq j, \ell \leq m+1$ .

Testing for possible changes was initiated by Quandt (1958, 1960) who suggested maximally selected statistics and provided practical advise how to get critical values. Gombay and Horváth (1994), Horváth (1995) and Horváth and Shao (1995) obtained the limit distributions of some of the test statistics proposed by Quandt (1958, 1960) including maximally selected  $F$ -statistics and the likelihood ratio. McCabe and Harrison (1980) also contribute to this literature and advise the use of ordinary least squares residuals rather than recursive in CUSUM-type tests. Later McCabe (1988), using a multiple decision theory approach, shows that the CUSUM test is Bayes for structural stability in scale and variance models, and also that the CUSUM-of-squares test is a localised Bayes rule for structural stability in variance of linear regression models. Turning to estimation of the time change of change,

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Hušková (1996) gave large sample approximation for the estimator of the time of change assuming that we have exactly one change in the regressor during the observation period. The independence of the error terms are assumed in these early papers. Andrews (1993) provides a general methodology to test for the stability of random systems from an economic view point. Ghysels et al. (1997), Bai (1999), Bai and Perron (1998), Hall et al. (2012) followed the suggestions of Andrews (1993) and they also used the maximally selected statistics but the maxima were not computed for all observations points, a fraction of early and late observations were excluded. Aue et al. (2008, 2012a) used the maximally selected likelihood ratio method to test for stability of the parameter against exactly one change. However, they also showed that the derived tests are consistent against several changes under the alternative.

Our test for  $H_0$  against  $H_A$  uses the residuals

$$\hat{\epsilon}_t = y_t - \mathbf{x}_t^\top \hat{\boldsymbol{\beta}}_T, \quad 1 \leq t \leq T, \quad (1.2)$$

where

$$\hat{\boldsymbol{\beta}}_T = (\mathbf{X}_T \mathbf{X}_T^\top)^{-1} \mathbf{X}_T^\top \mathbf{Y}_T,$$

with

$$\mathbf{Y}_T = \begin{bmatrix} y_1 \\ y_2 \\ \vdots \\ y_T \end{bmatrix} \quad \text{and} \quad \mathbf{X}_T = \begin{bmatrix} x_{1,1} & x_{2,1} & \cdots & x_{T,1} \\ x_{1,2} & x_{2,2} & \cdots & x_{T,2} \\ \vdots & \vdots & \ddots & \vdots \\ x_{1,d} & x_{2,d} & \cdots & x_{T,d} \end{bmatrix}.$$

In this paper we suggest three test statistics based on the sums of the residuals. The classical CUSUM statistic

$$D_T = T^{-1/2} \max_{1 \leq \ell \leq T} \left| \sum_{t=1}^{\ell} \hat{\epsilon}_t - \frac{\ell}{T} \sum_{t=1}^T \hat{\epsilon}_t \right|$$

which together with the standardized CUSUM

$$H_T = T^{1/2} \max_{1 \leq \ell < T} (\ell(T - \ell))^{-1/2} \left| \sum_{t=1}^{\ell} \hat{\epsilon}_t - \frac{\ell}{T} \sum_{t=1}^T \hat{\epsilon}_t \right|$$

are one of the most often used statistics in change point analysis. Aue and Horváth (2013) contains a review of change point detection in time series and it also provides a historical account of CUSUM procedures. Horváth and Rice (2014) explain how mathematical and probabilistic tools can be used to extend classical change point methods to time series models. The motivation for  $H_T$  and  $D_T$  is based on the likelihood ratio method when there is exactly one change in the parameters under the alternative. Since we allow up to  $\mathbf{m}$  changes under the alternative, we propose a modification of the CUSUM statistics. Let

$$M(k_1, \dots, k_m) = |M_1(k_1)| + |M_2(k_1, k_2)| + \dots + |M_m(k_{m-1}, k_m)| + |M_{m+1}(k_m)|,$$

$1 \leq k_1 \leq k_2 \leq \dots \leq k_m < T$ , where

$$M_1(k_1) = \frac{1}{\sqrt{k_1}} \left( \sum_{t=1}^{k_1} \hat{\epsilon}_t - \frac{k_1}{T} \sum_{t=1}^T \hat{\epsilon}_t \right),$$

$$M_i(k_{i-1}, k_i) = \frac{1}{\sqrt{T}} \left( \sum_{t=k_{i-1}+1}^{k_i} \hat{\epsilon}_t - \frac{k_i - k_{i-1}}{T} \sum_{t=1}^T \hat{\epsilon}_t \right), \quad 2 \leq i \leq \mathbf{m},$$

and

$$M_{m+1}(k_m) = \frac{1}{\sqrt{T - k_m}} \left( \sum_{t=k_m+1}^T \hat{e}_t - \frac{T - k_m}{T} \sum_{t=1}^T \hat{e}_t \right)$$

and define

$$M_T = \max_{\mathfrak{M}} M(k_1, k_2, \dots, k_m), \quad \text{where } \mathfrak{M} = \{1 \leq k_1 \leq k_2 \leq \dots \leq k_m < T\}$$

We would like to note that  $M(k_1, k_2, \dots, k_m), 1 \leq k_1 \leq k_2 \leq \dots \leq k_m < T$  is also a generalization of the classical CUSUM process (cf. Csörgő and Horváth (1997)), since  $M_1(k_1)$  and  $M_{m+1}(k_m), 1 \leq k_1, k_m < T$  are standardized CUSUM processes starting from the first and the last residual, respectively. The components  $M_1(k_1)$  and  $M_{m+1}(k_m)$  of  $M(k_1, k_2, \dots, k_m)$  are self-normalized and therefore they could be derived from a likelihood argument. First we obtain the joint asymptotic distribution of  $D_T, H_T$  and  $M_T$  in Theorem 2.1. In Theorem 2.2 we derive the joint limit distribution of  $M_T^*, M_T^{(1)}$  and  $M_T^{(2)}$ , where

$$M_T^* = \max_{\mathfrak{M}^*} \sum_{j=2}^m M_j(k_{j-1}, k_j) \quad \text{with } \mathfrak{M}^* = \{1 < k_2 \leq \dots \leq k_m < T\},$$

$$M_T^{(1)} = \max_{1 \leq k \leq T} M_1(k) \quad \text{and} \quad M_T^{(2)} = \max_{1 \leq k < T} M_{m+1}(k).$$

Due to the standardization, the limit distributions of  $M_1(k_1)$  and  $M_{m+1}(k_m)$  are non-standard, they do not follow from weak convergence type results. For the application of the Lagrange multiplier type statistics using the whole sample we refer to Hidalgo and Seo (2013). Jeng (2015) surveys CUSUM and related procedures in financial applications. We also discuss the behavior of  $M_T$  under the alternative  $H_A$ .

In this paper the test statistics are based on the residuals  $\hat{e}_t, 1 \leq t \leq T$  but in a similar matter we can use the weighted residuals  $\tilde{\mathbf{e}}_t = \mathbf{x}_t \hat{e}_t, 1 \leq t \leq T$  (cf. Hušková (1996)). Analogously to  $D_T$  and  $H_T$  one can define

$$\tilde{D}_T = \frac{1}{\sqrt{T}} \max_{1 \leq \ell \leq T} \left( \left( \sum_{t=1}^{\ell} \tilde{\mathbf{e}}_t - \frac{\ell}{T} \sum_{t=1}^T \tilde{\mathbf{e}}_t \right)^\top \Sigma^{-1} \left( \sum_{t=1}^{\ell} \tilde{\mathbf{e}}_t - \frac{\ell}{T} \sum_{t=1}^T \tilde{\mathbf{e}}_t \right) \right)^{1/2}$$

and

$$\tilde{H}_T^2 = \max_{1 \leq \ell < T} \left( \frac{T}{T(T - \ell)} \left( \sum_{t=1}^{\ell} \tilde{\mathbf{e}}_t - \frac{\ell}{T} \sum_{t=1}^T \tilde{\mathbf{e}}_t \right)^\top \Sigma^{-1} \left( \sum_{t=1}^{\ell} \tilde{\mathbf{e}}_t - \frac{\ell}{T} \sum_{t=1}^T \tilde{\mathbf{e}}_t \right) \right)^{1/2},$$

where

$$\Sigma = \sum_{t=-\infty}^{\infty} E [\mathbf{x}_0 e_0 (\mathbf{x}_t e_t)^\top],$$

the long run covariance matrix of the sum of the weighted innovations  $\mathbf{x}_t e_t$ . Now we define

$$\tilde{M}(k_1, \dots, k_m) = \tilde{M}_1^{1/2}(k_1) + \tilde{M}_2^{1/2}(k_1, k_2) + \dots + \tilde{M}_m^{1/2}(k_{m-1}, k_m) + \tilde{M}_{m+1}^{1/2}(k_m),$$

$1 \leq k_1 \leq k_2 \leq \dots \leq k_m < T$ , where

$$\begin{aligned} \tilde{M}_1(k_1) &= \frac{1}{k_1} \left( \sum_{t=1}^{k_1} \tilde{\mathbf{e}}_t - \frac{k_1}{T} \sum_{t=1}^T \tilde{\mathbf{e}}_t \right)^\top \Sigma^{-1} \left( \sum_{t=1}^{k_1} \tilde{\mathbf{e}}_t - \frac{k_1}{T} \sum_{t=1}^T \tilde{\mathbf{e}}_t \right), \\ \tilde{M}_i(k_{i-1}, k_i) &= \frac{1}{T} \left( \sum_{t=k_{i-1}+1}^{k_i} \tilde{\mathbf{e}}_t - \frac{k_i - k_{i-1}}{T} \sum_{t=1}^T \tilde{\mathbf{e}}_t \right)^\top \Sigma^{-1} \left( \sum_{t=k_{i-1}+1}^{k_i} \tilde{\mathbf{e}}_t - \frac{k_i - k_{i-1}}{T} \sum_{t=1}^T \tilde{\mathbf{e}}_t \right), \\ & \hspace{15em} 2 \leq i \leq m, \end{aligned}$$

and

$$\tilde{M}_{m+1}(k_m) = \frac{1}{T - k_m} \left( \sum_{t=k_m+1}^T \tilde{\mathbf{e}}_t - \frac{T - k_m}{T} \sum_{t=1}^T \tilde{\mathbf{e}}_t \right)^\top \Sigma^{-1} \left( \sum_{t=k_m+1}^T \tilde{\mathbf{e}}_t - \frac{T - k_m}{T} \sum_{t=1}^T \tilde{\mathbf{e}}_t \right).$$

The statistics  $\tilde{D}_T$ ,  $\tilde{H}_T$  and  $\max_{1 \leq k_1 \leq k_2 \leq \dots \leq k_m < T} \tilde{M}(k_1, k_2, \dots, k_m)$  can also be applied to test  $H_0$  against  $H_A$ . The derivation of their asymptotic properties can be the subject of future research.

## 2. ASSUMPTIONS AND MAIN RESULTS

The vectors  $\{\mathbf{x}_t, e_t, -\infty < t < \infty\}$  form a stationary time series. Our main assumption is that the sequence is a Bernoulli shift which can be approximated with finitely dependent time series. Let  $\|\cdot\|$  denote the Euclidean norm of vectors and matrices.

**Assumption 2.1.** *The sequence  $\{\mathbf{x}_t, e_t, -\infty < t < \infty\}$  is a Bernoulli shift, i.e. there are measurable functionals  $\mathbf{g}$  and  $f$  such that  $\mathbf{x}_t = \mathbf{g}(\varepsilon_t, \varepsilon_{t-1}, \dots)$  and  $e_t = f(\varepsilon_t, \varepsilon_{t-1}, \dots)$ , where  $\{\varepsilon_t, -\infty < t < \infty\}$  are independent and identically distributed random variables in some space. Also,*

$$\mathbb{E}e_0 = 0, \mathbb{E}|e_0|^\nu < \infty, \mathbb{E}e_0 x_{0,i} = 0 \text{ and } \mathbb{E}\|x_{0,i}\|^\nu < \infty \text{ with some } \nu > 4, \quad (2.1)$$

and

$$(\mathbb{E}|e_{t,\ell} - e_t|^\nu)^{1/\nu} = O(\ell^{-\alpha}) \text{ and } (\mathbb{E}\|\mathbf{x}_{t,\ell} - \mathbf{x}_t\|^\nu)^{1/\nu} = O(\ell^{-\alpha}) \text{ with some } \alpha > 2, \quad (2.2)$$

where  $e_{t,\ell} = f(\varepsilon_t, \varepsilon_{t-1}, \dots, \varepsilon_{t-\ell}, \varepsilon_{t,\ell,t-\ell-1}, \varepsilon_{t,\ell,t-\ell-2}, \dots)$ ,

$\mathbf{x}_{t,\ell} = \mathbf{g}(\varepsilon_t, \varepsilon_{t-1}, \dots, \varepsilon_{t-\ell}, \varepsilon_{t,\ell,t-\ell-1}, \varepsilon_{t,\ell,t-\ell-2}, \dots)$  and the  $\varepsilon_{i,j,k}$ 's are independent and identically distributed copies of  $\varepsilon_0$ .

The Bernoulli shifts  $\mathbf{x}_{t,m}$  and  $e_{t,m}$  are random variables that closely approximate  $\mathbf{x}_t$  and  $e_t$  in the sense specified in Assumption 2.1. They used to establish some of the theorems that follow. Assumption 2.1 implies immediately that  $e_t, \mathbf{x}_t, -\infty < t < \infty$  is a stationary sequence. For results on change point detection in linear models with nonstationary errors we refer to Hansen (1992), Busetti and Taylor (2004), Harvey et al. (2006), Cavaliere and Taylor (2008) and Kejriwal and Perron (2008).

We prove in Lemma A.1 that

$$\frac{1}{T} \mathbf{X}_T \mathbf{X}_T^\top \rightarrow \mathbf{A} \text{ a.s.}$$

The next assumption postulates that  $\mathbf{A}^{-1}$  exists.

**Assumption 2.2.**  $\mathbf{A}$  is a nonsingular matrix.

Let

$$\sigma^2 = \mathbb{E}e_0^2 + 2 \sum_{\ell=1}^{\infty} \mathbb{E}e_0 e_{\ell}. \quad (2.3)$$

We show in the proof of Lemma A.3 that  $\sigma^2 < \infty$ . To state our main result we need to introduce further notations. The random variables  $\xi_1$  and  $\xi_2$  are double exponential random variables, i.e.

$$\xi_1 \text{ and } \xi_2 \text{ are independent and } \mathbb{P}\{\xi_1 \leq x\} = \mathbb{P}\{\xi_2 \leq x\} = \exp(-e^{-x}) \text{ for all } x, \quad (2.4)$$

and define the numerical sequences

$$a_T = (2 \log \log T)^{1/2} \quad \text{and} \quad b_T = 2 \log \log T + \frac{1}{2} \log \log \log T - \frac{1}{2} \log \pi.$$

**Theorem 2.1.** *If  $H_0$  and Assumptions 2.1 and 2.2 hold, then we have*

$$\left( \frac{D_T}{\sigma}, a_T \frac{H_T}{\sigma} - b_T, a_T \frac{M_T}{\sigma} - 2b_T \right) \xrightarrow{\mathcal{D}} \left( \sup_{0 \leq t \leq 1} |B(t)|, \max(\xi_1, \xi_2), \xi_1 + \xi_2 \right), \quad (2.5)$$

where  $\{B(t), 0 \leq t \leq 1\}$  is a Brownian bridge, independent of  $\xi_1$  and  $\xi_2$  defined by (2.4).

Next we provide the joint asymptotic behaviour  $M_T^*$ ,  $M_T^{(1)}$  and  $M_T^{(2)}$ .

**Theorem 2.2.** *If  $H_0$  and Assumptions 2.1 and 2.2 hold, then we have*

$$\left( \frac{M_T^*}{\sigma}, a_T \frac{M_T^{(1)}}{\sigma} - b_T, a_T \frac{M_T^{(2)}}{\sigma} - b_T \right) \xrightarrow{\mathcal{D}} \left( \bar{B}, \xi_1, \xi_2 \right), \quad (2.6)$$

with

$$\bar{B} = \sup_{0 \leq u_1 \leq u_2 \leq \dots \leq u_m \leq 1} \sum_{j=2}^m |B(u_j) - B(u_{j-1})|,$$

$u_0 = 0$ , where  $\{B(t), 0 \leq t \leq 1\}$  is a Brownian bridge, independent of  $\xi_1$  and  $\xi_2$  defined by (2.4).

We demonstrate via Monte Carlo simulations in Section 3 that the properties of  $D_T$ ,  $H_T$  and  $M_T$  are different under the alternative and the power of these tests depend on the location of the change point(s). Theorem 2.1 makes it possible to combine the three tests to increase the power.

The norming sequences  $a_T$  and  $b_T$  are simple from a theoretical point of view but they are not the best choice in small to moderate sample sizes. Hence we provide an alternative version of (2.5). Let

$$a_{\phi, T} = (2 \log \log [T(\log T)^{\phi}])^{1/2} \quad (2.7)$$

and

$$b_{\phi, T} = 2 \log \log [T(\log T)^{\phi}] + \frac{1}{2} \log \log \log (T(\log T)^{\phi}) - \frac{1}{2} \log \pi, \quad (2.8)$$

where  $-\infty < \phi < \infty$ .

**Theorem 2.3.** *If  $H_0$  and Assumptions 2.1 and 2.2 hold, then we have for all  $-\infty < \phi < \infty$  that*

$$a_{\phi,T} \frac{H_T}{\sigma} - b_{\phi,T} \xrightarrow{\mathcal{D}} \max(\xi_1, \xi_2), \quad (2.9)$$

and

$$a_{\phi,T} \frac{M_T}{\sigma} - 2b_{\phi,T} \xrightarrow{\mathcal{D}} \xi_1 + \xi_2, \quad (2.10)$$

where  $\xi_1$  and  $\xi_2$  are defined in (2.4).

We discuss the choice of  $\phi$  in Section 3.

Next we study the consistency of testing procedures based on Theorem 2.1. Let

$$\tilde{\boldsymbol{\beta}}_T = \sum_{\ell=1}^{m+1} \frac{k_{\ell}^* - k_{\ell-1}^*}{T} \boldsymbol{\beta}^{(\ell)}$$

and

$$J_T = \sqrt{k_1^*} \left| \mathbf{c}^\top (\boldsymbol{\beta}^{(1)} - \tilde{\boldsymbol{\beta}}_T) \right| + \sum_{i=2}^m \frac{k_i^* - k_{i-1}^*}{\sqrt{T}} \left| \mathbf{c}^\top (\boldsymbol{\beta}^{(i)} - \tilde{\boldsymbol{\beta}}_T) \right| + \sqrt{T - k_m^*} \left| \mathbf{c}^\top (\boldsymbol{\beta}^{(m+1)} - \tilde{\boldsymbol{\beta}}_T) \right|,$$

where  $\mathbb{E} \mathbf{x}_0 = \mathbf{c}$ .

**Theorem 2.4.** *We assume that  $H_A$  and Assumptions 2.1 and 2.2 are satisfied.*

(i) *If*

$$J_T \rightarrow \infty, \quad (2.11)$$

then we have that

$$\frac{D_T}{\sigma} \xrightarrow{\mathbb{P}} \infty. \quad (2.12)$$

(ii) *If*

$$(\log \log T)^{-1/2} J_T \rightarrow \infty, \quad (2.13)$$

then we have that

$$(\log \log T)^{-1/2} \frac{H_T}{\sigma} \xrightarrow{\mathbb{P}} \infty. \quad (2.14)$$

and

$$(\log \log T)^{-1/2} \frac{M_T}{\sigma} \xrightarrow{\mathbb{P}} \infty. \quad (2.15)$$

Assumptions (2.11) and (2.13) quantify the relationship between the locations and the sizes of the changes. If the change is early, i.e.  $k_1^*/T \rightarrow 0$ , then the size of the change at  $k_1^*$  should be relatively larger to be detected than if the change occurs in the middle of the data. The same comment holds for a late change, i.e. when  $k_m^*/T \rightarrow 1$ .

The extra  $(\log \log T)^{-1/2}$  term in (2.13) are needed since the variables  $H_T$  and  $M_T$  are increasing to infinity with rate  $(\log \log T)^{1/2}$  under the null hypothesis.

Next we consider two immediate consequences of Theorem 2.4. Let  $\bar{\delta}^{(i)} = \mathbf{c}^\top (\boldsymbol{\beta}^{(i+1)} - \boldsymbol{\beta}^{(i)})$ ,  $1 \leq i \leq m$  denote the size of the change at  $k_i^*$ .

**Corollary 2.1.** *We assume that  $H_A$ , Assumptions 2.1 and 2.2 are satisfied and*

$$\lim_{T \rightarrow \infty} \frac{k_i^*}{T} = \theta_i \quad \text{and} \quad 0 < \theta_1 < \theta_2 < \dots < \theta_m < 1 \quad (2.16)$$

hold.

(i) If

$$T^{1/2} \max_{1 \leq i \leq m} |\bar{\delta}^{(i)}| \rightarrow \infty,$$

then we have (2.12).

(ii) If

$$T^{1/2} (\log \log T)^{-1/2} \max_{1 \leq i \leq m} |\bar{\delta}^{(i)}| \rightarrow \infty,$$

then we have (2.14) and (2.15).

Relation (2.16) means that the change occurs in the “middle” of the data. To illustrate the optimality of our results we consider a special case. We assume that  $m = 1$ , i.e. we have exactly one change and  $\bar{\delta}$  denotes the size of the change.

**Corollary 2.2.** *We assume that  $H_A$  holds with  $m = 1$  and Assumptions 2.1 and 2.2 are satisfied.*

(i) If

$$\left( \frac{k_1^*(T - k_1^*)}{T} \right)^{1/2} |\bar{\delta}| \rightarrow \infty$$

holds, then we have (2.12).

(ii) If

$$\left( \frac{k_1^*(T - k_1^*)}{T} \right)^{1/2} \frac{|\bar{\delta}|}{(\log \log T)^{1/2}} \rightarrow \infty$$

holds, then we have (2.14) and (2.15).

Conditions detailed in Corollary 2.2 are exactly the necessary and sufficient conditions for the consistency of the CUSUM and of the self-normalized CUSUM in case of independent and identically distributed errors (cf. Csörgő and Horváth (1997, p. 170–178)).

According to Corollaries 2.1 and 2.2, we can detect changes if at least one of the changes is larger than  $T^{-1/2}$  or  $((\log \log T)/T)^{1/2}$ , respectively, which also appeared as conditions for the consistency of CUSUM based tests (cf. Csörgő and Horváth (1997), Aue and Horváth (2013) and Horváth and Rice (2014)).

We show in Section A that

$$\sup_{0 \leq u \leq 1} \left\| \frac{1}{T} \sum_{t=1}^{\lfloor Tu \rfloor} \mathbf{x}_t \mathbf{x}_t^\top - u \mathbf{A} \right\| \rightarrow 0 \quad \text{a.s.} \quad (2.17)$$

and

$$\frac{1}{T} \sum_{t=1}^T \|\mathbf{x}_t\|^\nu < \infty \quad \text{a.s. with some } \nu > 2. \quad (2.18)$$

The results of Theorems 2.1–2.4 remain true if we condition on  $\mathbf{x}_t$ ,  $-\infty < t < \infty$  assuming that (2.17) and (2.18) hold and  $\mathbf{x}_t$ ,  $-\infty < t < \infty$  and  $e_t$ ,  $-\infty < t < \infty$  are independent. Assumption (2.17) immediately rules out linear, polynomial, time trend and trigonometric regression. However, in these cases the likelihood method leads to weight functions different



from the square function in the definitions of  $H_T$ ,  $M_1(k_1)$  and  $M_{m+1}(k_m)$ . We refer to Jarušková (1999, 2003) and Albin and Jarušková (2003) for the limit of the maximally selected likelihood ratio with changing trends and to Aue et. al. (2008, 2009, 2012b) for the more general case.

### 3. FINITE SAMPLE PERFORMANCE

**3.1. Estimation of  $\sigma$ .** The long run variance of (2.3) is unknown and must be estimated from the sample. First we consider the case when the errors are uncorrelated, i.e.

**Assumption 3.1.**

$$\mathbb{E}e_t e_s = \begin{cases} 0, & \text{if } t \neq s \\ \sigma^2, & \text{if } t = s. \end{cases}$$

In case of uncorrelated errors we can use the sample variance

$$S_T^2 = \frac{1}{T} \sum_{t=1}^T \hat{e}_t^2. \quad (3.1)$$

**Theorem 3.1.** *We assume that Assumptions 2.1, 2.2 and 3.1 are satisfied.*

(i) *If  $H_0$  holds, then we have that*

$$|S_T^2 - \sigma^2| = O_{\mathbb{P}}(T^{-1/2}). \quad (3.2)$$

(ii) *If  $H_A$  holds, then we have that*

$$S_T^2 = O_{\mathbb{P}}(1). \quad (3.3)$$

It follows immediately from (3.2) and (3.3) that the conclusions of Theorems 2.1–2.4 remain true when  $\sigma$  is replaced with  $S_T$  under Assumption 3.1.

If the errors are correlated we need to use a long run variance kernel estimator

$$\hat{\sigma}_T^2 = \hat{\gamma}_0 + 2 \sum_{\ell=1}^{T-1} K(\ell/h) \hat{\gamma}_\ell, \quad (3.4)$$

where

$$\hat{\gamma}_\ell = \frac{1}{T} \sum_{t=1}^{T-\ell} \hat{e}_t \hat{e}_{t+\ell}$$

denotes the sample correlation of lag  $\ell$  between the residuals. The kernel  $K$  and the window  $h$  satisfy the standard conditions:

**Assumption 3.2.**  $K \geq 0$ ,  $K(0) = 1$ ,  $K(u) = 0$ , if  $|u| > c$  with some  $c > 0$ , and  $K(u)$  is Lipschitz continuous on the real line.

We refer to Taniguchi and Kakizawa (2000) and Politis and Romano (1995) for discussion on the choice of  $K(\cdot)$ .

**Assumption 3.3.**  $h = h(T) \rightarrow \infty$  and  $h/T \rightarrow 0$ .

Parzen (1957) points out that Assumption 3.3 is the necessary condition for the asymptotic consistency of the kernel based long run variance estimator. For the optimal choice of  $h$  we refer to Andrews (1991) and Newey and West (1994). Assumption 3.3 is sufficiently general that it includes the optimal windows specified in these references. The adaptive choice of  $h$  of Politis (2003) can also be used in our set up.

**Theorem 3.2.** *We assume that Assumptions 2.1, 2.2, 3.2 and 3.3 are satisfied.*

(i) *If  $H_0$  holds, then we have that*

$$\hat{\sigma}_T^2 - \sigma^2 = O_{\mathbb{P}} \left( \left( \frac{h}{T} \right)^{1/2} + \frac{1}{h} \right). \quad (3.5)$$

(ii) *If  $H_A$  holds, then we have that*

$$\hat{\sigma}_T^2 = O_{\mathbb{P}}(1) \quad (3.6)$$

Under Assumption 3.3 the convergence of the first coordinate in (2.5),(2.6) and Theorem 2.4(i) remain true when  $\sigma$  is replaced with  $\hat{\sigma}_T$ . If the smoothing parameter  $h$  satisfies  $(h \log \log T)/T \rightarrow 0$  and  $(\log \log T)^{1/2}/h \rightarrow 0$  we can replace the theoretical  $\sigma$  with the estimator  $\hat{\sigma}_T$  in Theorems 2.1, 2.2 and 2.4.

**3.2. Monte Carlo simulations under the null hypothesis.** To assess how well the asymptotic distributions detailed in Theorems 2.1 and 2.3 approximates the finite sample distributions, Monte Carlo simulations are performed. There are several results on the rate of convergence in the functional central limit theorem even in case of dependent variables, we only deal with the choice of the tuning parameter  $\phi$  in Theorem 2.3. The choice of  $\phi$  in (2.9) has been discussed in the literature already (cf. Csörgő and Horváth (1997) and Davis et al. (1995)) so we investigate the finite sample properties of (2.10) when  $\sigma$  is estimated. We consider independent standard normal, GARCH (1,1) and AR(1) errors for various sample sizes. In all cases we investigate, the choice of  $\phi = 1$  gives the best results and therefore only those are reported. In our experiments  $\{\mathbf{x}_t, 1 \leq t \leq T\}$  and  $\{e_t, 1 \leq t \leq T\}$  are independent. Also,  $\boldsymbol{\beta} = (0, 2)^\top$  and  $\mathbf{x}_t = (1, x_{t,2})^\top$ , where  $x_{t,2}, 1 \leq t \leq T$  are independent and identically distributed random variables with  $\mathbb{E}x_{t,2} = 1$  and  $\text{var}(x_{t,2}) = 1$ . The outcomes of the simulations are based on 5,000 repetitions.

**Example 3.1.** First we consider the simplest case when the errors  $\{e_t, 1 \leq t \leq T\}$  are independent and standard normal random variables. Since Assumption (3.1) holds, we used

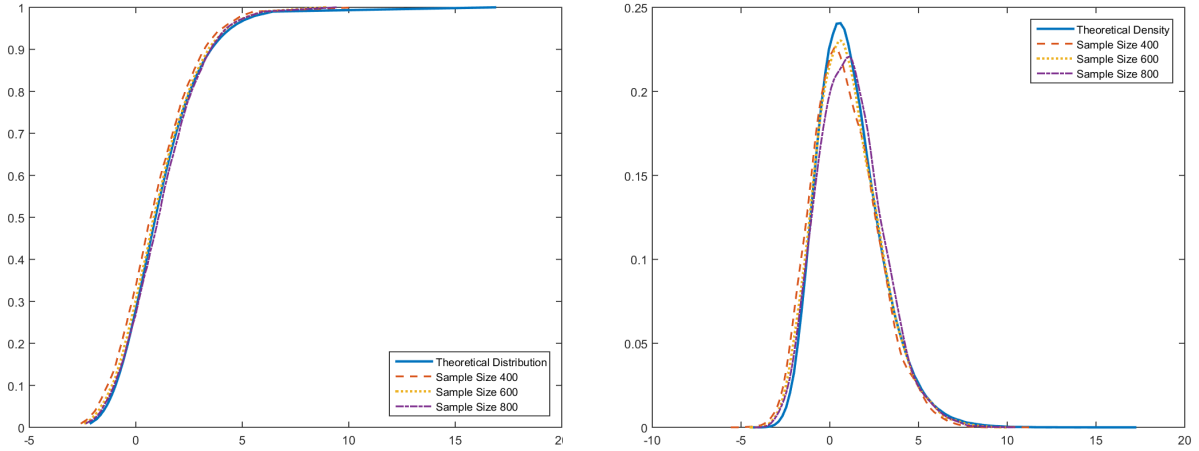
$$V_{T,1}^{(3)} = a_{1,T} \frac{M_T}{S_T} - 2b_{1,T}, \quad (3.7)$$

where  $a_{1,T}$  and  $b_{1,T}$  are defined in (2.7) and (2.8), and  $S_T^2$  is the average of the squared residuals of (3.1). In Figure 3.1 we report the distribution and the density functions of  $V_{T,1}^{(3)}$  for  $T = 400, 600$  and  $800$ . According to Figure 3.1, putting together Theorems 2.3 and 3.1, we obtain a good approximation for the distribution of the test statistic with the choice of  $\phi = 1$ .

**Example 3.2.** In the second study the errors  $e_t$  satisfy a GARCH (1,1) model, i.e.

$$e_t = v_t \varepsilon_t \quad \text{and} \quad v_t^2 = \alpha_0 + \alpha_1 e_{t-1}^2 + \alpha_2 v_{t-1}^2,$$

FIGURE 3.1. Plots of the distribution (left panel) and density functions (right panel) of  $V_{T,1}^{(3)}$  for  $T = 400, 600$  and  $800$  with the distribution and density function of  $\xi_1 + \xi_2$



where the  $\varepsilon_t$ 's are independent and identically distributed standard normal random variables. In our study, we used  $\alpha_0 = .25, \alpha_1 = .25$  and  $\alpha_2 = 0.5$ . For a survey on GARCH and related processes, we refer to Francq and Zakoian (2010). It follows from Aue et al. (2014) that GARCH (1,1) with the present choice of parameters satisfies Assumption 2.1. Since GARCH (1,1) errors satisfy Assumption 3.1, one can use  $V_{T,1}$  as test statistics. However, since

$$\sigma^2 = \frac{\alpha_0}{1 - \alpha_1 - \alpha_2},$$

we can use

$$\tilde{\sigma}_T^2 = \frac{\hat{\alpha}_{T,0}}{1 - \hat{\alpha}_{T,1} - \hat{\alpha}_{T,2}},$$

where  $\hat{\alpha}_{T,0}, \hat{\alpha}_{T,1}$  and  $\hat{\alpha}_{T,2}$  are the quasi maximum likelihood estimators for the GARCH parameters from the residuals  $\hat{e}_t, 1 \leq t \leq T$  of (1.2). Using the basic properties of the quasi maximum likelihood estimators of the GARCH parameters discussed in Francq and Zakoian (2010), one can verify that

$$|\tilde{\sigma}_T^2 - \sigma^2| = O_{\mathbb{P}}(T^{-1/2}).$$

Hence Theorem 2.3 implies

$$V_{T,2}^{(3)} \xrightarrow{\mathcal{D}} \xi_1 + \xi_2,$$

where

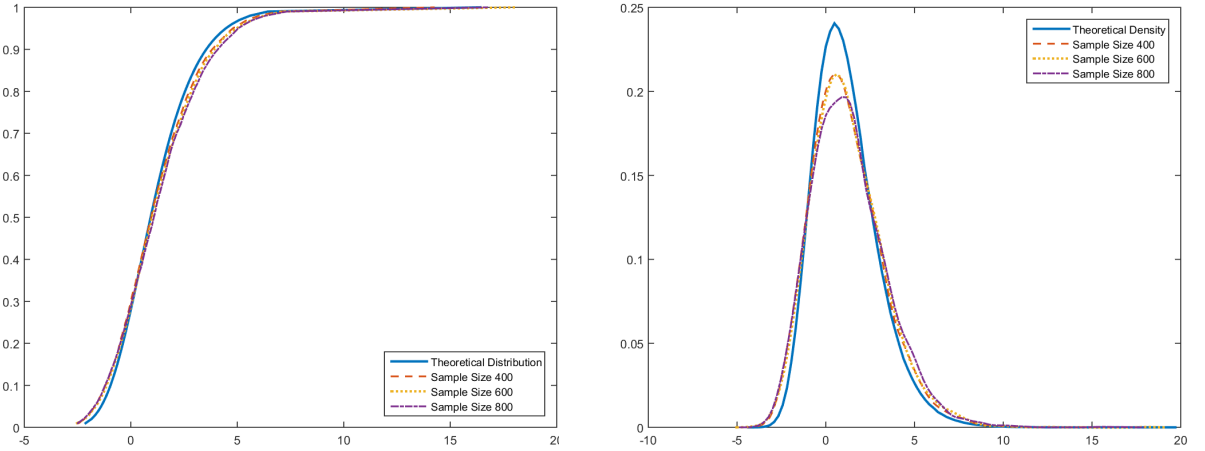
$$V_{T,2}^{(3)} = a_{1,T} \frac{M_T}{\tilde{\sigma}_T} - 2b_{1,T}.$$

The outcome of the Monte Carlo experiment is reported in Figure 3.2. Figure 3.2 shows that the dependence in the GARCH errors causes only minor difference compared to the case of independent and identically distributed  $e_t$ 's.

**Example 3.3.** In our last experiment we simulated AR(1) errors:

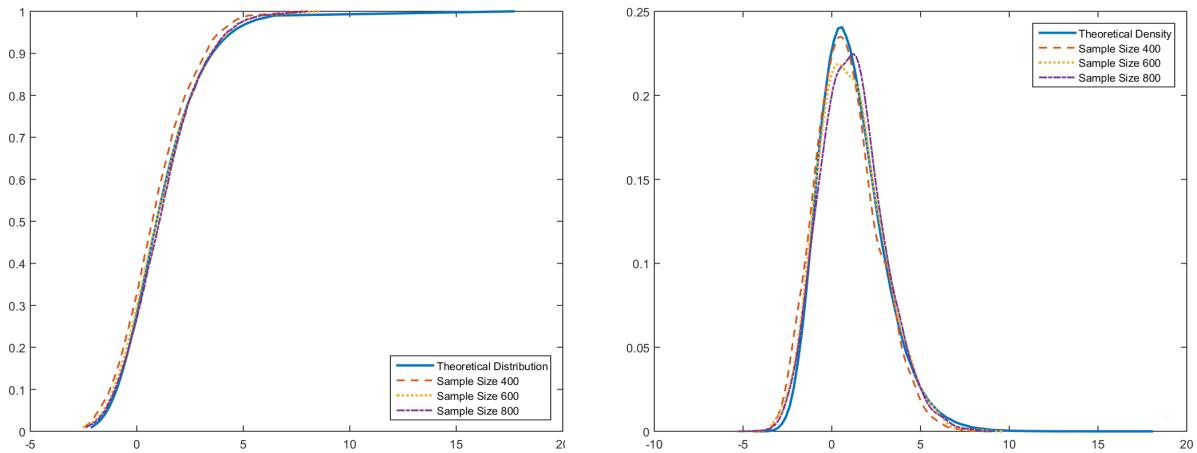
$$e_t = \frac{1}{2}e_{t-1} + \varepsilon_t,$$

FIGURE 3.2. Plots of the distribution (left panel) and density functions (right panel) of  $V_{T,2}^{(3)}$  for  $T = 400, 600$  and  $800$  with the distribution and density function of  $\xi_1 + \xi_2$



where  $\{\varepsilon_t, -\infty < t < \infty\}$  is a sequence of independent standard normal random variables.

FIGURE 3.3. Plots of the distribution (left panel) and density functions (right panel) of  $V_{T,3}^{(3)}$  for  $T = 400, 600$  and  $800$  with the distribution and density function of  $\xi_1 + \xi_2$



We use now

$$V_{T,3}^{(3)} = a_{1,T} \frac{1}{\hat{\sigma}_T} M_T - 2b_{1,T},$$

where  $\hat{\sigma}_T^2$  is the long run variance estimator of (3.4). We used the Bartlett kernel  $K(x) = (1 - |x|)I\{|x| \leq 1\}$  and the window  $h(T) = \lfloor 4(T/100)^{2/9} \rfloor + 1$  following the advise of Andrews and Monahan (1992). Due to the kernel estimation of the long run variance, we need somewhat larger sample sizes to achieve the same empirical accuracy as in the previous experiments and the critical values are slightly underestimated by the limit distribution.

**3.3. Monte Carlo simulations under the alternative hypothesis.** As in the numerical experiments under  $H_0$ , we assume that  $\{\mathbf{x}_t, 1 \leq t \leq T\}$  and  $\{e_t, 1 \leq t \leq T\}$  are independent,  $\mathbf{x}_t = (1, x_{t,2})^\top$ , where  $x_{t,2}, 1 \leq t \leq T$  are independent and identically distributed normal random variables with  $\mathbb{E}x_t = 1$ ,  $\text{var}(x_t) = 1$ . and  $e_1, e_2, \dots$  are independent standard normal random variables. As previously, we used 5,000 repetitions.

We compare our method to the widely used maximally selected CUSUM statistic  $D_T/\sigma$  where  $\sigma$  is defined by (2.3). It is known that under mild conditions (cf. Aue and Horváth (2013)) that

$$\frac{D_T}{\sigma} \xrightarrow{\mathcal{D}} \sup_{0 \leq t \leq 1} |B(t)|, \quad (3.8)$$

where  $B(t), 0 \leq t \leq 1$  denotes a Brownian bridge. Since  $\sigma$  is unknown, we estimate  $\sigma$  with  $S_T$  in case of independent and identically distributed errors resulting in

$$V_{T,1}^{(1)} = \frac{D_T}{S_T},$$

where  $S_T$  is the average of the squared residuals as in Example 3.1. Similarly, in the GARCH (1,1) model of Example 3.2 we use

$$V_{T,2}^{(1)} = \frac{D_T}{\tilde{\sigma}_T},$$

and

$$V_{T,3}^{(1)} = \frac{D_T}{\hat{\sigma}_T}$$

in case of a general stationary model (cf. Example 3.3). It follows from (3.8) and the discussions in Examples 3.1–3.3 that

$$V_{T,i}^{(1)} \xrightarrow{\mathcal{D}} \sup_{0 \leq t \leq 1} |B(t)|, \quad i = 1, 2, 3.$$

Similarly we introduce the statistics for AR(1) sequences

$$V_{T,1}^{(2)} = a_{1,T} \frac{H_T}{S_T} - b_{1,T}, \quad V_{T,2}^{(2)} = a_{1,T} \frac{H_T}{\tilde{\sigma}_T} - b_{1,T} \quad \text{and} \quad V_{T,3}^{(2)} = a_{1,T} \frac{H_T}{\hat{\sigma}_T} - b_{1,T}.$$

By Theorem 2.3 and Section 3.2 we have under  $H_0$  that

$$V_{T,i}^{(2)} \xrightarrow{\mathcal{D}} \max(\xi_1, \xi_2) \quad i = 1, 2, 3.$$

First we consider the case when there is exactly one change in the parameter  $\beta_t$  at  $k_1^*$ .

**Model I.** We assume that  $\mathbf{m} = 1$  and

$$y_t = \begin{cases} \mathbf{x}_t^\top \boldsymbol{\beta}^{(1)} + e_t, & \text{if } 1 \leq t \leq k_1^*, \\ \mathbf{x}_t^\top \boldsymbol{\beta}^{(2)} + e_t, & \text{if } k_1^* + 1 \leq t \leq T. \end{cases}$$

Following Gombay (2010), we selected  $\boldsymbol{\beta}^{(1)} = (0, 1)^\top$  and  $\boldsymbol{\beta}^{(2)} = (0, 1 + \delta)^\top$ , where  $\delta = -2, -1.8, \dots, 1.8, 2$ . We considered three cases for the time of change  $k_1^* = \lfloor T\theta_1 \rfloor$  where  $\theta_1 = .2$  (early change),  $\theta_2 = .5$  (change in the middle) and  $\theta_2 = .9$  (late change).

Figures 3.4–3.6 exhibit the power functions of the statistics  $V_{T,i}^{(j)}, 1 \leq i, j \leq 3$  in Model I. The power of  $V_{T,i}^{(2)}$  and  $V_{T,i}^{(3)}, i = 1, 2, 3$  are essentially the same but the CUSUM statistics

FIGURE 3.4. The power functions of the tests based on  $V_{100,1}^{(1)}$  (upper panel),  $V_{100,1}^{(2)}$  (lower left panel) and  $V_{100,1}^{(3)}$  (lower right panel) in Model I under the conditions of Example 3.1

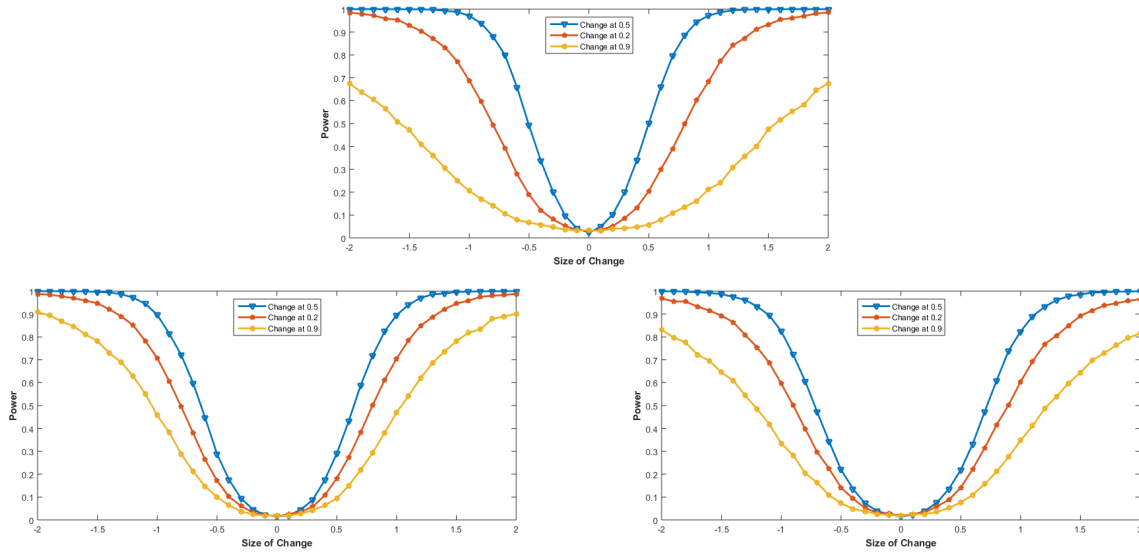
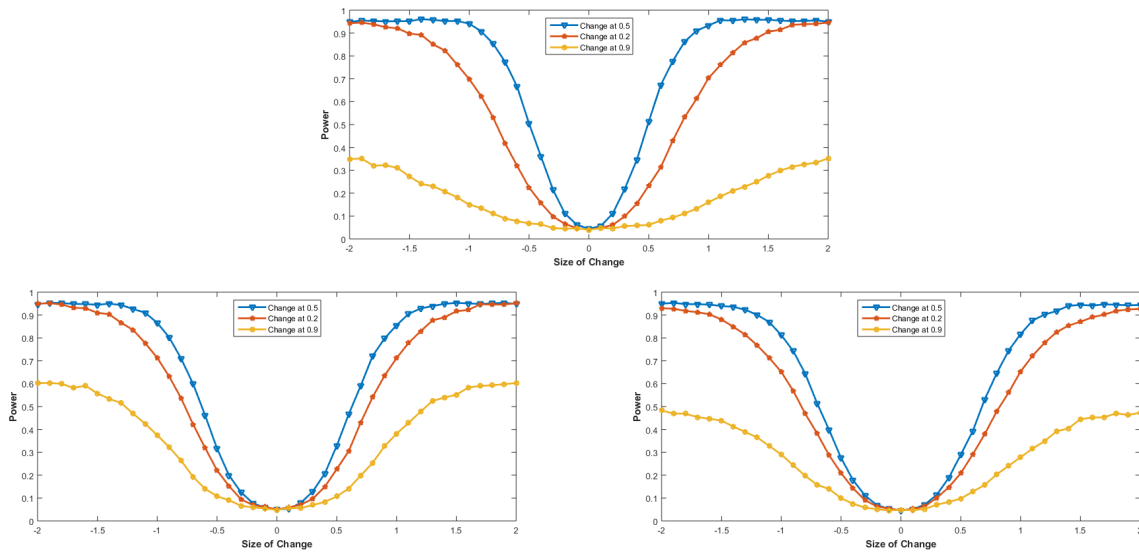


FIGURE 3.5. The power functions of the tests based on  $V_{100,2}^{(1)}$  (upper panel),  $V_{100,2}^{(2)}$  (lower left panel) and  $V_{100,2}^{(3)}$  (lower right panel) in Model I under the conditions of Example 3.2



have higher power when the change occurs in the middle of the data.

FIGURE 3.6. The power functions of the tests based on  $V_{400,3}^{(1)}$  (upper panel),  $V_{400,3}^{(2)}$  (lower left panel) and  $V_{400,3}^{(3)}$  (lower right panel) in Model I under the conditions of Example 3.3

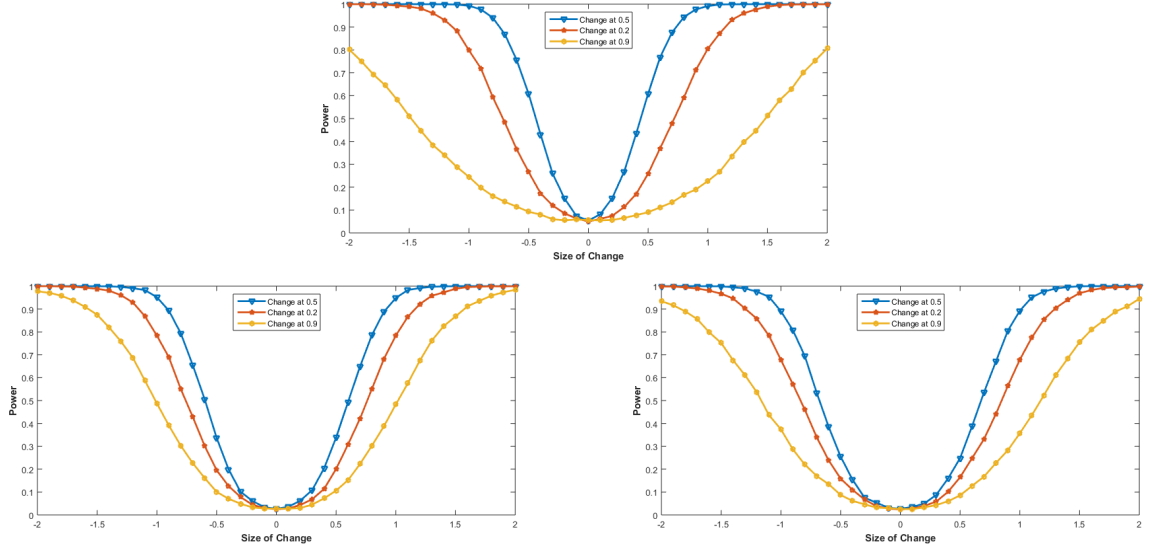
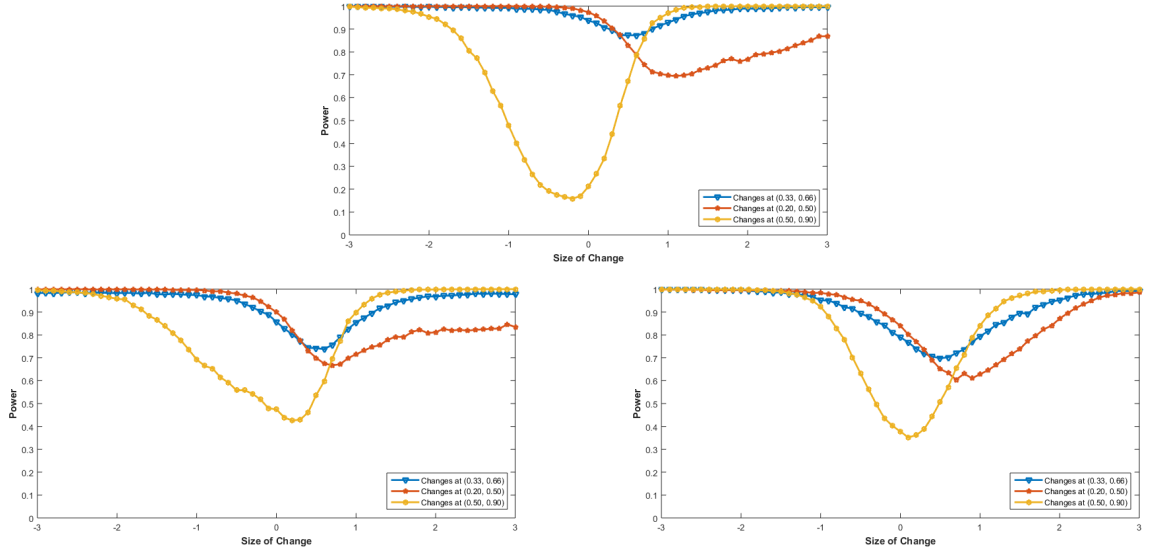


FIGURE 3.7. The power functions of the tests based on  $V_{100,1}^{(1)}$  (upper panel),  $V_{100,1}^{(2)}$  (lower left panel) and  $V_{100,1}^{(3)}$  (lower right panel) in Model II under the conditions of Example 3.1

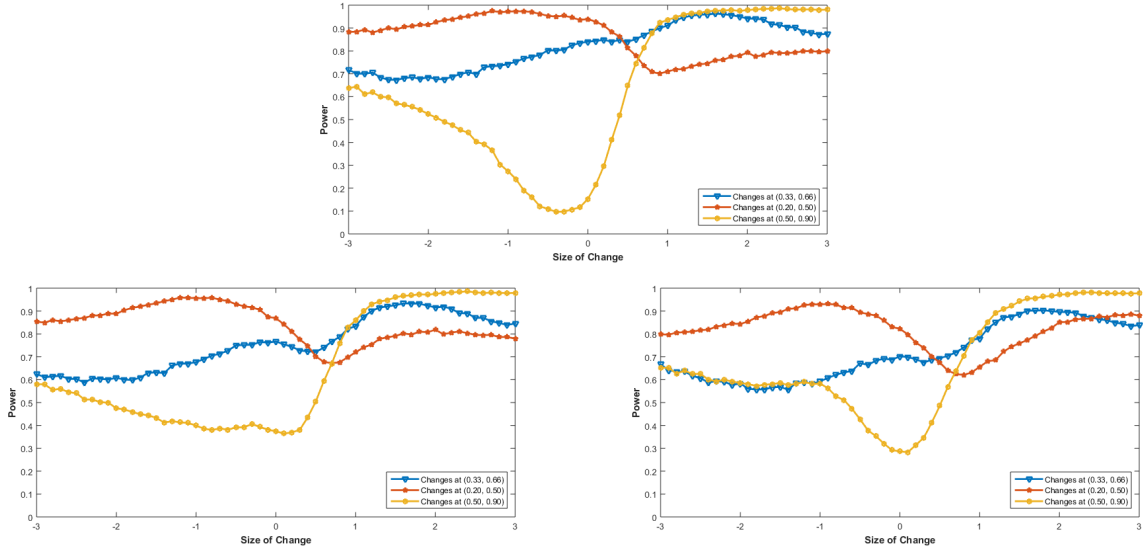


**Model II.** In this case  $m = 2$  and

$$y_t = \begin{cases} \mathbf{x}_t^\top \boldsymbol{\beta}^{(1)} + e_t, & \text{if } 1 \leq t \leq k_1^* \\ \mathbf{x}_t^\top \boldsymbol{\beta}^{(2)} + e_t, & \text{if } k_1^* + 1 \leq t \leq k_2^* \\ \mathbf{x}_t^\top \boldsymbol{\beta}^{(3)} + e_t, & \text{if } k_2^* + 1 \leq t \leq T \end{cases}$$

with  $\beta^{(1)} = (0, 1)^\top$ ,  $\beta^{(2)} = (0, 1 + \delta)^\top$ , where  $\delta = -3, -2.9, \dots, 2.9, 3$  and  $\beta^{(3)} = (0, 2)^\top$ . The times of the changes in Figures 3.7–3.9 are  $k_1^* = \lfloor T\theta_1 \rfloor$  and  $k_2^* = \lfloor T\theta_2 \rfloor$ , when  $(\theta_1, \theta_2) = (.33, .66)$  (blue curves),  $(.2, .5)$  (red curves) and  $(.5, .9)$  (yellow curves). Due to the selection of the parameters, there is at least one change in Model II. As it is expected, the CUSUM statistics  $V_{T,i}^{(1)}$ ,  $i = 1, 2, 3$  have the lowest power nearly in all cases when the size of the change is small. Both  $V_{T,i}^{(2)}$  and  $V_{T,i}^{(3)}$ ,  $i = 1, 2, 3$  have high power and  $V_{T,i}^{(3)}$ ,  $i = 1, 2, 3$  are better when the second change is late.

FIGURE 3.8. The power functions of the tests based on  $V_{100,2}^{(1)}$  (upper panel),  $V_{100,2}^{(2)}$  (lower left panel) and  $V_{100,2}^{(3)}$  (lower right panel) in Model II under the conditions of Example 3.2



#### 4. CHANGE DETECTION IN THE CAPM PARAMETERS

The capital asset pricing model (CAPM) of Sharpe (1964), Lintner (1965) and Merton (1973) and its extensions and modifications have been in the focus of research in applied as well as theoretical finance. In this application we follow Barras et al. (2010) and Fama and French (2010), and use the unconditional four factor CAPM of Carhart (1997) defined as

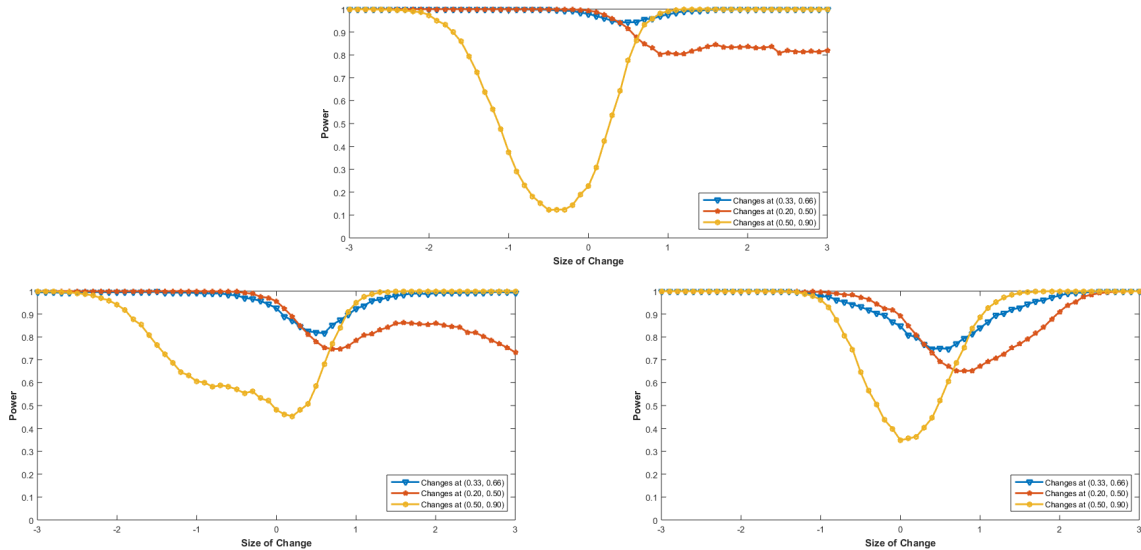
$$R_t - R_t^f = \alpha_t + (R_t^M - R_t^f)\beta_t^M + R_t^{\text{HML}}\beta_t^{\text{HML}} + R_t^{\text{SMB}}\beta_t^{\text{SMB}} + R_t^{\text{MOM}}\beta_t^{\text{MOM}} + e_t, \quad (4.1)$$

$1 \leq t \leq T$ , where  $R_t - R_t^f$  denotes the excess return on the mutual fund,  $R_t^M - R_t^f$  is the excess return on the market portfolio,  $R_t^{\text{HML}}$  refers to the average return on three small portfolios minus the average return on three big portfolios. The value of  $R_t^{\text{SMB}}$  is constructed as the average return on two value portfolios minus the average return on the two growth portfolios and  $R_t^{\text{MOM}}$  gives the returns on a portfolio consisting of stocks with high returns. The monthly return history of US mutual funds is available for the period January 1986 to November 2014 at the web site <http://finance.yahoo.com> and the factors are available at [http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/data\\_library.html](http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/data_library.html). In the notation of model (1.1) we have that

$$\mathbf{x}_t = (1, R_t^M - R_t^f, R_t^{\text{HML}}, R_t^{\text{SMB}}, R_t^{\text{MOM}})^\top \quad (4.2)$$



FIGURE 3.9. The power functions of the tests based on  $V_{400,3}^{(1)}$  (upper panel),  $V_{400,3}^{(2)}$  (lower left panel) and  $V_{100,3}^{(3)}$  (lower right panel) in Model II under the conditions of Example 3.3



and

$$\beta_t = (\alpha_t, \beta_t^M, \beta_t^{\text{HML}}, \beta_t^{\text{SMB}}, \beta_t^{\text{MOM}})^\top. \quad (4.3)$$

Barras et al. (2010) as well as Fama and French (2010) use (4.1) to evaluate the performance of the managers of 2,076 actively managed US mutual funds that existed between 1975 and 2006 assuming that  $\beta_t$  of (4.3) is not time dependent. Fama and French (2010) collect data from January 1984 to December 2006. Baras et al. (2010) provide data from 1975 to 2006. They call a fund manager skilled if the non time dependent  $\alpha$  is positive and unskilled otherwise.

The original CAPMs assume that  $\beta_t$  is constant (not time dependent) which has been criticized by Harvey (1989) and Ferson and Harvey (1993). Jagannathan and Wang (1996), Lettau and Ludvigson (2001) and Beach (2011) advocate time-dependent betas and provide examples where time-varying beta outperform the unconditional CAPM with constant coefficients. On the other hand, Ghysels (1998) argues that a structural break model might be more suitable in applications. Caporale (2012) provides an example of structural breaks in the CAPM betas in the banking sector, finding only three breaks during the period February 1941 and January 2008. For sequential testing of the stability of high-frequency portfolio betas, we refer to Aue et al. (2012a).

If the null hypothesis is rejected, under the assumption that we have two changes, we estimate the times of change by

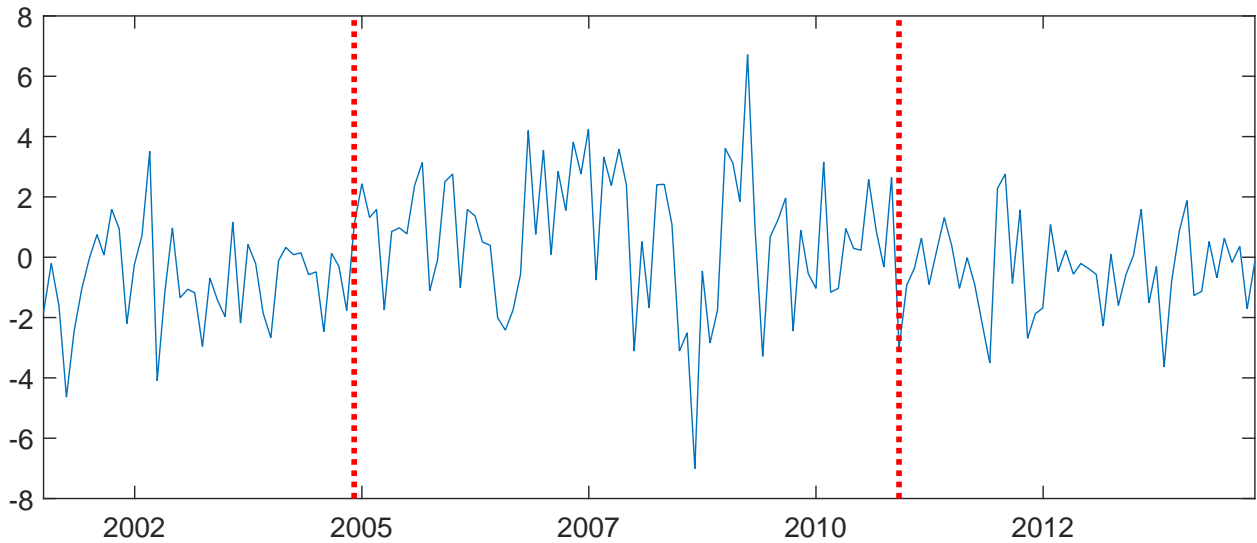
$$(\hat{k}_{1,T}, \hat{k}_{2,T}) = \operatorname{argmax}\{M(k_1, k_2), 1 \leq k_1 < k_2 \leq T\}.$$

If (2.16) holds and  $T^{1/2}\bar{\delta}^{(1)} \rightarrow \infty$  and  $T^{1/2}\bar{\delta}^{(2)} \rightarrow \infty$ , then

$$\frac{1}{T}\hat{k}_{1,T} \xrightarrow{\mathbb{P}} \theta_1 \quad \text{and} \quad \frac{1}{T}\hat{k}_{2,T} \xrightarrow{\mathbb{P}} \theta_2.$$

Then the standardized CUSUM procedure is calculated in sub-samples to confirm there are only two changes. For our study we selected two mutual funds, *American Century Heritage C* (code AHGCX) and *Voya Growth and Income Port I* (code IIVGX). AHGCX seeks long-term capital growth. The fund normally invests in stocks of medium-sized and smaller companies that the adviser believes will increase in value over time. Our procedure found two changes, and the estimated time of changes are December 2004 and December 2010. The standardized CUSUM statistics for five sub-samples are reported in Table C.1 and confirms only two changes. The model estimation result of the segmentation is in Table C.2. In the terminology of Barras et al. (2010), the manager of this fund is “unskilled” in two periods, July 2001 to November 2004 and December 2010 to November 2014, since the portfolio  $\alpha$  are negative for the two periods.

FIGURE 4.1. Plot of the residuals of the AHGCX time series

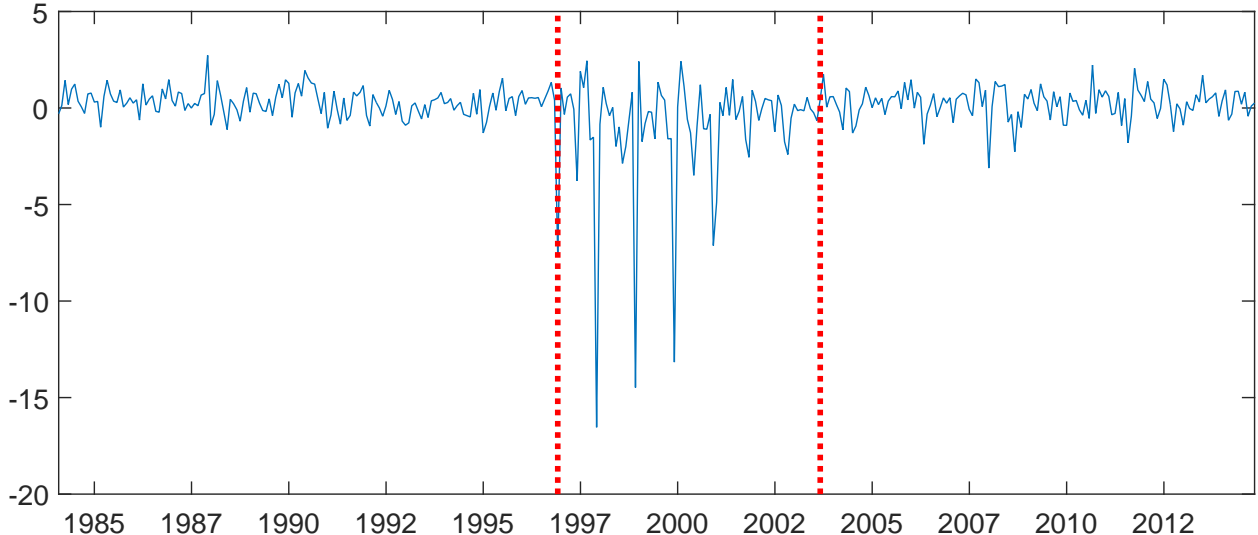


[Tables 4.1 -4.2 Go After Plot AHGCX]

IIVGX’s investment aim is to maximize total return through investments in a diversified portfolio of common stock and securities convertible into common stocks. We plot the residuals for the IIVGX data in Figure 4.2. The estimated times for the changes are December 1996 and September 2003. Table C.3 presents the standardized CUSUM statistics for five sub-samples and confirms that there are only two changes. Table C.4 summarizes the outcome of the segmentation procedure. The portfolio  $\alpha$  is very low in the period between December 1996 to August 2003.

[Tables 4.2- 4.3 Go After Plot of IIVGX]

FIGURE 4.2. Plot of the residuals of the IIVGX time series



## 5. SUMMARY

A new procedure has been developed that can test for an arbitrary but fixed number of changes in parameters of time-dependent regression models. This is achieved by modifying the CUSUM statistic so that it can test for at-most- $m$  changes in this model. The asymptotic properties of our modified statistic are explored under the null hypothesis of no change as well as under the alternative hypothesis. It is documented there that it converges in distribution to a sum of two independent double exponential random variables and that our test is asymptotically consistent under the alternative. Simulations show that our test statistic can detect one change when there is only one change in the parameters and when there are two changes. These simulations also allow comparison of our statistic with the standard CUSUM statistic. Our statistic is further illustrated through application to detecting time-varying risk factors in the capital asset pricing model.

### A. APPENDIX: PROOFS OF THEOREMS 2.1 AND 2.2

**Lemma A.1.** *If Assumption 2.1 holds, then we have that*

$$\frac{1}{T} \mathbf{X}_T \mathbf{X}_T^\top \rightarrow \mathbf{A} \quad a.s.$$

*Proof.* According to the Cauchy-Schwarz inequality and (2.1),  $\mathbb{E}|x_{0,i}x_{0,j}| \leq (\mathbb{E}x_{0,i}^2x_{0,j}^2)^{1/2} < \infty$ ,  $1 \leq i, j \leq d$ . Bernoulli shifts are stationary and ergodic sequences (cf. Stout (1974)) and therefore the ergodic theorem (cf. Breiman (1968)) yields

$$\frac{1}{T} \sum_{t=1}^T x_{t,i}x_{t,j} \longrightarrow \mathbb{E}x_{0,i}x_{0,j}. \quad (\text{A.1})$$

□

Under the null hypothesis (1.1) reduces to

$$y_t = \mathbf{x}_t^\top \boldsymbol{\beta} + e_t, \quad 1 \leq t \leq T,$$

where  $\boldsymbol{\beta}$  denotes the common regressor.

**Lemma A.2.** *If  $H_0$ , Assumptions 2.1 and 2.2 hold, then we have that*

$$\hat{\boldsymbol{\beta}}_T - \boldsymbol{\beta} = O_{\mathbb{P}}(T^{-1/2})$$

*Proof.* Since under  $H_0$

$$\hat{\boldsymbol{\beta}}_T - \boldsymbol{\beta} = (\mathbf{X}_T^\top \mathbf{X}_T)^{-1} \mathbf{X}_T^\top \mathbf{E}_T,$$

where  $\mathbf{E}_T = [e_1, e_2, \dots, e_T]^\top$ . Thus we get via Lemma A.1 that

$$\|\hat{\boldsymbol{\beta}}_T - \boldsymbol{\beta}\| = O_{\mathbb{P}}(1/T) \|\mathbf{X}_T^\top \mathbf{E}_T\|.$$

Since  $\mathbb{E}x_{t,i}e_t = 0$  is assumed in (2.1), it is enough to show that

$$E\|\mathbf{X}_T^\top \mathbf{E}_T\|^2 = O(T). \tag{A.2}$$

It follows from stationarity that

$$\begin{aligned} \mathbb{E} \left( \sum_{t=1}^T x_{t,i} e_t \right)^2 &\leq \sum_{t=1}^T \sum_{s=1}^T |\mathbb{E}x_{t,i} e_t x_{s,i} e_s| \\ &= T \mathbb{E}(x_{t,0} e_0)^2 + 2 \sum_{1 \leq t < s \leq T} |\mathbb{E}x_{t,i} e_t x_{s,i} e_s| \\ &= T \mathbb{E}(x_{t,0} e_0)^2 + 2 \sum_{u=1}^{T-1} (T-u) |\mathbb{E}x_{0,i} e_0 x_{u,i} e_u| \\ &\leq T \sum_{u=0}^{\infty} |\mathbb{E}x_{0,i} e_0 x_{u,i} e_u|. \end{aligned}$$

Using the notation  $\mathbf{x}_{t,m} = [x_{t,m,1}, x_{t,m,2}, \dots, x_{t,m,d}]$ , we write

$$\mathbb{E}x_{0,i} e_0 x_{u,i} e_u = \mathbb{E}x_{0,i} e_0 (x_{u,i} e_u - x_{u,u,i} e_{u,u})$$

since by independence  $\mathbb{E}x_{0,i}e_0x_{u,u,i}e_{u,u} = 0$ . Hence by the Cauchy–Schwarz inequality we conclude

$$\begin{aligned}
\sum_{u=1}^{\infty} |\mathbb{E}x_{0,i}e_0x_{u,i}e_u| &= \sum_{u=1}^{\infty} |\mathbb{E}x_{0,i}e_0(x_{u,i}e_u - x_{u,u,i}e_{u,u})| \\
&\leq (\mathbb{E}x_{0,i}^4)^{1/4} (\mathbb{E}e_0^4)^{1/4} \sum_{u=1}^{\infty} (\mathbb{E}(x_{u,i}e_u - x_{u,u,i}e_{u,u})^2)^{1/2} \\
&= (\mathbb{E}x_{0,i}^4)^{1/4} (\mathbb{E}e_0^4)^{1/4} \sum_{u=1}^{\infty} (\mathbb{E}(x_{0,i}e_0 - x_{0,u,i}e_{0,u})^2)^{1/2} \\
&\leq 2(\mathbb{E}x_{0,i}^4)^{1/4} (\mathbb{E}e_i^4)^{1/4} \sum_{u=1}^{\infty} (\mathbb{E}(x_{0,i} - x_{0,u,i})e_0)^2 + \mathbb{E}(x_{0,u,i}(e_0 - e_{0,u})^2)^{1/2} \\
&\leq 2(\mathbb{E}x_{0,i}^4)^{1/4} (\mathbb{E}e_0^4)^{1/2} \sum_{u=1}^{\infty} (\mathbb{E}(x_{0,i} - x_{0,u,i})^4)^{1/4} \\
&\quad + 2(\mathbb{E}x_{0,i}^4)^{1/2} (\mathbb{E}e_0^4)^{1/4} \sum_{u=1}^{\infty} (\mathbb{E}(e_0 - e_{0,u})^4)^{1/4},
\end{aligned}$$

and therefore (A.2) follows from (2.2).  $\square$

Since under  $H_0$  we have

$$\hat{e}_t = e_t - \mathbf{x}_t^\top (\hat{\boldsymbol{\beta}}_T - \boldsymbol{\beta}), \quad (\text{A.3})$$

we can decompose  $M_1(k)$  as

$$M_1(k) = \frac{1}{\sqrt{k}} \left( \sum_{t=1}^k e_t + R_k \right), \quad 1 \leq k \leq T, \quad (\text{A.4})$$

where

$$R_k = R_{k,1} + R_{k,2} + R_{k,3}, \quad 1 \leq k \leq T,$$

with

$$R_{k,1} = - \left( \sum_{t=1}^k \mathbf{x}_t \right)^\top (\hat{\boldsymbol{\beta}}_T - \boldsymbol{\beta}), \quad R_{k,2} = -\frac{k}{T} \sum_{t=1}^T e_t \quad \text{and} \quad R_{k,3} = \frac{k}{T} \left( \sum_{t=1}^T \mathbf{x}_t \right)^\top (\hat{\boldsymbol{\beta}}_T - \boldsymbol{\beta}).$$

**Lemma A.3.** *If  $H_0$ , Assumptions 2.1 and 2.2 hold, then we have that*

$$\max_{1 \leq k \leq T} k^{-1/2} |R_k| = O_{\mathbb{P}}(1).$$

*Proof.* By Assumption 2.1,  $\mathbf{x}_t$  is stationary and ergodic, and therefore by the ergodic theorem (cf. Breiman (1968)) we have

$$\max_{1 \leq k \leq T} \frac{1}{k} \left| \sum_{t=1}^k \mathbf{x}_t \right| = O_{\mathbb{P}}(1).$$

Hence Lemma A.2 yields

$$\max_{1 \leq k \leq T} k^{-1/2} |R_{k,1}| = O_{\mathbb{P}}(1).$$

Following the arguments in the proof of Lemma A.2 we get that

$$\mathbb{E} \left( \sum_{t=1}^T e_t \right) \leq 2T \sum_{s=0}^T |\mathbb{E} e_0 e_s| < \infty$$

and since  $\mathbb{E} e_t = 0$  we conclude

$$\left| \sum_{t=1}^T e_t \right| = O_{\mathbb{P}}(T^{1/2}). \quad (\text{A.5})$$

Hence

$$\max_{1 \leq k \leq T} k^{-1/2} |R_{k,2}| = O_{\mathbb{P}}(1).$$

By the ergodic theorem we have that

$$\frac{1}{T} \sum_{t=1}^T |x_{t,i}| = O_{\mathbb{P}}(1). \quad (\text{A.6})$$

Applying Lemma A.2 we get that

$$\max_{1 \leq k \leq T} k^{-1/2} |R_{k,3}| = \max_{1 \leq k \leq T} \left| \frac{\sqrt{k}}{T} \left( \sum_{t=1}^T \mathbf{x}_t \right)^{\top} (\hat{\boldsymbol{\beta}}_T - \boldsymbol{\beta}) \right| = O_{\mathbb{P}}(1),$$

completing the proof of lemma A.3.  $\square$

**Lemma A.4.** *If Assumptions 2.1 and 2.2 hold, then we have that*

$$(2 \log \log T)^{-1/2} \max_{1 \leq k \leq T} \frac{1}{\sqrt{k}} \left| \sum_{t=1}^k e_t \right| \xrightarrow{\mathbb{P}} \sigma,$$

where  $\sigma$  is defined in (2.3).

*Proof.* It follows from Lemma 5.4 of Aue et al. (2014), that we can define Wiener processes  $W_T$  such that

$$\max_{1 \leq k \leq T} k^{-1/2+\delta} \left| \sum_{t=1}^k e_t - \sigma W_T(k) \right| = O_{\mathbb{P}}(1) \text{ with some } \delta > 0. \quad (\text{A.7})$$

Hence for any  $1 \leq c_1 = c_1(T) \leq c_2 = c_2(T) \leq T$  we have

$$\max_{c_1 \leq k \leq c_2} \frac{1}{\sqrt{k}} \left| \sum_{t=1}^k e_t \right| - \max_{c_1 \leq k \leq c_2} \frac{\sigma}{\sqrt{k}} |W_T(k)| = O_{\mathbb{P}}(c_1^{-\delta}). \quad (\text{A.8})$$

Since the distribution of  $W_T(\cdot)$  does not depend on  $T$ , by the law of the iterated logarithm for Wiener processes we get

$$(2 \log \log T)^{-1/2} \max_{1 \leq k \leq T} \frac{1}{\sqrt{k}} |W_T(k)| \xrightarrow{\mathbb{P}} 1,$$

which completes the proof.  $\square$

**Lemma A.5.** *If  $H_0$ , Assumptions 2.1 and 2.2 hold, then we have that*

$$\mathbb{P} \left( \max_{1 \leq k \leq T} |M_1(k)| = \max_{n(T) \leq k \leq m(T)} |M_1(k)| \right) \xrightarrow{\mathbb{P}} 1,$$

where  $n(T) = (\log T)^\kappa$ , and  $m(T) = T/(\log T)^\kappa$  with any  $\kappa > 0$ .

*Proof.* It follows from Lemmas A.3 and A.4 that

$$(2 \log \log T)^{-1/2} \max_{1 \leq k \leq T} |M_1(k)| \xrightarrow{\mathbb{P}} \sigma, \quad (\text{A.9})$$

and therefore we need to show only that

$$\max_{1 \leq k \leq n(T)} |M_1(k)| = o_{\mathbb{P}}((\log \log T)^{1/2})$$

and

$$\max_{m(T) \leq k \leq T} |M_1(k)| = o_{\mathbb{P}}((\log \log T)^{1/2}).$$

On account of Lemma A.3 we need to prove only that

$$\max_{1 \leq k \leq n(T)} \frac{1}{\sqrt{k}} \left| \sum_{t=1}^k e_t \right| = o_{\mathbb{P}}((\log \log T)^{1/2}) \quad (\text{A.10})$$

and

$$\max_{m(T) \leq k \leq T} \frac{1}{\sqrt{k}} \left| \sum_{t=1}^k e_t \right| = o_{\mathbb{P}}((\log \log T)^{1/2}). \quad (\text{A.11})$$

For any Wiener process  $W(\cdot)$  we have that

$$\max_{1 \leq k \leq n(T)} \frac{1}{\sqrt{k}} |W(k)| = O_{\mathbb{P}}((\log \log \log T)^{1/2}) \quad (\text{A.12})$$

and

$$\max_{m(T) \leq k \leq T} \frac{1}{\sqrt{k}} |W(k)| = O_{\mathbb{P}}((\log \log \log T)^{1/2}) \quad (\text{A.13})$$

(cf. Csörgő and Horváth (1997)). The claims in (A.10) and (A.11) follow immediately from (A.8) and (A.12), (A.13).  $\square$

According to Lemma A.5,  $|M_1(k)|$  reaches its largest value on the interval  $n(T), m(T)$  with probability converging to 1.

**Lemma A.6.** *If  $H_0$ , Assumptions 2.1 and 2.2 hold, then we have that*

$$\max_{n(T) \leq k \leq m(T)} |M_1(k)| - \max_{n(T) \leq k \leq m(T)} \frac{1}{\sqrt{k}} \left| \sum_{t=1}^k e_t \right| = o_{\mathbb{P}}((\log \log T)^{-1/2}),$$

where  $n(T)$  and  $m(T)$  are defined in Lemma A.5.

*Proof.* We use again (A.4). Combining Lemma A.2 and (A.1) we get that

$$\max_{n(T) \leq k \leq m(T)} k^{-1/2} |R_{k,1}| = O_{\mathbb{P}}(1) \max_{n(T) \leq k \leq m(T)} k^{1/2} T^{-1/2} = O_{\mathbb{P}}(1) (\log T)^{-\kappa/2}$$

and similarly

$$\max_{n(T) \leq k \leq m(T)} k^{-1/2} |R_{k,3}| = O_{\mathbb{P}}(1) (\log T)^{-\kappa/2}$$

It follows from (A.5) that

$$\max_{n(T) \leq k \leq m(T)} k^{-1/2} |R_{k,2}| = O_{\mathbb{P}}(1) T^{-1/2} m^{1/2}(T) = O_{\mathbb{P}}(1) (\log T)^{-\kappa/2},$$

which concludes the proof.  $\square$

Next we consider  $M_{m+1}(\ell)$ , since its definition is similar to that of  $M_1(k)$ . However, due to time reversal, i.e. the CUSUM starts with residual  $\hat{e}_T$ , we need to modify Lemmas A.3–A.6. As in the decomposition of  $M_1(k)$ , we have

$$M_{m+1}(\ell) = \frac{1}{\sqrt{T-\ell}} \left( \sum_{t=\ell}^T e_t + R_\ell^* \right), \quad 1\ell < T,$$

where

$$R_\ell^* = R_{\ell,1}^* + R_{\ell,2}^* + R_{\ell,3}^*, \quad 1\ell < T,$$

with

$$R_{\ell,1}^* = - \left( \sum_{t=\ell}^T \mathbf{x}_t \right)^\top (\hat{\boldsymbol{\beta}}_T - \boldsymbol{\beta}), \quad R_{\ell,2}^* = -\frac{T-\ell}{T} \sum_{t=1}^T e_t$$

and

$$R_{\ell,3}^* = \frac{T-\ell}{T} \left( \sum_{t=1}^T \mathbf{x}_t \right)^\top (\hat{\boldsymbol{\beta}}_T - \boldsymbol{\beta}).$$

**Lemma A.7.** *If  $H_0$ , Assumptions 2.1 and 2.2 hold, then we have that*

$$\max_{1 \leq \ell < T} \frac{|R_\ell^*|}{\sqrt{T-\ell}} = O_{\mathbb{P}}(1).$$

*Proof.* First we show that

$$\max_{1 \leq \ell < T} \frac{1}{T-\ell} \left| \sum_{t=\ell}^T x_{t,i} \right| = O_{\mathbb{P}}(1). \quad (\text{A.14})$$

Using Aue et al. (2014), we can find Wiener processes  $W_{T,i}(\cdot)$  such that

$$\max_{1 \leq \ell < T} \frac{1}{\sqrt{T-\ell}} \left| \sum_{t=\ell}^T (x_{t,i} - \mathbb{E}x_{t,i}) \right| - \max_{1 \leq \ell < T} \left( \frac{\text{var}(x_{0,i})}{T-\ell} \right)^{1/2} \left| W_{T,i}(T-\ell) \right| = O_{\mathbb{P}}(1),$$

and therefore (A.14) follows from the law of the iterated logarithm for Wiener processes. Hence Lemma A.2 yields

$$\max_{1 \leq \ell < T} (T-\ell)^{-1/2} |R_{\ell,1}^*| = O_{\mathbb{P}}(1) \quad \text{and} \quad \max_{1 \leq \ell < T} (T-\ell)^{-1/2} |R_{\ell,3}^*| = O_{\mathbb{P}}(1).$$

Using (A.5) we get immediately that

$$\max_{1 \leq \ell < T} (T-\ell)^{-1/2} |R_{\ell,2}^*| = O_{\mathbb{P}}(1).$$

□

We continue with the analogue of Lemma A.4.

**Lemma A.8.** *If Assumptions 2.1 and 2.2 hold, then we have that*

$$(2 \log \log T)^{-1/2} \max_{1 \leq \ell < T} \frac{1}{\sqrt{T-\ell}} \left| \sum_{t=\ell}^T e_t \right| \xrightarrow{\mathbb{P}} \sigma,$$

where  $\sigma$  is defined in (2.3).



*Proof.* Using again Lemma 5.4 of Aue et al. (2014), we can define Wiener processes  $W_T^*$  such that

$$\max_{1 \leq \ell < T} (T - \ell)^{-1/2+\delta} \left| \sum_{t=\ell}^T e_t - \sigma W_T^*(T - \ell) \right| = O_{\mathbb{P}}(1) \text{ with some } \delta > 0. \quad (\text{A.15})$$

Change of variable and the Darling–Erdős law (cf. Csörgő and Horváth (1997)) gives

$$(2 \log \log T)^{-1/2} \max_{1 \leq \ell < T} \frac{1}{\sqrt{T - \ell}} |W_T^*(T - \ell)| \xrightarrow{\mathbb{P}} 1,$$

and therefore Lemma A.8 follows from (A.15) along the lines of (A.7) and (A.8).  $\square$

**Lemma A.9.** *If  $H_0$ , Assumptions 2.1 and 2.2 hold, then we have that*

$$\mathbb{P} \left( \max_{1 \leq \ell < T} |M_{\mathfrak{m}+1}(\ell)| = \max_{T-m(T) \leq \ell \leq T-n(T)} |M_{\mathfrak{m}+1}(\ell)| \right) \xrightarrow{\mathbb{P}} 1,$$

where  $n(T)$  and  $m(T)$  are defined in Lemma A.5.

*Proof.* It follows from Lemmas A.7 and A.8 that

$$(2 \log \log T)^{-1/2} \max_{1 \leq \ell < T} (T - \ell)^{-1/2} |M_{\mathfrak{m}+1}(\ell)| \xrightarrow{\mathbb{P}} \sigma.$$

Thus, in light of Lemmas A.7 and A.8, it is enough to establish that

$$\max_{1 \leq \ell \leq T-m(T)} (T - \ell)^{-1/2} \left| \sum_{t=\ell}^T e_t \right| = o_{\mathbb{P}}((\log \log T)^{1/2}) \quad (\text{A.16})$$

and

$$\max_{T-n(T) \leq \ell < T} (T - \ell)^{-1/2} \left| \sum_{t=\ell}^T e_t \right| = o_{\mathbb{P}}((\log \log T)^{1/2}). \quad (\text{A.17})$$

Using (A.15) we get for any  $1 \leq c_1 = c_1(T) \leq c_2 = c_2(T) < T$

$$\max_{T-c_2 \leq \ell \leq T-c_1} \frac{1}{\sqrt{T - \ell}} \left| \sum_{t=\ell}^T e_t \right| - \max_{T-c_2 \leq \ell \leq T-c_1} \frac{\sigma}{\sqrt{T - \ell}} |W_T^*(T - \ell)| = O_{\mathbb{P}}(c_1^{-\delta}),$$

and therefore (A.12) and (A.13) imply (A.16) and (A.17).  $\square$

**Lemma A.10.** *If  $H_0$ , Assumptions 2.1 and 2.2 hold, then we have that*

$$\max_{T-m(T) \leq \ell \leq T-n(T)} |M_{\mathfrak{m}+1}(\ell)| - \max_{T-m(T) \leq \ell \leq T-n(T)} \frac{1}{\sqrt{T - \ell}} \left| \sum_{t=\ell}^T e_t \right| = o_{\mathbb{P}}((\log \log T)^{-1/2}),$$

where  $n(T)$  and  $m(T)$  are defined in Lemma A.5.

*Proof.* Putting together Lemma A.2 and (A.14) we conclude

$$\max_{T-m(T) \leq \ell \leq T-n(T)} (T - \ell)^{-1/2} |R_{\ell,1}^*| = O_{\mathbb{P}}(1)(\log T)^{-\kappa/2}.$$

Similar arguments give

$$\max_{T-m(T) \leq \ell \leq T-n(T)} (T - \ell)^{-1/2} |R_{\ell,2}^*| = O_{\mathbb{P}}(1)(\log T)^{-\kappa/2}$$

and

$$\max_{T-m(T) \leq \ell \leq T-n(T)} (T - \ell)^{-1/2} |R_{\ell,3}^*| = O_{\mathbb{P}}(1)(\log T)^{-\kappa/2}.$$

□

**Lemma A.11.** *If  $H_0$ , Assumptions 2.1 and 2.2 hold, then we have that*

$$\max_{1 \leq k \leq \ell \leq T} |M_j(k, \ell)| = O_{\mathbb{P}}(1), \quad 2 \leq j \leq m.$$

*Proof.* Using (A.3) we write

$$M_j(k, \ell) = R_{k,\ell,1} + \dots + R_{k,\ell,4},$$

where

$$R_{k,\ell,1} = T^{-1/2} \sum_{t=k}^{\ell} e_t, \quad R_{k,\ell,2} = -T^{-1/2} \left( \sum_{t=k}^{\ell} \mathbf{x}_t \right)^{\top} (\hat{\boldsymbol{\beta}}_T - \boldsymbol{\beta}), \quad R_{k,\ell,3} = -\frac{\ell - k}{T^{3/2}} \sum_{t=1}^T e_t$$

and

$$R_{k,\ell,4} = \frac{\ell - k}{T^{3/2}} \left( \sum_{t=1}^T \mathbf{x}_t \right)^{\top} (\hat{\boldsymbol{\beta}}_T - \boldsymbol{\beta}).$$

It follows from Lemma 5.4 of Aue et al. (2014) that

$$T^{-1/2} \sum_{t=1}^{Tu} e_t \xrightarrow{\mathcal{D}[0,1]} \sigma W(t),$$

where  $W(\cdot)$  is a Wiener process. Hence

$$\sup_{0 \leq u, v \leq 1} T^{-1/2} \left| \sum_{t=Tu}^{Tv} e_t \right| \xrightarrow{\mathcal{D}} \sigma \sup_{0 \leq u, v \leq 1} |W(u) - W(v)|,$$

which implies

$$\max_{1 \leq k \leq \ell \leq T} |R_{k,\ell,1}| = O_{\mathbb{P}}(1).$$

Applying Lemma A.2 and the ergodic theorem we conclude

$$\max_{1 \leq k \leq \ell \leq T} |R_{k,\ell,2}| = O_{\mathbb{P}}(1) \frac{1}{T} \max_{1 \leq k \leq \ell \leq T} \left\| \sum_{t=k}^{\ell} \mathbf{x}_t \right\| = O_{\mathbb{P}}(1) \frac{1}{T} \max_{1 \leq k \leq \ell \leq T} \sum_{t=k}^{\ell} \|\mathbf{x}_t\| = O_{\mathbb{P}}(1).$$

Similar arguments yield

$$\max_{1 \leq k \leq \ell \leq T} |R_{k,\ell,3}| = O_{\mathbb{P}}(1) \quad \text{and} \quad \max_{1 \leq k \leq \ell \leq T} |R_{k,\ell,4}| = O_{\mathbb{P}}(1).$$

□

Let us define  $A_T = [n(T), m(T)] \times [T - m(T), T - n(T)]$ , where  $n(T)$  and  $m(T)$  are defined in Lemma A.5.

**Lemma A.12.** *If  $H_0$ , Assumptions 2.1 and 2.2 hold, then we have that*

$$\max_{(k,\ell) \in A_T} |M_j(k, \ell)| = o_{\mathbb{P}}((\log \log T)^{-1/2}), \quad 2 \leq j \leq m.$$

*Proof.* On account of (A.3) we can decompose  $M_j$  as

$$M_j(k, \ell) = M_{2,1}(k) + M_{2,2}(k) + M_{2,3}(\ell) + M_{2,4}(\ell),$$

where

$$M_{2,1}(k) = -T^{-1/2} \sum_{t=1}^{k-1} \hat{e}_t, \quad M_{2,2}(k) = \frac{k}{T^{3/2}} \sum_{t=1}^T \hat{e}_t, \quad M_{2,3}(\ell) = -T^{-1/2} \sum_{t=\ell+1}^T \hat{e}_t,$$

and

$$M_{2,4}(\ell) = \frac{T-\ell}{T^{3/2}} \sum_{t=1}^T \hat{e}_t.$$

We note that by (A.3) that

$$\sum_{t=u}^v \hat{e}_t = \sum_{t=u}^v e_t - \left( \sum_{t=u}^v \mathbf{x}_t \right)^\top (\hat{\beta}_T - \beta) = O_{\mathbb{P}}(T^{1/2}).$$

Thus we obtain that from Lemma A.2, (A.5) and (A.6) that

$$\max_{n(T) \leq k \leq m(T)} |M_{2,2}(k)| = O_{\mathbb{P}}((\log T)^{-\kappa/2}) \quad \text{and} \quad \max_{T-m(T) \leq \ell \leq T-n(T)} |M_{2,4}(\ell)| = O_{\mathbb{P}}((\log T)^{-\kappa/2}).$$

Applying (A.7) we get

$$\begin{aligned} & T^{-1/2} \max_{n(T) \leq k \leq m(T)} \left| \sum_{t=1}^k e_t \right| \tag{A.18} \\ & \leq T^{-1/2} \max_{n(T) \leq k \leq m(T)} \sigma |W_T(k)| + T^{-1/2} \max_{n(T) \leq k \leq m(T)} \left| \sum_{t=1}^k e_t - \sigma W_T(k) \right| \\ & = O_{\mathbb{P}}(1) T^{-1/2} \max_{n(T) \leq k \leq m(T)} k^{1/2-\delta} + T^{-1/2} \max_{n(T) \leq k \leq m(T)} \sigma |W_T(k)| \\ & = O_{\mathbb{P}}(1) \left( T^{-1/2} m^{1/2-\delta}(T) + (\log T)^{-\kappa/2} \right) \\ & = o_{\mathbb{P}} \left( (\log \log T)^{-1/2} \right), \end{aligned}$$

since by the scale transformation of the Wiener process  $W(\cdot)$  we have

$$\sup_{0 \leq u \leq m(T)} |W(t)| \stackrel{\mathcal{D}}{=} m^{1/2}(T) \sup_{0 \leq v \leq 1} |W(v)|.$$

By the ergodic theorem and Lemma A.2 we obtain that

$$T^{-1/2} \max_{n(T) \leq k \leq m(T)} \left| \left( \sum_{t=1}^{k-1} \mathbf{x}_t \right)^\top (\hat{\beta}_T - \beta) \right| = O_{\mathbb{P}}(1),$$

and therefore (A.18) implies

$$\max_{n(T) \leq k \leq m(T)} |M_{2,1}(k)| = o_{\mathbb{P}}((\log \log T)^{-1/2}).$$

Next we write

$$|M_{2,3}(\ell)| \leq \left| \sum_{t=\ell+1}^T e_t \right| + \left| \left( \sum_{t=\ell+1}^T \mathbf{x}_t \right)^\top (\hat{\beta}_T - \beta) \right|.$$

Replacing (A.7) with (A.15), one can establish along the lines of the proof of (A.18) that

$$T^{-1/2} \max_{T-m(T) \leq k \leq T-n(T)} \left| \sum_{t=\ell+1}^T e_t \right| = o_{\mathbb{P}}((\log \log T)^{-1/2}).$$

Lemma A.2 and (A.14) give

$$\begin{aligned} T^{-1/2} \max_{T-m(T) \leq \ell \leq T-n(T)} \left| \left( \sum_{t=\ell+1}^T \mathbf{x}_t \right)^{\top} (\hat{\boldsymbol{\beta}}_T - \boldsymbol{\beta}) \right| &= O_{\mathbb{P}}(1/T) \max_{T-m(T) \leq \ell \leq T-n(T)} (T - \ell) \\ &= O_{\mathbb{P}}(m(T)/T), \end{aligned}$$

which completes the proof of

$$\max_{T-m(T) \leq \ell \leq T-n(T)} M_{2,3}(\ell) = o_{\mathbb{P}}((\log \log T)^{-1/2}).$$

□

*Proof of Theorem 2.1.* It follows from Lemmas A.5–A.12 that

$$H_T = \max \left( \max_{n(T) \leq k \leq m(T)} \frac{1}{\sqrt{k}} \left| \sum_{t=1}^k e_t \right|, \max_{T-m(T) \leq \ell \leq T-n(T)} \frac{1}{\sqrt{T-\ell}} \left| \sum_{t=\ell}^T e_t \right| \right) + o_{\mathbb{P}}((\log \log T)^{-1/2})$$

and

$$M_T = \max_{n(T) \leq k \leq m(T)} \frac{1}{\sqrt{k}} \left| \sum_{t=1}^k e_t \right| + \max_{T-m(T) \leq \ell \leq T-n(T)} \frac{1}{\sqrt{T-\ell}} \left| \sum_{t=\ell}^T e_t \right| + o_{\mathbb{P}}((\log \log T)^{-1/2}).$$

We get from (A.3) that

$$D_T = T^{-1/2} \max_{1 \leq \ell \leq T} \left| \sum_{t=1}^{\ell} e_t - \frac{\ell}{T} \sum_{t=1}^T e_t - \left( \sum_{t=1}^{\ell} \mathbf{x}_t - \frac{\ell}{T} \sum_{t=1}^T \mathbf{x}_t \right)^{\top} (\hat{\boldsymbol{\beta}}_T - \boldsymbol{\beta}) \right|.$$

By Assumption 2.1 we can use the approximation in Aue et al. (2014) and get

$$T^{-1/2} \max_{1 \leq \ell \leq T} \left\| \sum_{t=1}^{\ell} \mathbf{x}_t - \frac{\ell}{T} \sum_{t=1}^T \mathbf{x}_t \right\| = O_{\mathbb{P}}(1)$$

and therefore by Lemma A.2 we have

$$D_T = T^{-1/2} \max_{1 \leq \ell \leq T} \left| \sum_{t=1}^{\ell} e_t - \frac{\ell}{T} \sum_{t=1}^T e_t \right| + o_{\mathbb{P}}(1).$$

Also, by Aue et al. (2014) we also conclude that

$$\begin{aligned} T^{-1/2} \max_{1 \leq \ell \leq T} \left| \sum_{t=1}^{\ell} e_t - \frac{\ell}{T} \sum_{t=1}^T e_t \right| & \tag{A.19} \\ &= T^{-1/2} \max_{m(T)+1 \leq \ell \leq T-m(T)-1} \left| \sum_{t=m(T)+1}^{\ell} e_t - \frac{\ell}{T} \sum_{t=m(T)+1}^{T-m(T)-1} e_t \right| + o_{\mathbb{P}}(1). \end{aligned}$$

Thus we need to show only

$$\begin{aligned} & \left( a_T \max_{n(T) \leq k \leq m(T)} \frac{1}{\sigma\sqrt{k}} \left| \sum_{t=1}^k e_t \right| - b_T, a_T \max_{T-m(T) \leq \ell \leq T-n(T)} \frac{1}{\sigma\sqrt{T-\ell}} \left| \sum_{t=\ell}^T e_t \right| - b_T, \right. \\ & \quad \left. \frac{1}{\sigma\sqrt{T}} \max_{m(T)+1 \leq \ell \leq T-m(T)-1} \left| \sum_{t=m(T)+1}^{\ell} e_t - \frac{\ell}{T} \sum_{t=m(T)+1}^{T-m(T)-1} e_t \right| \right) \\ & \xrightarrow{\mathcal{D}} \left( \xi_1, \xi_2, \sup_{0 \leq t \leq 1} |B(t)| \right), \end{aligned} \quad (\text{A.20})$$

where  $\xi_1$  and  $\xi_2$  are random variables with distribution defined in (2.4) and  $\{B(t), 0 \leq t \leq 1\}$  is a Brownian bridge. Also,  $\xi_1, \xi_2$  and  $\{B(t), 0 \leq t \leq 1\}$  are independent. In (A.7) and (A.15) we obtained weighted approximations for  $\sum_{t=1}^k e_t$  and  $\sum_{t=\ell}^T e_t$  with some Wiener processes. However, the proof of (A.20) requires a joint approximation of these partial sums on  $(k, \ell) \in A_T$  and  $\sum_{t=u}^v e_t, m(T) + 1 \leq u < v \leq T - m(T) - 1$ . The approximation in Aue et al. (2014) uses the blocking technique and therefore we can define three independent Wiener processes  $W_{T,1}(\cdot), W_{T,2}(\cdot)$  and  $W_{T,3}(\cdot)$  such that

$$\max_{n(T) \leq k \leq m(T)} k^{1/2-\delta} \left| \sum_{t=1}^k e_t - \sigma W_{T,1}(k) \right| = O_{\mathbb{P}}(1), \quad (\text{A.21})$$

$$\max_{T-m(T) \leq \ell \leq T-n(T)} (T-\ell)^{1/2-\delta} \left| \sum_{t=\ell}^T e_t - \sigma W_{T,2}(T-\ell) \right| = O_{\mathbb{P}}(1) \quad (\text{A.22})$$

with some  $\delta > 0$  and

$$\max_{m(T)+1 \leq k \leq T-m(T)-1} \left| \sum_{t=m(T)+1}^k e_t - \sigma W_{T,3}(k-m(T)) \right| = o_{\mathbb{P}}(1). \quad (\text{A.23})$$

The approximation in (A.23) yields that

$$\begin{aligned} & \frac{1}{\sqrt{T}} \max_{m(T)+1 \leq \ell \leq T-m(T)-1} \left| \sum_{t=m(T)+1}^{\ell} e_t - \frac{\ell}{T} \sum_{t=m(T)+1}^{T-m(T)-1} e_t \right. \\ & \quad \left. - \sigma \left( W_{T,3}(\ell - m(T)) - \frac{\ell}{T} W_{T,3}(T - 2m(T) - 1) \right) \right| = o_{\mathbb{P}}(1) \end{aligned}$$

and the modulus of continuity of the Wiener process (cf. Csörgő and Révész (1981)) gives

$$\begin{aligned} & \frac{1}{\sqrt{T}} \max_{m(T)+1 \leq \ell \leq T-m(T)-1} \left| W_{T,3}(\ell - m(T)) - \frac{\ell}{T} W_{T,3}(T - 2m(T) - 1) \right| \\ & = \frac{1}{\sqrt{T}} \sup_{0 \leq x \leq T} \left| W_{T,3}(x) - \frac{x}{T} W_{T,3} \right| + o_{\mathbb{P}}(1). \end{aligned}$$

Thus we have

$$\begin{aligned} & \frac{1}{\sqrt{T}} \max_{m(T)+1 \leq \ell \leq T-m(T)-1} \left| \sum_{t=m(T)+1}^{\ell} e_t - \frac{\ell}{T} \sum_{t=m(T)+1}^{T-m(T)-1} e_t \right| \\ &= \frac{\sigma}{\sqrt{T}} \sup_{0 \leq x \leq T} \left| W_{T,3}(x) - \frac{x}{T} W_{T,3}(x) \right| + o_{\mathbb{P}}(1). \end{aligned} \quad (\text{A.24})$$

It is easy to see that

$$\frac{1}{\sqrt{T}} \sup_{0 \leq x \leq T} \left| W_{T,3}(x) - \frac{x}{T} W_{T,3}(x) \right| \stackrel{\mathcal{D}}{=} \sup_{0 \leq t \leq 1} |B(t)|, \quad (\text{A.25})$$

where  $B(t)$  denotes a Brownian bridge. It follows from (A.21) that

$$\max_{n(T) \leq k \leq m(T)} \frac{1}{\sqrt{k}} \left| \sum_{t=1}^k e_t \right| - \max_{n(T) \leq k \leq m(T)} \frac{\sigma}{\sqrt{k}} \left| W_{T,1}(k) \right| = O_{\mathbb{P}}((\log T)^{-\kappa\delta}) \quad (\text{A.26})$$

and (A.22) yields

$$\begin{aligned} & \max_{T-m(T) \leq k \leq T-n(T)} \frac{1}{\sqrt{T-\ell}} \left| \sum_{t=\ell}^T e_t \right| - \max_{T-m(T) \leq \ell \leq T-n(T)} \frac{\sigma}{\sqrt{T-\ell}} \left| W_{T,2}(T-\ell) \right| \\ &= O_{\mathbb{P}}((\log T)^{-\kappa\delta}). \end{aligned} \quad (\text{A.27})$$

The asymptotic independence in (A.20) is an immediate consequence of (A.24), (A.26) and (A.27).

The Darling–Erdős limit result (cf. Appendix A in Csörgő and Horváth (1997)) states that

$$a_{m(T)/n(T)} \max_{n(T) \leq k \leq m(T)} \frac{1}{\sqrt{k}} |W(k)| - b_{m(T)/n(T)} \xrightarrow{\mathcal{D}} \xi, \quad (\text{A.28})$$

for any Wiener process  $W(\cdot)$ , where  $\xi$  is a random variable with distribution function  $\exp(-e^{-x})$ . Elementary calculations give

$$|a_T - a_{m(T)/n(T)}| = O((\log \log T)^{1/2} / \log T) = o((\log \log T)^{-1/2})$$

and similarly

$$|b_T - b_{m(T)/n(T)}| = o(1).$$

Hence (A.28) can be rewritten as

$$a_T \max_{n(T) \leq k \leq m(T)} \frac{1}{\sqrt{k}} |W(k)| - b_T \xrightarrow{\mathcal{D}} \xi. \quad (\text{A.29})$$

Now (A.20) follows from (A.24)–(A.27) and (A.29).  $\square$

*Proof of Theorem 2.2.* We follow the proof of Theorem 2.1. By Lemmas A.5–A.12 we have that

$$M_T^{(1)} = \max_{n(T) \leq k \leq m(T)} \frac{1}{\sqrt{k}} \left| \sum_{t=1}^k e_t \right| + o_{\mathbb{P}}((\log \log T)^{-1/2}), \quad (\text{A.30})$$

$$M_T^{(2)} = \max_{T-m(T) \leq k \leq T-n(T)} \frac{1}{\sqrt{T-\ell}} \left| \sum_{t=\ell}^T e_t \right| + o_{\mathbb{P}}((\log \log T)^{-1/2}), \quad (\text{A.31})$$

and

$$M_T^* = \max_{\mathfrak{M}_0^*} \sum_{i=2}^m \frac{1}{\sqrt{T}} \left| \sum_{t=k_{i-1}}^{k_i} e_t - \frac{k_i - k_{i-1}}{T} \sum_{t=m(T)+1}^{T-m(T)-1} e_t \right|, \quad (\text{A.32})$$

where  $\mathfrak{M}_0^* = \{n(T) \leq k_1 \leq k_2 \leq \dots \leq k_m \leq T - m(T) - 1\}$  (cf. (A.19)). Repeating the arguments used in the proof of Theorem 2.1, we obtain the asymptotic independence of  $a(T)M_T^{(1)}/\sigma - b(T)$ ,  $a(T)M_T^{(2)}/\sigma - b(T)$  and  $M_T^*$  from (A.30)–(A.32). We also conclude from Lemmas A.5–A.12 and (A.30), (A.31) that

$$\left( a_T \frac{M_T^{(1)}}{\sigma} - b_T, a_T \frac{M_T^{(2)}}{\sigma} - b_T \right) \xrightarrow{\mathcal{D}} (\xi_1, \xi_2),$$

where  $\xi_1$  and  $\xi_2$  defined in (2.4). We obtain from (A.24) and the continuity of the Wiener process that

$$\begin{aligned} & \max_{\mathfrak{M}_0^*} \sum_{i=2}^m \frac{1}{\sqrt{T}} \left| \sum_{t=k_{i-1}}^{k_i} e_t - \frac{k_i - k_{i-1}}{T} \sum_{t=m(T)+1}^{T-m(T)-1} e_t \right| \\ & \xrightarrow{\mathcal{D}} \sigma \max_{0 \leq u_1 \leq u_2 \leq \dots \leq u_m \leq 1} \sum_{i=2}^m |W(u_i) - W(u_{i-1}) - (u_i - u_{i-1})W(1)|, \end{aligned}$$

with  $u_0 = 0$ , where  $W(t)$ ,  $0 \leq t \leq 1$  is a Wiener process. Since  $B(t) = W(t) - tW(1)$  is a Brownian bridge, the proof of Theorem 2.2 is complete.  $\square$

## B. APPENDIX: PROOF OF THEOREM 2.4.

The OLS estimator  $\beta_T$  under  $H_A$  can be decomposed as

$$\begin{aligned} \hat{\beta}_T &= (\mathbf{X}_T \mathbf{X}_T^\top)^{-1} \mathbf{X}_T \mathbf{Y}_T \\ &= (\mathbf{X}_T \mathbf{X}_T^\top)^{-1} \left( \sum_{i=1}^{m+1} \sum_{t=k_{i-1}^*+1}^{k_i^*} \mathbf{x}_t \mathbf{x}_t^\top \beta^{(i)} + \sum_{t=1}^T \mathbf{x}_t e_t \right). \end{aligned}$$

We continue with the analogue of Lemma A.2 under  $H_A$ .

**Lemma B.1.** *If  $H_A$ , Assumptions 2.1 and 2.2 hold, then we have that*

$$\hat{\beta}_T = \tilde{\beta}_T + O_{\mathbb{P}}(T^{-1/2}). \quad (\text{B.1})$$

*Proof.* Since  $x_{t,i}x_{j,t}$ ,  $-\infty < t < \infty$  is a Bernoulli shift, repeating the arguments leading to A.2 we obtain that

$$\mathbb{E} \left( \sum_{t=1}^N (x_{t,i}x_{j,t} - \mathbb{E}x_{t,i}x_{j,t}) \right)^2 = O(N), \quad \text{as } N \rightarrow \infty.$$

and therefore

$$(\mathbf{X}_T \mathbf{X}_T^\top)^{-1} = \frac{1}{T} \mathbf{A}^{-1} + O_{\mathbb{P}}(T^{-3/2})$$

and

$$\sum_{t=k_{i-1}^*+1}^{k_i^*} \mathbf{x}_t \mathbf{x}_t^\top = (k_i^* - k_{i-1}^*) \mathbf{A} + O_{\mathbb{P}} \left( \sqrt{k_i^* - k_{i-1}^*} \right),$$

and therefore by (A.2) we have (B.1).  $\square$

*Proof of Theorem 2.4.* Similarly to A.2 we have

$$\mathbb{E} \left\| \sum_{t=T_1}^{T_2} (\mathbf{x}_t - \mathbf{c}) \right\|^2 = O(T_2 - T_1), \quad \text{as } N \rightarrow \infty, \quad (\text{B.2})$$

and therefore

$$\sum_{t=1}^T \mathbf{x}_t = T \mathbf{c} + O_{\mathbb{P}}(\sqrt{T}) \quad \text{and} \quad \sum_{t=k_{i-1}^*+1}^{k_i^*} \mathbf{x}_t + O_{\mathbb{P}} \left( \sqrt{k_i^* - k_{i-1}^*} \right), \quad 1 \leq i \leq \mathbf{m} + 1.$$

Since

$$\hat{e}_t = y_t - \mathbf{x}_t^\top \hat{\boldsymbol{\beta}}_T = e_t - \mathbf{x}_t^\top (\hat{\boldsymbol{\beta}}_T - \boldsymbol{\beta}^{(i)}), \quad \text{if } k_{i-1}^* + 1 \leq t \leq k_i^*, 1 \leq i \leq \mathbf{m} + 1, \quad (\text{B.3})$$

we get

$$\begin{aligned} M_1(k_1^*) &= \frac{1}{\sqrt{k_1^*}} \left( \sum_{t=1}^{k_1^*} \mathbf{x}_t^\top (\hat{\boldsymbol{\beta}}_T - \boldsymbol{\beta}^{(1)}) - \frac{k_1^*}{T} \sum_{t=1}^T \mathbf{x}_t^\top (\hat{\boldsymbol{\beta}}_T - \tilde{\boldsymbol{\beta}}_T) \right) + O_{\mathbb{P}}(1) \\ &= \sqrt{k_1^*} \mathbf{c}^\top (\tilde{\boldsymbol{\beta}}_T - \boldsymbol{\beta}^{(1)}) (1 + o_{\mathbb{P}}(1)) + O_{\mathbb{P}}(1) \end{aligned} \quad (\text{B.4})$$

and

$$M_{\mathbf{m}+1}(k_{\mathbf{m}}^*) = \sqrt{T - k_{\mathbf{m}}^*} \mathbf{m} \mathbf{c}^\top (\tilde{\boldsymbol{\beta}}_T - \boldsymbol{\beta}^{(\mathbf{m}+1)}) (1 + o_{\mathbb{P}}(1)) + O_{\mathbb{P}}(1). \quad (\text{B.5})$$

Using again Lemma B.1 and (B.2) we conclude for all  $2 \leq i \leq \mathbf{m}$  that

$$M_i(k_{i-1}^*, k_i^*) = \frac{k_i^* - k_{i-1}^*}{\sqrt{T}} \mathbf{c}^\top (\tilde{\boldsymbol{\beta}}_T - \boldsymbol{\beta}^{(i)}) (1 + o_{\mathbb{P}}(1)) + O_{\mathbb{P}}(1). \quad (\text{B.6})$$

The result now follows from (B.4)–(B.6).  $\square$

*Proofs of Corollaries 2.1 and 2.2.* The results can be derived from Theorem 2.4 by elementary calculations.  $\square$

### C. APPENDIX: PROOFS OF THEOREMS 3.1 AND 3.2

*Proof of Theorem 3.1.* Using (A.3) we get

$$S_T^2 = \sum_{t=1}^T e_t^2 - 2 \left( \sum_{t=1}^T e_t \mathbf{x}_t \right)^\top (\hat{\boldsymbol{\beta}}_T - \boldsymbol{\beta}) + \sum_{t=1}^T \left( \mathbf{x}_t^\top (\hat{\boldsymbol{\beta}}_T - \boldsymbol{\beta}) \right)^2. \quad (\text{C.1})$$

Assumption 2.1 implies along the lines of the proof of (A.5) that

$$\left| \sum_{t=1}^T e_t^2 - T \sigma^2 \right| = O_{\mathbb{P}}(T^{1/2}). \quad (\text{C.2})$$



Putting together Lemma A.1 and (A.2) we conclude

$$\left| \left( \sum_{t=1}^T e_t \mathbf{x}_t \right)^\top (\hat{\boldsymbol{\beta}}_T - \boldsymbol{\beta}) \right| = O_{\mathbb{P}}(1),$$

and by the ergodic theorem we get

$$\sum_{t=1}^T \left( \mathbf{x}_t^\top (\hat{\boldsymbol{\beta}}_T - \boldsymbol{\beta}) \right)^2 = O_{\mathbb{P}}(1).$$

This completes the proof of (3.2).

Lemma B.1 and (A.2) give

$$\left| \left( \sum_{t=1}^T e_t \mathbf{x}_t \right)^\top (\hat{\boldsymbol{\beta}}_T - \boldsymbol{\beta}) \right| = O_{\mathbb{P}}(T), \quad \sum_{t=1}^T \left( \mathbf{x}_t^\top (\hat{\boldsymbol{\beta}}_T - \boldsymbol{\beta}) \right)^2 = O_{\mathbb{P}}(1),$$

and therefore (3.3) follows from (C.2).

*Proof of Theorem 3.2.* It follows from (A.3)

$$\begin{aligned} \hat{\gamma}_\ell &= \frac{1}{T} \sum_{t=1}^{T-\ell} e_t e_{t+\ell} - \frac{1}{T} \sum_{t=1}^{T-\ell} e_t \mathbf{x}_{t+\ell}^\top (\hat{\boldsymbol{\beta}}_T - \boldsymbol{\beta}) - \frac{1}{T} \sum_{t=1}^{T-\ell} e_{t+\ell} \mathbf{x}_t^\top (\hat{\boldsymbol{\beta}}_T - \boldsymbol{\beta}) \\ &\quad + \frac{1}{T} \sum_{t=1}^{T-\ell} \mathbf{x}_t^\top (\hat{\boldsymbol{\beta}}_T - \boldsymbol{\beta}) \mathbf{x}_{t+\ell}^\top (\hat{\boldsymbol{\beta}}_T - \boldsymbol{\beta}). \end{aligned} \quad (\text{C.3})$$

It is easy to see that

$$\begin{aligned} &\mathbb{E} \left( \sum_{\ell=1}^{T-1} K(\ell/h) \frac{1}{T} \sum_{t=1}^{T-\ell} (e_t e_{t+\ell} - \mathbb{E} x_0 x_\ell) \right)^2 \\ &= \frac{1}{T^2} \sum_{\ell, \ell'=1}^{T-1} \sum_{t=1}^{T-\ell} \sum_{s=1}^{T-\ell'} K(\ell/h) K(\ell'/h) (\mathbb{E}[e_t e_{t+\ell} e_s e_{s+\ell'}] - \mathbb{E} e_0 e_\ell \mathbb{E} e_0 e_{\ell'}) \\ &= Q_T. \end{aligned}$$

Following the proof of Horváth and Rice (2015) we write by Assumption 3.2

$$\begin{aligned} Q_T &= O(1/T^2) \sum_{1 \leq t \leq s \leq T-1} \sum_{1 \leq \ell, \ell' \leq ch} |\mathbb{E}[e_t e_{t+\ell} e_s e_{s+\ell'}] - \mathbb{E} e_0 e_\ell \mathbb{E} e_0 e_{\ell'}| \\ &= O(1/T^2) \sum_{1 \leq t \leq s \leq T-1} \sum_{1 \leq \ell, \ell' \leq ch} |\mathbb{E}[e_t e_{t+\ell} e_s e_{s+\ell'}] - \mathbb{E} e_0 e_\ell \mathbb{E} e_0 e_{\ell'}| \\ &= O(1/T) \sum_{1 \leq v \leq T-1} \sum_{1 \leq \ell, \ell' \leq ch} |\mathbb{E}[e_0 e_\ell e_v e_{v+\ell'}] - \mathbb{E} e_0 e_\ell \mathbb{E} e_0 e_{\ell'}| \\ &= O(1/T)(Q_{T,1} + Q_{T,2}), \end{aligned}$$

where

$$Q_{T,1} = \sum_{(v, \ell, \ell') \in G_{T,1}} |\mathbb{E}[e_0 e_\ell e_v e_{v+\ell'}] - \mathbb{E} e_0 e_\ell \mathbb{E} e_0 e_{\ell'}|$$

and

$$Q_{T,2} = \sum_{(v,\ell,\ell') \in G_{T,2}} |\mathbb{E}[e_0 e_\ell e_v e_{v+\ell'}] - \mathbb{E}e_0 e_\ell \mathbb{E}e_0 e_{\ell'}|$$

with

$$G_{T,1} = \{(v, \ell, \ell') : h+1 \leq v \leq T, 1 \leq \ell, \ell' \leq h\},$$

$$G_{T,2} = \{(v, \ell, \ell') : 1 \leq v \leq h, 1 \leq \ell, \ell' \leq h\}.$$

Next we define

$$\bar{e}_{v,v-\ell-1} = f(\varepsilon_v, \varepsilon_{v-1}, \dots, \varepsilon_{\ell+1}, \varepsilon'_\ell, \varepsilon'_{\ell-1}, \dots)$$

and

$$\bar{e}_{v+\ell', v+\ell'-\ell-1} = f(\varepsilon_{v+\ell'}, \varepsilon_{v+\ell'-1}, \dots, \varepsilon_{\ell+1}, \varepsilon'_\ell, \varepsilon'_{\ell-1}, \dots),$$

where  $\varepsilon'_\ell, -\infty < \ell < \infty$  are independent copies of  $\varepsilon_0$ , independent of  $\varepsilon_j, -\infty < j < \infty$ . It follows from the Bernoulli assumption that  $(e_0, e_\ell)$  is independent of  $(\bar{e}_{v,v-\ell-1}, \bar{e}_{v+\ell', v+\ell'-\ell-1})$ . Also, according to the construction,  $(\bar{e}_{v,v-\ell-1}, \bar{e}_{v+\ell', v+\ell'-\ell-1})$  and  $(e_v, e_{v+\ell'})$  are identically distributed. Hence

$$\mathbb{E}[e_0 e_\ell e_v e_{v+\ell'}] - \mathbb{E}e_0 e_\ell \mathbb{E}e_0 e_{\ell'} = \mathbb{E}e_0 e_\ell [e_v e_{v+\ell'} - \bar{e}_{v,v-\ell-1} \bar{e}_{v+\ell', v+\ell'-\ell-1}].$$

It follows from Assumption 2.1 that

$$(\mathbb{E}(e_v - \bar{e}_{v,v-\ell-1})^4)^{1/4} \leq c(v-\ell)^{-\alpha} \quad \text{and} \quad (\mathbb{E}(e_{v+\ell'} - \bar{e}_{v+\ell', v+\ell'-\ell-1})^4)^{1/4} \leq c(v+\ell'-\ell)^{-\alpha}$$

with some constant  $c$  for all  $(v, \ell, \ell') \in G_{T,1}$ . Hence

$$\begin{aligned} & \mathbb{E}|e_0 e_\ell [e_v e_{v+\ell'} - \bar{e}_{v,v-\ell-1} \bar{e}_{v+\ell', v+\ell'-\ell-1}]| \\ & \leq \mathbb{E}|e_0 e_\ell e_v [e_{v+\ell'} - \bar{e}_{v+\ell', v+\ell'-\ell-1}]| + \mathbb{E}|e_0 e_\ell \bar{e}_{v+\ell', v+\ell'-\ell-1} [e_v - \bar{e}_{v,v-\ell-1}]| \\ & \leq (\mathbb{E}e_0^4 \mathbb{E}e_\ell^4 \mathbb{E}e_v^4 \mathbb{E}|e_{v+\ell'} - \bar{e}_{v+\ell', v+\ell'-\ell-1}|^4)^{1/4} + (\mathbb{E}e_0^4 \mathbb{E}e_\ell^4 \mathbb{E}\bar{e}_{v+\ell', v+\ell'-\ell-1}^4 \mathbb{E}|e_v - \bar{e}_{v,v-\ell-1}|^4)^{1/4} \\ & \leq c_*(v-\ell)^{-\alpha} \end{aligned}$$

with some constant  $c_*$ . Hence elementary arguments give that

$$Q_{T,1} = O(1) \sum_{(v,\ell,\ell') \in G_{T,1}} (v-\ell)^{-\alpha} = O(h) \int_1^h \int_{h+1}^T (x-y)^{-\alpha} dx dy = O(h).$$

Next we note

$$Q_{T,2} \leq \sum_{(v,\ell,\ell') \in G_{T,2}} |\mathbb{E}[e_0 e_\ell e_v e_{v+\ell'}]| + \sum_{(v,\ell,\ell') \in G_{T,2}} |\mathbb{E}e_0 e_\ell \mathbb{E}e_0 e_{\ell'}|$$

and

$$\sum_{(v,\ell,\ell') \in G_{T,2}} |\mathbb{E}e_0 e_\ell \mathbb{E}e_0 e_{\ell'}| \leq h \left( \sum_{\ell=1}^{\infty} |\mathbb{E}e_0 e_\ell| \right)^2 < \infty.$$

Using the variables  $e_{t,m}$  defined in Assumption 2.1 we write for all  $0 \leq s \leq t \leq v \leq 2h$

$$\begin{aligned} e_0 e_s e_t e_v &= e_0 e_{s,s} (e_t - e_{t,t-s}) (e_v - e_{v,v-t}) + e_0 e_{s,s} (e_t - e_{t,t-s}) e_{v,v-t} + e_0 e_{s,s} e_{t,t-s} (e_v - e_{v,v}) \\ & \quad + e_0 e_{s,s} e_{t,t-s} e_{v,v} + e_0 (e_s - e_{s,s}) (e_t - e_{t,t-s}) e_v + e_0 (e_s - e_{s,s}) e_{t,t-s} e_{v,v-t} \\ & \quad + e_0 (e_s - e_{s,s}) e_{t,t-s} (e_v - e_{v,v-t}). \end{aligned}$$

It follows from the definitions of  $e_{t,m}$  that  $e_{v,v-t}$  is independent of  $e_0 e_{s,s} (e_t - e_{t,t-s})$ ,  $e_0$  is independent of  $e_{s,s} e_{t,t-s} e_{v,v}$ ,  $e_{v,v-t}$  is independent of  $e_0 (e_s - e_{s,s}) e_{t,t-s}$  and therefore

$$\mathbb{E}[e_0 e_{s,s} (e_t - e_{t,t-s}) e_{v,v-t}] = 0, \quad \mathbb{E}[e_0 e_{s,s} e_{t,t-s} e_{v,v}] = 0 \quad \text{and} \quad \mathbb{E}[e_0 (e_s - e_{s,s}) e_{t,t-s} e_{v,v-t}] = 0.$$

Using Assumption 2.1 we obtain that

$$\begin{aligned} & \sum_{0 \leq s \leq t \leq v \leq 2h} |\mathbb{E} e_0 e_{s,s} (e_t - e_{t,t-s})(e_v - e_{v,v-t})| \\ & \leq (\mathbb{E} e_0^4)^{1/2} \sum_{0 \leq s \leq t \leq v \leq 2h} (\mathbb{E} (e_t - e_{t,t-s})^4)^{1/4} (\mathbb{E} (e_v - e_{v,v-t})^4)^{1/4} \\ & = O(h). \end{aligned}$$

Similarly

$$\begin{aligned} & \sum_{0 \leq s \leq t \leq v \leq 2h} |\mathbb{E} e_0 (e_s - e_{s,s})(e_t - e_{t,t-s}) e_v| = O(h), \\ & \sum_{0 \leq s \leq t \leq v \leq 2h} |\mathbb{E} e_0 (e_s - e_{s,s}) e_{t,t-s} (e_v - e_{v,v-t})| = O(h) \end{aligned}$$

and

$$\sum_{0 \leq s \leq t \leq v \leq 2h} |\mathbb{E} e_0 e_{s,s} e_{t,t-s} (e_v - e_{v,v})| = O(1) \int_0^{2h} \int_s^{2h} \int_t^{2h} v^{-\alpha} dv dt ds = O(h).$$

Thus we conclude that

$$\mathbb{E} \left( \sum_{\ell=1}^{T-1} \frac{1}{T} \sum_{t=1}^{T-\ell} (e_t e_{t+\ell} - \mathbb{E} x_0 x_\ell) \right)^2 = O\left(\frac{h}{T}\right)$$

and therefore

$$\sum_{\ell=1}^{T-1} K(\ell/T) \frac{1}{T} \sum_{t=1}^{T-\ell} e_t e_{t+\ell} = \sum_{\ell=1}^{T-1} K(\ell/h) \frac{T-\ell}{T} \mathbb{E} e_0 e_\ell + O_{\mathbb{P}}((h/T)^{1/2}). \quad (\text{C.4})$$

Let

$$w_{\ell,i} = \frac{1}{T} \sum_{t=1}^{T-\ell} e_t x_{t+\ell,i}$$

and define

$$\hat{a}_{T,i} = \sum_{\ell=1}^{T-1} K(\ell/h) w_{\ell,i}.$$

Similarly to (C.4) we have for all  $1 \leq i \leq d$

$$\begin{aligned} \hat{a}_{T,i} &= \sum_{\ell=1}^{T-1} K(\ell/h) \frac{T-\ell}{T} \mathbb{E} e_0 x_{\ell,i} + O_{\mathbb{P}}((h/T)^{1/2}) \\ &= O(1) \sum_{\ell=1}^{\infty} |\mathbb{E} e_0 x_{\ell,i}| + O_{\mathbb{P}}((h/T)^{1/2}) \\ &= O_{\mathbb{P}}(1), \end{aligned} \quad (\text{C.5})$$

since by Assumption 2.1 we have  $|\mathbb{E} e_0 x_{\ell,i}| = \mathbb{E} (e_0 - e_{0,\ell}) x_{\ell,i}$  and therefore by the Cauchy-Schwartz inequality

$$\sum_{\ell=1}^{\infty} |\mathbb{E} e_0 x_{\ell,i}| = \sum_{\ell=1}^{\infty} |\mathbb{E} (e_0 - e_{0,\ell}) x_{\ell,i}| \leq \mathbb{E} (x_{0,i}^2)^{1/2} \sum_{\ell=1}^{\infty} (\mathbb{E} (e_0 - e_{0,\ell}))^{1/2} < \infty.$$

Along the lines of the arguments leading to (C.4) and (C.5) we get

$$\hat{b}_{T,i} = \sum_{\ell=1}^{T-1} K(\ell/h) \frac{1}{T} \sum_{t=1}^{T-\ell} x_{t,i} e_{t+\ell} = O_{\mathbb{P}}(1) \text{ for all } 1 \leq i \leq d, \quad (\text{C.6})$$

and

$$\hat{c}_{T,i,j} = \sum_{\ell=1}^{T-1} K(\ell/h) \frac{1}{T} \sum_{t=1}^{T-\ell} x_{t,i} x_{t+\ell,j} = O_{\mathbb{P}}(1) \text{ for all } 1 \leq i, j \leq d. \quad (\text{C.7})$$

Thus we conclude by Lemma

$$\hat{\sigma}_T^2 = \sum_{\ell=1}^{T-1} K(\ell/h) \frac{T-\ell}{T} \mathbb{E} e_0 e_\ell + O_{\mathbb{P}}((h/T)^{1/2}) + O_{\mathbb{P}}(T^{-1/2}).$$

Assumption 3.2 yields

$$\sum_{\ell=1}^{T-1} K(\ell/h) \frac{T-\ell}{T} \mathbb{E} e_0 e_\ell = \sum_{\ell=1}^{\infty} \mathbb{E} e_0 e_\ell + \sum_{\ell=1}^{hc} \left( K(\ell/h) \frac{T-\ell}{T} - 1 \right) \mathbb{E} e_0 e_\ell - \sum_{\ell=hc+1}^{\infty} \mathbb{E} e_0 e_\ell$$

and

$$\begin{aligned} \left| \sum_{\ell=1}^{hc} \left( K(\ell/h) \frac{T-\ell}{T} - 1 \right) \mathbb{E} e_0 e_\ell \right| &\leq \sum_{\ell=1}^{hc} |K(\ell/h) - 1| |\mathbb{E} e_0 e_\ell| + \sum_{\ell=1}^{hc} K(\ell/h) \left| \frac{T-\ell}{T} - 1 \right| |\mathbb{E} e_0 e_\ell| \\ &= O(1) \left( \frac{1}{h} + \frac{1}{T} \right) \sum_{\ell=1}^{\infty} \ell |\mathbb{E} e_0 e_\ell|. \end{aligned}$$

By Assumption 2.1 we get

$$\sum_{\ell=1}^{\infty} \ell |\mathbb{E} e_0 e_\ell| = \sum_{\ell=1}^{\infty} \ell |\mathbb{E} e_0 e_\ell| = \sum_{\ell=1}^{\infty} \ell |\mathbb{E} e_0 (e_\ell - e_{\ell,\ell})| \leq (\mathbb{E} e_0^2)^{1/2} \sum_{\ell=1}^{\infty} \ell (\mathbb{E} (e_\ell - e_{\ell,\ell})^2)^{1/2} < \infty$$

and

$$\sum_{\ell=hc+1}^{\infty} |\mathbb{E} e_0 e_\ell| \leq \frac{1}{h} \sum_{\ell=1}^{\infty} \ell |\mathbb{E} e_0 e_\ell|.$$

Since  $\hat{\gamma}_0 - \mathbb{E} e_0^2 = O_P(T^{-1/2})$ , (3.5) is proven.

We use (B.3) to prove (3.6). It follows from (C.5)–(C.7) and Lemma B.1 that

$$|\hat{a}_{T,i}| \|\hat{\beta}_T - \beta\| = O_{\mathbb{P}}(1), \quad |\hat{b}_{T,i}| \|\hat{\beta}_T - \beta\| = O_{\mathbb{P}}(1) \quad \text{and} \quad |\hat{c}_{T,i,j}| \|\hat{\beta}_T - \beta\| = O_{\mathbb{P}}(1). \quad (\text{C.8})$$

The result in (3.6) is an immediate consequence of (C.3), (C.4) and (C.8).

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TABLE C.1. Standardized CUSUM for Sub-samples

	Standardized CUSUM	P-Value
Jul. 2001–Nov. 2010	3.46	3.08%
Nov. 2004–Nov. 2014	3.23	3.89%
Jul. 2001–Nov. 2004	0.87	34.32%
Dec. 2004–Nov. 2010	0.69	39.48%
Dec. 2010–Nov. 2014	0.37	49.86%

TABLE C.2. The segmentation for the AHGCX time series

	Jul. 2001–Nov. 2004	Dec. 2004–Nov. 2010	Dec. 2010–Nov. 2014
$\alpha$	-1.09	0.65	-0.39
$\beta^M$	0.99	1.22	1.05
$\beta^{HML}$	0.05	-0.30	-0.41
$\beta^{SMB}$	0.48	0.35	0.40
$\beta^{MOM}$	0.25	0.18	0.09



TABLE C.3. Standardized CUSUM for Sub-samples

	Standardized CUSUM	P-Value
Feb. 1984–Aug. 2003	5.84	0.29%
Nov. 1996–Nov. 2014	6.02	0.24%
Feb. 1984–Nov. 1996	1.99	12.80%
Dec. 1996–Aug. 2003	0.66	40.30%
Sep. 2003–Nov. 2014	-0.62	84.54%

TABLE C.4. The segmentation for the IIVGX time series

	Feb. 1984 – Nov. 1996	Dec. 1996 – Aug. 2003	Sep. 2003 – Nov. 2014
$\alpha$	0.11	-1.21	-0.06
$\beta^M$	0.83	0.80	1.01
$\beta^{HML}$	-0.06	0.10	-0.03
$\beta^{SMB}$	-0.14	-0.07	-0.10
$\beta^{MOM}$	0.03	-0.11	0.01

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