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Leonid Galtchouk, Serguey Pergamenshchikov. Adaptive asymptotically efficient estimation in heteroscedastic nonparametric regression via model selection.. 2008. hal-00326910

HAL Id: hal-00326910

<https://hal.science/hal-00326910>

Preprint submitted on 7 Oct 2008

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Adaptive asymptotically efficient estimation in heteroscedastic nonparametric regression via model selection.

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October 7, 2008

Abstract

The paper deals with asymptotic properties of the adaptive procedure proposed in the author paper, 2007, for estimating a unknown nonparametric regression. We prove that this procedure is asymptotically efficient for a quadratic risk, i.e. the asymptotic quadratic risk for this procedure coincides with the Pinsker constant which gives a sharp lower bound for the quadratic risk over all possible estimates. ¹
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¹*AMS 2000 Subject Classification* : primary 62G08; secondary 62G05, 62G20

²*Key words*: asymptotic bounds, adaptive estimation, efficient estimation, heteroscedastic regression, nonparametric regression, Pinsker's constant.

1 Introduction

The paper deals with the estimation problem in the heteroscedastic non-parametric regression model

$$y_j = S(x_j) + \sigma_j(S) \xi_j, \quad (1.1)$$

where the design points $x_j = j/n$, $S(\cdot)$ is an unknown function to be estimated, $(\xi_j)_{1 \leq j \leq n}$ is a sequence of centered independent random variables with unit variance and $(\sigma_j(S))_{1 \leq j \leq n}$ are unknown scale functionals depending on the design points and the regression function S .

Typically, the notion of asymptotic optimality is associated with the optimal convergence rate of the minimax risk (see e.g., Ibragimov, Hasminskii, 1981; Stone, 1982). An important question in optimality results is to study the exact asymptotic behavior of the minimax risk. Such results have been obtained only in a limited number of investigations. As to the nonparametric estimation problem for heteroscedastic regression models we should mention the papers by Efromovich, 2007, Efromovich, Pinsker, 1996, and Galtchouk, Pergamenshchikov, 2005, concerning the exact asymptotic behavior of the \mathcal{L}_2 -risk and the paper by Brua, 2007, devoted to the efficient pointwise estimation for heteroscedastic regressions.

Heteroscedastic regression models are largely used in financial mathematics, in particular, in problem of calibrating (see e.g., Belomestny, Reiss, 2006). An example of heteroscedastic regression models is given by econometrics (see, for example, Goldfeld, Quandt, 1972, p. 83), where for consumer budget problems one uses some parametric version of model (1.1) with the scale coefficients defined as

$$\sigma_j^2(S) = c_0 + c_1 x_j + c_2 S^2(x_j), \quad (1.2)$$

where c_0 , c_1 and c_2 are some unknown positive constants.

The purpose of the article is to study asymptotic properties of the adaptive estimation procedure proposed in Galtchouk, Pergamenshchikov, 2007, for which a non-asymptotic oracle inequality was proved for quadratic risks. We will prove that this oracle inequality is asymptotically sharp, i.e. the asymptotic quadratic risk is minimal. It means the adaptive estimation procedure is efficient under some the conditions on the scales $(\sigma_j(S))_{1 \leq j \leq n}$ which are satisfied in the case (1.2). Note that in Efromovich, 2007, Efromovich, Pinsker, 1996, an efficient adaptive procedure is constructed for heteroscedastic regression when the scale coefficient is independent of S , i.e. $\sigma_j(S) = \sigma_j$.

In Galtchouk, Pergamenshchikov, 2005, for the model (1.1) the asymptotic efficiency was proved under strong the conditions on the scales which are not satisfied in the case (1.2). Moreover in the cited papers the efficiency was proved for the gaussian random variables $(\xi_j)_{1 \leq j \leq n}$ that is very restrictive for applications of proposed methods to practical problems.

In the paper we modify the risk. We take a additional supremum over the family of unknown noise distributions like to Galtchouk, Pergamenshchikov, 2006. This modification allows us to eliminate from the risk dependence on the noise distribution. Moreover for this risk a efficient procedure is robust with respect to changing the noise distribution.

It is well known to prove the asymptotic efficiency one has to show that the asymptotic quadratic risk coincides with the lower bound which is equal to the Pinsker constant. In the paper two problems are resolved: in the first one a upper bound for the risk is obtained by making use of the non-asymptotic oracle inequality from Galtchouk, Pergamenshchikov, 2007, in the second one we prove that this upper bound coincides with the Pinsker constant. Let us remember that the adaptive procedure proposed in Galtchouk, Pergamenshchikov, 2007, is based on weighted least-squares estimates, where the weights are proper modifications of the Pinsker weights for the homogeneous case (when $\sigma_1(S) = \dots = \sigma_n(S) = 1$) relative to a certain smoothness of the function S and this procedure chooses a best estimator for the quadratic risk among these estimators. To obtain the Pinsker constant for the model (1.1) one has to prove a sharp asymptotic lower bound for the quadratic risk in the case when the noise variance depends on the unknown regression function. In this case, as usually, we minorize the minimax risk by a bayesian one for a respective parametric family. Then for the bayesian risk we make use of a lower bound (see Theorem 6.1) which is a modification of the van Trees inequality (see, Gill, Levit, 1995).

The paper is organized as follows. In Section 2 we construct an adaptive estimation procedure. In Section 3 we formulate principal the conditions. The main results are presented in Section 4. The upper bound for the quadratic risk is given in Section 5. In Section 6 we give all main steps of proving the lower bound. In Subsection 6.1 we find the lower bound for the bayesian risk which minorizes the minimax risk. In Subsection 6.2 we study a special parametric functions family used to define the bayesian risk. In Subsection 6.3 we choose a prior distribution for bayesian risk to maximize the lower bound. Section 7 is devoted to explain how to use the given procedure in the case when the unknown regression function is non periodic.

In Section 8 we discuss the main results and their practical importance. The proofs are given in Section 9. The Appendix contains some technical results.

2 Adaptive procedure

In this section we describe the adaptive procedure proposed in Galtchouk, Pergamenschikov, 2006. We make use of the standard trigonometric basis $(\phi_j)_{j \geq 1}$ in $\mathcal{L}_2[0, 1]$, i.e.

$$\phi_1(x) = 1, \quad \phi_j(x) = \sqrt{2} Tr_j(2\pi[j/2]x), \quad j \geq 2, \quad (2.1)$$

where the function $Tr_j(x) = \cos(x)$ for even j and $Tr_j(x) = \sin(x)$ for odd j ; $[x]$ denotes the integer part of x .

To evaluate the error of estimation in the model (1.1) we will make use of the empiric norm in the Hilbert space $\mathcal{L}_2[0, 1]$, generated by the design points $(x_j)_{1 \leq j \leq n}$ of model (1.1). To this end, for any functions u and v from $\mathcal{L}_2[0, 1]$, we define the empiric inner product

$$(u, v)_n = \frac{1}{n} \sum_{l=1}^n u(x_l) v(x_l).$$

Moreover, we will use this inner product for vectors in \mathbb{R}^n as well, i.e. if $u = (u_1, \dots, u_n)'$ and $v = (v_1, \dots, v_n)'$, then

$$(u, v)_n = \frac{1}{n} u'v = \frac{1}{n} \sum_{l=1}^n u_l v_l.$$

The prime denotes the transposition.

Notice that if n is odd, then the functions $(\phi_j)_{1 \leq j \leq n}$ are orthonormal with respect to this inner product, i.e. for any $1 \leq i, j \leq n$,

$$(\phi_i, \phi_j)_n = \frac{1}{n} \sum_{l=1}^n \phi_i(x_l) \phi_j(x_l) = \mathbf{K}r_{ij}, \quad (2.2)$$

where $\mathbf{K}r_{ij}$ is Kronecker's symbol, $\mathbf{K}r_{ij} = 1$ if $i = j$ and $\mathbf{K}r_{ij} = 0$ for $i \neq j$.

Remark 2.1. *Note that in the case of even n , the basis (2.1) is orthogonal and it is orthonormal except the n th function for which the normalizing constant should be changed. The corresponding modifications of the formulas for even n one can see in Galtchouk, Pergamenschikov, 2005. To avoid these complications of formulas related to even n , we suppose n to be odd.*

Thanks to this basis we pass to the discrete Fourier transformation of model (1.1):

$$\widehat{\theta}_{j,n} = \theta_{j,n} + \frac{1}{\sqrt{n}} \xi_{j,n}, \quad (2.3)$$

where $\widehat{\theta}_{j,n} = (Y, \phi_j)_n$, $Y = (y_1, \dots, y_n)'$, $\theta_{j,n} = (S, \phi_j)_n$ and

$$\xi_{j,n} = \frac{1}{\sqrt{n}} \sum_{l=1}^n \sigma_l(S) \xi_l \phi_j(x_l).$$

We estimate the function S by the weighted least squares estimator

$$\widehat{S}_\lambda = \sum_{j=1}^n \lambda(j) \widehat{\theta}_{j,n} \phi_j, \quad (2.4)$$

where the weight vector $\lambda = (\lambda(1), \dots, \lambda(n))'$ belongs to some finite set Λ from $[0, 1]^n$ with $n \geq 3$.

Here we make use of the weight family Λ introduced in Galtchouk, Pergamshchikov, 2008, i.e.

$$\Lambda = \{\lambda_\alpha, \alpha \in \mathcal{A}\}, \quad \mathcal{A} = \{1, \dots, k^*\} \times \{t_1, \dots, t_m\}, \quad (2.5)$$

where $t_i = i\varepsilon$ and $m = [1/\varepsilon^2]$. We suppose that the parameters $k^* \geq 1$ and $0 < \varepsilon \leq 1$ are functions of n , i.e. $k^* = k_n^*$ and $\varepsilon = \varepsilon_n$, such that,

$$\begin{cases} \lim_{n \rightarrow \infty} k_n^* = +\infty, & \lim_{n \rightarrow \infty} \frac{k_n^*}{\ln n} = 0, \\ \lim_{n \rightarrow \infty} \varepsilon_n = 0 & \text{and} \quad \lim_{n \rightarrow \infty} n^\nu \varepsilon_n = +\infty, \end{cases} \quad (2.6)$$

for any $\nu > 0$. For example, one can take for $n \geq 3$

$$\varepsilon_n = 1/\ln n \quad \text{and} \quad k_n^* = \bar{k} + \sqrt{\ln n},$$

where \bar{k} is any nonnegative constant.

For each $\alpha = (\beta, t) \in \mathcal{A}$ we define the weight vector $\lambda_\alpha = (\lambda_\alpha(1), \dots, \lambda_\alpha(n))'$ as

$$\lambda_\alpha(j) = \mathbf{1}_{\{1 \leq j \leq j_0\}} + (1 - (j/\omega(\alpha))^\beta) \mathbf{1}_{\{j_0 < j \leq \omega(\alpha)\}}. \quad (2.7)$$

Here $j_0 = j_0(\alpha) = [\omega(\alpha) \varepsilon_n]$ with

$$\omega(\alpha) = \bar{\omega} + (A_\beta t)^{1/(2\beta+1)} n^{1/(2\beta+1)}, \quad (2.8)$$

where $\bar{\omega}$ is any nonnegative constant and

$$A_\beta = \frac{(\beta + 1)(2\beta + 1)}{\beta\pi^{2\beta}}.$$

Remark 2.2. *Note that the weighted least squares estimators (2.4) have been introduced by Pinsker, 1981, for continuous time optimal signal filtering in the gaussian noise. He proved that the mean-square asymptotic risk is minimized by weighted least squares estimators with weights of type (2.7). Moreover he has found the sharp minimal value of the mean-square asymptotic risk, which was called later as the Pinsker constant. Nussbaum, 1985, used the same method with proper modification for efficient estimation of the function S of known smoothness in the homogeneous gaussian model (1.1), i.e. when $\sigma_1(S) = \dots = \sigma_n(S) = 1$ and $(\xi_j)_{1 \leq j \leq n}$ is i.i.d. $\mathcal{N}(0, 1)$ sequence.*

To choose weights from the set (2.5) we minimize the special cost function introduced by Galtchouk, Pergamenshchikov, 2007. This cost function is as follows

$$J_n(\lambda) = \sum_{j=1}^n \lambda^2(j) \hat{\theta}_{j,n}^2 - 2 \sum_{j=1}^n \lambda(j) \tilde{\theta}_{j,n} + \rho \hat{P}_n(\lambda), \quad (2.9)$$

where

$$\tilde{\theta}_{j,n} = \hat{\theta}_{j,n}^2 - \frac{1}{n} \hat{\zeta}_n \quad \text{with} \quad \hat{\zeta}_n = \sum_{j=l_n+1}^n \hat{\theta}_{j,n}^2 \quad (2.10)$$

and $l_n = [n^{1/3} + 1]$. The penalty term we define as

$$\hat{P}_n(\lambda) = \frac{|\lambda|^2 \hat{\zeta}_n}{n}, \quad |\lambda|^2 = \sum_{j=1}^n \lambda^2(j) \quad \text{and} \quad \rho = \frac{1}{3 + L_n},$$

where $L_n \geq 0$ is any slowly increasing sequence, i.e.

$$\lim_{n \rightarrow \infty} L_n = +\infty \quad \text{and} \quad \lim_{n \rightarrow \infty} \frac{L_n}{n^\nu} = 0, \quad (2.11)$$

for any $\nu > 0$.

Finally, we set

$$\hat{\lambda} = \operatorname{argmin}_{\lambda \in \Lambda} J_n(\lambda) \quad \text{and} \quad \hat{S}_* = \hat{S}_{\hat{\lambda}}. \quad (2.12)$$

The goal of this paper is to study asymptotic (as $n \rightarrow \infty$) properties of this estimation procedure.

Remark 2.3. Now we explain why does one choose the cost function in the form (2.9). Developing the empiric quadratic risk for estimate (2.4), one obtains

$$\|\widehat{S}_\lambda - S\|_n^2 = \sum_{j=1}^n \lambda^2(j) \widehat{\theta}_{j,n}^2 - 2 \sum_{j=1}^n \lambda(j) \widehat{\theta}_{j,n} \theta_{j,n} + \|S\|_n^2.$$

It's natural to choose the weight vector λ for which this function reaches the minimum. Since the last term on the right-hand part is independent of λ , it can be dropped and one has to minimize with respect to λ the function equals the difference of two first terms on the right-hand part. It's clear that the minimization problem can't be solved directly because the Fourier coefficients $(\theta_{j,n})$ are unknown. To overcome this difficulty, we replace the product $\widehat{\theta}_{j,n} \theta_{j,n}$ by its asymptotically unbiased estimator $\widetilde{\theta}_{j,n}$ (see, Galtchouk, Pergamenshchikov, 2007, 2008). Moreover, to pay this substitution, we introduce into the cost function the penalty term \widehat{P}_n with a small coefficient $\rho > 0$. The form of the penalty term is provided by the principal term of the quadratic risk for weighted least-squares estimator, see Galtchouk, Pergamenshchikov, 2007, 2008. The coefficient $\rho > 0$ means, that the penalty is small, because the estimator $\widetilde{\theta}_{j,n}$ approximates in mean the quantity $\widehat{\theta}_{j,n} \theta_{j,n}$ asymptotically, as $n \rightarrow \infty$.

Note that the principal difference between the procedure (2.12) and the adaptive procedure proposed by Golubev, Nussbaum, 1993, for a homogeneous gaussian regression, consists in presence of the penalty term in the cost function (2.9).

Remark 2.4. As it was noted at Remark 2.2, Nussbaum, 1985, has shown that the weight coefficients of type (2.7) provide the asymptotic minimum of the mean-squared risk at the regression function estimation problem for the homogeneous gaussian model (1.1), when the smoothness of the function S is known. In fact, to obtain an efficient estimator one needs to take a weighted least squares estimator (2.4) with the weight vector λ_α , where the index α depends on smoothness of function S and on coefficients $(\sigma_j(S))_{1 \leq j \leq n}$, (see (5.2) below), which are unknown in our case. For this reason, Galtchouk, Pergamenshchikov, 2007, 2008, have proposed to make use of the family of coefficients (2.5), which contains the weight vector providing the minimum of the mean-squared risk. Moreover, they proposed the adaptive procedure (2.12) for which a non-asymptotic oracle inequality (see, Theorem 4.1 below)

was proved under some weak conditions on the coefficients $(\sigma_j(S))_{1 \leq j \leq n}$. It is important to note that due the properties of the parametric family (2.6), the secondary term in the oracle inequality is slowly increasing (slower than any degree of n).

3 Conditions

First we impose some conditions on unknown function S in the model (1.1).

Let $\mathcal{C}_{per,1}^k(\mathbb{R})$ be the set of 1-periodic k times differentiable $\mathbb{R} \rightarrow \mathbb{R}$ functions. We assume that S belongs to the following set

$$W_r^k = \{f \in \mathcal{C}_{per,1}^k(\mathbb{R}) : \sum_{j=0}^k \|f^{(j)}\|^2 \leq r\}, \quad (3.1)$$

where $\|\cdot\|$ denotes the norm in $\mathcal{L}_2[0,1]$, i.e.

$$\|f\|^2 = \int_0^1 f^2(t)dt. \quad (3.2)$$

Moreover, we suppose that $r > 0$ and $k \geq 1$ are unknown parameters.

Note that, we can represent the set W_r^k as an ellipse in $\mathcal{L}_2[0,1]$, i.e.

$$W_r^k = \{f \in \mathcal{L}_2[0,1] : \sum_{j=1}^{\infty} a_j \theta_j^2 \leq r\}, \quad (3.3)$$

where

$$\theta_j = (f, \phi_j) = \int_0^1 f(t)\phi_j(t)dt \quad (3.4)$$

and

$$a_j = \sum_{l=0}^k \|\phi_j^{(l)}\|^2 = \sum_{i=0}^k (2\pi[j/2])^{2i}. \quad (3.5)$$

Here $(\phi_j)_{j \geq 1}$ is the trigonometric basis defined in (2.1).

Now we describe the conditions on the scale coefficients $(\sigma_j(S))_{j \geq 1}$.

H₁) $\sigma_j(S) = g(x_j, S)$ for some unknown function $g : [0,1] \times \mathcal{L}_1[0,1] \rightarrow \mathbb{R}_+$, which is square integrable with respect to x such that

$$\lim_{n \rightarrow \infty} \sup_{S \in W_r^k} \left| \frac{1}{n} \sum_{j=1}^n g^2(x_j, S) - \varsigma(S) \right| = 0, \quad (3.6)$$

where $\varsigma(S) := \int_0^1 g^2(x, S)dx$. Moreover,

$$g_* = \inf_{0 \leq x \leq 1} \inf_{S \in W_r^k} g^2(x, S) > 0 \quad (3.7)$$

and

$$\sup_{S \in W_r^k} \varsigma(S) < \infty. \quad (3.8)$$

H₂) For any $x \in [0, 1]$, the operator $g^2(x, \cdot) : \mathcal{C}[0, 1] \rightarrow \mathbb{R}$ is differentiable in the Fréchet sense for any fixed function f_0 from $\mathcal{C}[0, 1]$, i.e. for any f from some vicinity of f_0 in $\mathcal{C}[0, 1]$,

$$g^2(x, f) = g^2(x, f_0) + \mathbf{L}_{x, f_0}(f - f_0) + \Upsilon(x, f_0, f),$$

where the Fréchet derivative $\mathbf{L}_{x, f_0} : \mathcal{C}[0, 1] \rightarrow \mathbb{R}$ is a bounded linear operator and the residual term $\Upsilon(x, f_0, f)$, for each $x \in [0, 1]$, satisfies the following property:

$$\lim_{\|f - f_0\|_\infty \rightarrow 0} \frac{|\Upsilon(x, f_0, f)|}{\|f - f_0\|_\infty} = 0,$$

where $\|f\|_\infty = \sup_{0 \leq t \leq 1} |f(t)|$.

H₃) There exists some positive constant C^* such that for any function S from $\mathcal{C}[0, 1]$ the operator $\mathbf{L}_{x, S}$ defined in the condition **H₂)** satisfies the following inequality for any function f from $\mathcal{C}[0, 1]$:

$$|\mathbf{L}_{x, S}(f)| \leq C^* (|S(x)f(x)| + |f|_1 + \|S\| \|f\|), \quad (3.9)$$

where $|f|_1 = \int_0^1 |f(t)|dt$.

H₄) The function $g_0(\cdot) = g(\cdot, S_0)$ corresponding to $S_0 \equiv 0$ is continuous on the interval $[0, 1]$. Moreover,

$$\lim_{\delta \rightarrow 0} \sup_{0 \leq x \leq 1} \sup_{\|S\|_\infty \leq \delta} |g(x, S) - g(x, S_0)| = 0.$$

Remark 3.1. Let us explain the conditions **H₁)**–**H₄)**. In fact, this is the regularity conditions of the function $g(x, S)$ generating the scale coefficients $(\sigma_j(S))_{1 \leq j \leq n}$.

Condition \mathbf{H}_1) means that the function $g(\cdot, S)$ should be uniformly integrable with respect to the first argument in the sense of convergence (3.6). Moreover, this function should be separated from zero (see inequality (3.7)) and bounded on the class (3.1) (see inequality (3.8)). Boundedness away from zero provides that the distribution of observations $(y_j)_{1 \leq j \leq n}$ isn't degenerate in \mathbb{R}^n , and the boundedness means that the intensity of the noise vector should be finite, otherwise the estimation problem hasn't any sense.

Conditions \mathbf{H}_2) and \mathbf{H}_3) mean that the function $g(x, \cdot)$ is regular, at any fixed $0 \leq x \leq 1$, with respect to S in the sense, that it is differentiable in the Fréchet sense (see e.g., Kolmogorov, Fomin, 1989) and moreover the Fréchet derivative satisfies the growth condition given by the inequality (3.9) which permits to consider the example (1.2).

Last the condition \mathbf{H}_4) is the usual uniform continuity the condition of the function $g(\cdot, \cdot)$ at the function S_0 .

Now we give some examples of functions satisfying the conditions \mathbf{H}_1)- \mathbf{H}_4).

We set

$$g^2(x, S) = c_0 + c_1 x + c_2 S^2(x) + c_3 \int_0^1 S^2(t) dt \quad (3.10)$$

with some coefficients $c_0 > 0$, $c_i \geq 0$, $i = 1, 2, 3$.

In this case

$$\varsigma(S) = c_0 + \frac{c_1}{2} + (c_2 + c_3) \int_0^1 S^2(t) dt.$$

The Fréchet derivative is given by

$$\mathbf{L}_{x,S}(f) = 2S(x)f(x) + 2 \int_0^1 S(t)f(t) dt.$$

It is easy to see that the function (3.10) satisfies the conditions \mathbf{H}_1)- \mathbf{H}_4). Moreover, the conditions \mathbf{H}_1)- \mathbf{H}_4) are satisfied by any function of type

$$g^2(x, S) = G(x, S(x)) + \int_0^1 V(S(t)) dt, \quad (3.11)$$

where the functions G and V satisfy the following the conditions:

- G is a $[0, 1] \times \mathbb{R} \rightarrow [c_0, +\infty)$ function (with $c_0 > 0$) such that

$$\lim_{\delta \rightarrow 0} \max_{|u-v| \leq \delta} \sup_{y \in \mathbb{R}} |G(u, y) - G(v, y)| = 0 \quad (3.12)$$

and

$$m_1 = \sup_{0 \leq x \leq 1} \sup_{y \in \mathbb{R}} \frac{|G_y(x, y)|}{|y|} < \infty; \quad (3.13)$$

- V is a continuously differentiable $\mathbb{R} \rightarrow \mathbb{R}_+$ function such that

$$m_2 = \sup_{y \in \mathbb{R}} \frac{|\dot{V}(y)|}{1 + |y|} < \infty,$$

where $\dot{V}(\cdot)$ is the derivative of V .

In this case

$$\varsigma(S) = \int_0^1 G(t, S(t)) dt + \int_0^1 V(S(t)) dt$$

and

$$\begin{aligned} \left| n^{-1} \sum_{j=1}^n g^2(x_j, S) - \varsigma(S) \right| &\leq \sum_{j=1}^n \int_{x_{j-1}}^{x_j} |G(x_j, S(x_j)) - G(t, S(t))| dt \\ &\leq \Delta_n + \sum_{j=1}^n \int_{x_{j-1}}^{x_j} |G(t, S(x_j)) - G(t, S(t))| dt, \end{aligned}$$

where $\Delta_n = \max_{|u-v| \leq 1/n} \sup_{y \in \mathbb{R}} |G(u, y) - G(v, y)|$. Now to estimate the last term in this inequality note that

$$G(t, S(x_j)) - G(t, S(t)) = \int_t^{x_j} G_y(t, S(z)) \dot{S}(z) dz.$$

Therefore, from the condition (3.13) we get

$$|G(t, S(x_j)) - G(t, S(t))| \leq m_1 \int_{x_{j-1}}^{x_j} |S(z)| |\dot{S}(z)| dz,$$

and through the Bounyakovskii-Cauchy-Schwarz inequality, for any $S \in W_r^k$,

$$\begin{aligned} \left| n^{-1} \sum_{j=1}^n g^2(x_j, S) - \varsigma(S) \right| &\leq \sum_{j=1}^n \int_{x_{j-1}}^{x_j} |G(x_j, S(x_j)) - G(t, S(t))| dt \\ &\leq \Delta_n + \frac{m_1}{n} \int_0^1 |S(t)| |\dot{S}(t)| dt \\ &\leq \Delta_n + \frac{m_1}{n} \|S\| \|\dot{S}\| \leq \Delta_n + \frac{m_1}{n} r. \end{aligned}$$

Now, the condition (3.12) implies \mathbf{H}_1 .

Moreover, the Fréchet derivative in this case is given by

$$\mathbf{L}_{x,S}(f) = G_y(x, S(x))f(x) + \int_0^1 \dot{V}(S(t))f(t)dt.$$

One can check directly that this operator satisfies the inequality (3.9) with $C^* = m_1 + m_2$.

4 Main results

Denote by \mathcal{P}_n the family of distributions p in \mathbb{R}^n of the vectors $(\xi_1, \dots, \xi_n)'$ in the model (1.1) such that the components ξ_j are jointly independent, centered with unit variance and

$$\max_{1 \leq k \leq n} \mathbf{E} \xi_k^4 \leq l_n^*, \quad (4.1)$$

where $l_n^* \geq 3$ is slowly increasing sequence, that is it satisfies the property (2.11).

It is easy to see that, for any $n \geq 1$, the centered gaussian distribution in \mathbb{R}^n with unit covariation matrix belongs to the family \mathcal{P}_n . We will denote by q this gaussian distribution.

For any estimator \widehat{S} we define the following quadratic risk

$$\mathcal{R}_n(\widehat{S}, S) = \sup_{p \in \mathcal{P}_n} \mathbf{E}_{S,p} \|\widehat{S} - S\|_n^2, \quad (4.2)$$

where $\mathbf{E}_{S,p}$ is the expectation with respect to the distribution $\mathbf{P}_{S,p}$ of the observations (y_1, \dots, y_n) with the fixed function S and the fixed distribution $p \in \mathcal{P}_n$ of random variables $(\xi_j)_{1 \leq j \leq n}$ in the model (1.1).

Moreover, to make the risk independent of the design points, in this paper we will make use of the risk with respect to the usual norm in $\mathcal{L}_2[0, 1]$ (3.2) also, i.e.

$$\mathcal{T}_n(\widehat{S}, S) = \sup_{p \in \mathcal{P}_n} \mathbf{E}_{S,p} \|\widehat{S} - S\|^2. \quad (4.3)$$

If an estimator \widehat{S} is defined only at the design points $(x_j)_{1 \leq j \leq n}$, then we extend it as step function onto the interval $[0, 1]$ by setting $\widehat{S}(x) = T(\widehat{S}(x))$, for all $0 \leq x \leq 1$, where

$$T(f)(x) = f(x_1) \mathbf{1}_{[0, x_1]}(x) + \sum_{k=2}^n f(x_k) \mathbf{1}_{(x_{k-1}, x_k]}(x). \quad (4.4)$$

In Galtchouk, Pergamenshchikov, 2007, 2008 the following non-asymptotic oracle inequality has been shown for the procedure (2.12) .

Theorem 4.1. *Assume that in the model (1.1) the function S belongs to W_r^1 . Then, for any odd $n \geq 3$ and $r > 0$, the estimator \widehat{S}_* satisfies the following oracle inequality*

$$\mathcal{R}_n(\widehat{S}_*, S) \leq \frac{1 + 3\rho - 2\rho^2}{1 - 3\rho} \min_{\lambda \in \Lambda} \mathcal{R}_n(\widehat{S}_\lambda, S) + \frac{1}{n} \mathcal{B}_n(\rho), \quad (4.5)$$

where the function $\mathcal{B}_n(\rho)$ is such that, for any $\nu > 0$,

$$\lim_{n \rightarrow \infty} \frac{\mathcal{B}_n(\rho)}{n^\nu} = 0. \quad (4.6)$$

Remark 4.1. *Note that in Galtchouk, Pergamenshchikov, 2007, 2008, the oracle inequality is proved for the model (1.1), where the random variables $(\xi_j)_{1 \leq j \leq n}$ are independent identically distributed. In fact, the result and the proof are true for independent random variables which are not identically distributed, i.e. for any distribution of the random vector $(\xi_1, \dots, \xi_n)'$ from \mathcal{P}_n .*

Now we formulate the main asymptotic results. To this end, for any function $S \in W_r^k$, we set

$$\gamma_k(S) = \Gamma_k^* r^{\frac{1}{2k+1}} (\varsigma(S))^{2k/(2k+1)}, \quad (4.7)$$

where

$$\Gamma_k^* = (2k + 1)^{\frac{1}{2k+1}} (k/(\pi(k + 1)))^{2k/(2k+1)}.$$

It is well known (see e.g., Nussbaum, 1985) that the optimal rate of convergence is $n^{2k/(2k+1)}$ when the risk is taken uniformly over W_r^k .

Theorem 4.2. *Assume that in the model (1.1) the sequence $(\sigma_j(S))$ fulfills the condition \mathbf{H}_1). Then the estimator \widehat{S}_* from (2.12) satisfies the inequalities*

$$\limsup_{n \rightarrow \infty} n^{\frac{2k}{2k+1}} \sup_{S \in W_r^k} \frac{\mathcal{R}_n(\widehat{S}_*, S)}{\gamma_k(S)} \leq 1 \quad (4.8)$$

and

$$\limsup_{n \rightarrow \infty} n^{\frac{2k}{2k+1}} \sup_{S \in W_r^k} \frac{\mathcal{T}_n(\widehat{S}_*, S)}{\gamma_k(S)} \leq 1. \quad (4.9)$$

The following result gives the sharp lower bound for risk (4.2) and show that $\gamma_k(S)$ is the *Pinsker* constant.

Theorem 4.3. *Assume that in the model (1.1) the sequence $(\sigma_j(S))$ satisfies the conditions \mathbf{H}_2)– \mathbf{H}_4). Then the risks (4.2) and (4.3) admit the following asymptotic lower bounds*

$$\liminf_{n \rightarrow \infty} n^{\frac{2k}{2k+1}} \inf_{\widehat{S}_n} \sup_{S \in W_r^k} \frac{\mathcal{R}_n(\widehat{S}_n, S)}{\gamma_k(S)} \geq 1 \quad (4.10)$$

and

$$\liminf_{n \rightarrow \infty} n^{\frac{2k}{2k+1}} \inf_{\widehat{S}_n} \sup_{S \in W_r^k} \frac{\mathcal{T}_n(\widehat{S}_n, S)}{\gamma_k(S)} \geq 1. \quad (4.11)$$

Remark 4.2. *Note that in Galtchouk, Pergamenshchikov, 2005, an asymptotically efficient estimator has been constructed and results similar to Theorems 4.2 and 4.3 were claimed for the model (1.1). In fact the upper bound is true there under some additional condition on the smoothness of the function S , i.e. on the parameter k . In the cited paper this additional condition is not formulated since erroneous inequality (A.6). To avoid using this inequality we modify the estimating procedure by introducing the penalty term $\rho \widehat{P}_n(\lambda)$ in the cost function (2.9). By this way we remove all additional conditions on the smoothness parameter k .*

Remark 4.3. *In fact to obtain the non-asymptotic oracle inequality (4.5), it isn't necessary to make use of equidistant design points and the trigonometric basis. One may take any design points (deterministic or random) and*

any orthonormal basis satisfying (2.2). But to obtain the property (4.6) one needs to impose some technical conditions (see Galtchouk, Pergamenshchikov, 2008).

Note that the results of Theorem 4.2 and Theorem 4.3 are based on equidistant design points and the trigonometric basis.

5 Upper bound

In this section we prove Theorem 4.2. To this end we will make use of the oracle inequality (4.5). We have to find an estimator from the family (2.4)-(2.5) for which we can show the upper bound (4.8). We start with the construction of such an estimator. First we put

$$\tilde{l}_n = \inf\{i \geq 1 : i\varepsilon \geq \bar{r}(S)\} \wedge m \quad \text{and} \quad \bar{r}(S) = r/\zeta(S), \quad (5.1)$$

where $a \wedge b = \min(a, b)$.

Then we choose an index from the set \mathcal{A}_ε as

$$\tilde{\alpha} = (k, \tilde{t}_n), \quad (5.2)$$

where k is the parameter of the set W_r^k and $\tilde{t}_n = \tilde{l}_n \varepsilon$. Finally, we set

$$\tilde{S} = \hat{S}_{\tilde{\lambda}} \quad \text{and} \quad \tilde{\lambda} = \lambda_{\tilde{\alpha}}. \quad (5.3)$$

Now we show the upper bound (4.8) for this estimator.

Theorem 5.1. *Assume that the condition \mathbf{H}_1) holds. Then*

$$\limsup_{n \rightarrow \infty} n^{\frac{2k}{2k+1}} \sup_{S \in W_r^k} \frac{\mathcal{R}_n(\tilde{S}, S)}{\gamma_k(S)} \leq 1. \quad (5.4)$$

Remark 5.1. *Note that the estimator \tilde{S} belongs to the family (2.4)-(2.5), but we can't use directly this estimator because the parameters k , r and $\bar{r}(S)$ are unknown. We can use this upper bound only through the oracle inequality (4.5) proved for procedure (2.12).*

Now Theorem 4.1 and Theorem 5.1 imply the upper bound (4.8). To obtain the upper bound (4.9) we need the following auxiliary result.

Lemma 5.2. For any $0 < \delta < 1$ and any estimate \widehat{S}_n of $S \in W_r^k$,

$$\|\widehat{S}_n - S\|_n^2 \geq (1 - \delta)\|T_n(\widehat{S}) - S\|^2 - (\delta^{-1} - 1)r/n^2,$$

where the function $T_n(\widehat{S})(\cdot)$ is defined in (4.4).

Proof of this Lemma is given in Appendix A.1.

Now inequality (4.8) and this lemma imply the upper bound (4.9). Hence Theorem 4.2.

6 Lower bound

In this section we give the main steps of proving the lower bounds (4.10) and (4.11). In common, we follow the same scheme as Nussbaum, 1985. We begin with minorizing the minimax risk by a bayesian one constructed on a parametric functional family introduced in Section 6.2 (see (6.9)) and using the prior distribution (6.10). Further, a special modification of the van Trees inequality (see, Theorem 6.1) yields a lower bound for the bayesian risk depending on the chosen prior distribution, of course. Finally, in section 6.3, we choose parameters of the prior distribution (see (6.10)) providing the maximal value of the lower bound for the bayesian risk. This value coincides with the Pinsker constant as it is shown in Section 9.2.

6.1 Lower bound for parametric heteroscedastic regression models

Let $(\mathbb{R}^n, \mathcal{B}(\mathbb{R}^n), \mathbf{P}_\vartheta, \vartheta \in \Theta \subseteq \mathbb{R}^l)$ be a statistical model relative to the observations $(y_j)_{1 \leq j \leq n}$ governed by the regression equation

$$y_j = S_\vartheta(x_j) + \sigma_j(\vartheta) \xi_j, \quad (6.1)$$

where ξ_1, \dots, ξ_n are i.i.d. $\mathcal{N}(0, 1)$ random variables, $\vartheta = (\vartheta_1, \dots, \vartheta_l)'$ is a unknown parameter vector, $S_\vartheta(x)$ is a unknown (or known) function and $\sigma_j(\vartheta) = g(x_j, S_\vartheta)$, with the function $g(x, S)$ defined in the condition \mathbf{H}_1 . Assume that a prior distribution μ_ϑ of the parameter ϑ in \mathbb{R}^l is defined by the density $\Phi(\cdot)$ of the following form

$$\Phi(z) = \Phi(z_1, \dots, z_l) = \prod_{i=1}^l \varphi_i(z_i),$$

where φ_i is a continuously differentiable bounded density on \mathbb{R} with

$$I_i = \int_{\mathbb{R}} \frac{\dot{\varphi}_i^2(u)}{\varphi_i(u)} du < \infty.$$

Let $\tau(\cdot)$ be a continuously differentiable $\mathbb{R}^l \rightarrow \mathbb{R}$ function such that, for any $1 \leq i \leq l$,

$$\lim_{|z_i| \rightarrow \infty} \tau(z) \varphi_i(z_i) = 0 \quad \text{and} \quad \int_{\mathbb{R}^l} |\tau'_i(z)| \Phi(z) dz < \infty, \quad (6.2)$$

where

$$\tau'_i(z) = (\partial/\partial z_i) \tau(z).$$

Let $\hat{\tau}_n$ be an estimator of $\tau(\vartheta)$ based on observations $(y_j)_{1 \leq j \leq n}$. For any $\mathcal{B}(\mathbb{R}^n \times \mathbb{R}^l)$ -measurable integrable function $G(x, z)$, $x \in \mathbb{R}^n$, $z \in \mathbb{R}^l$, we set

$$\tilde{\mathbf{E}} G(Y, \vartheta) = \int_{\mathbb{R}^l} \mathbf{E}_z G(Y, z) \Phi(z) dz,$$

where \mathbf{E}_ϑ is the expectation with respect to the distribution \mathbf{P}_ϑ of the vector $Y = (y_1, \dots, y_n)$. Note that in this case

$$\mathbf{E}_\vartheta G(Y, \vartheta) = \int_{\mathbb{R}^n} G(v, \vartheta) f(v, \vartheta) dv,$$

where

$$f(v, z) = \prod_{j=1}^n \frac{1}{\sqrt{2\pi}\sigma_j(z)} \exp \left\{ -\frac{(v_j - S_z(x_j))^2}{2\sigma_j^2(z)} \right\}. \quad (6.3)$$

We prove the following result.

Theorem 6.1. *Assume that the conditions $\mathbf{H}_1) - \mathbf{H}_2)$ hold. Moreover, assume that the function $S_z(\cdot)$ with $z = (z_1, \dots, z_l)'$ is uniformly over $0 \leq x \leq 1$ differentiable with respect to z_i , $1 \leq i \leq l$, i.e. for any $1 \leq i \leq l$ there exists a function $S'_{z,i} \in \mathcal{C}[0, 1]$ such that*

$$\lim_{h \rightarrow 0} \max_{0 \leq x \leq 1} \left| \left(S_{z+h\mathbf{e}_i}(x) - S_z(x) - S'_{z,i}(x)h \right) / h \right| = 0, \quad (6.4)$$

where $\mathbf{e}_i = (0, \dots, 1, \dots, 0)'$, all coordinates are 0, except the i -th equals to 1. Then for any square integrable estimator $\hat{\tau}_n$ of $\tau(\vartheta)$ and any $1 \leq i \leq l$,

$$\tilde{\mathbf{E}} (\hat{\tau}_n - \tau(\vartheta))^2 \geq \frac{\overline{\tau}_i^2}{F_i + B_i + I_i}, \quad (6.5)$$

where $\bar{\tau}_i = \int_{\mathbb{R}^l} \tau'_i(z) \Phi(z) dz$, $F_i = \sum_{j=1}^n \int_{\mathbb{R}^l} (S'_{z,i}(x_j)/\sigma_j(z))^2 \Phi(z) dz$ and

$$B_i = \frac{1}{2} \sum_{j=1}^n \int_{\mathbb{R}^l} \frac{\tilde{\mathbf{L}}_i^2(x_j, S_z)}{\sigma_j^4(S_z)} \Phi(z) dz,$$

$\tilde{\mathbf{L}}_i(x, z) = \mathbf{L}_{x, S_z}(S'_{z,i})$, the operator $\mathbf{L}_{x, S}$ is defined in the condition **H**₂).

Proof is given in Appendix A.2.

Remark 6.1. Note that the inequality (6.5) is some modification of the van Trees inequality (see, Gill, Levit, 1995) adapted to the model (6.1).

6.2 Parametric family of kernel functions

In this section we define and study some special parametric family of kernel function which will be used to prove the sharp lower bound (4.10).

Let us begin by kernel functions. We fix $\eta > 0$ and we set

$$\chi_\eta(x) = \eta^{-1} \int_{\mathbb{R}} \mathbf{1}_{(|u| \leq 1 - \eta)} V\left(\frac{u - x}{\eta}\right) du, \quad (6.6)$$

where $\mathbf{1}_A$ is the indicator of a set A , the kernel $V \in \mathcal{C}^\infty(\mathbb{R})$ is such that

$$V(u) = 0 \quad \text{for} \quad |u| \geq 1 \quad \text{and} \quad \int_{-1}^1 V(u) du = 1.$$

It is easy to see that the function $\chi_\eta(x)$ possesses the properties :

$$\begin{aligned} 0 \leq \chi_\eta \leq 1, \quad \chi_\eta(x) = 1 \quad \text{for} \quad |x| \leq 1 - 2\eta \quad \text{and} \\ \chi_\eta(x) = 0 \quad \text{for} \quad |x| \geq 1. \end{aligned}$$

Moreover, for any $c > 0$ and $\nu \geq 0$

$$\lim_{\eta \rightarrow 0} \sup_{f: \|f\|_\infty \leq c} \left| \int_{\mathbb{R}} f(x) \chi_\eta^\nu(x) dx - \int_{-1}^1 f(x) dx \right| = 0. \quad (6.7)$$

We divide the interval $[0, 1]$ into M equal subintervals of length $2h$ and on each of them we construct a kernel-type function which equals to zero at the boundary of the subinterval together with all derivatives.

It provides that the Fourier partial sums with respect to the trigonometric basis in $\mathcal{L}_2[-1, 1]$ give a natural parametric approximation to the function on each subinterval.

Let $(e_j)_{j \geq 1}$ be the trigonometric basis in $\mathcal{L}_2[-1, 1]$, i.e.

$$e_1 = 1/\sqrt{2}, \quad e_j(x) = \text{Tr}_j(\pi[j/2]x), \quad j \geq 2, \quad (6.8)$$

where the functions $(\text{Tr}_j)_{j \geq 2}$ are defined in (2.1).

Now, for any array $z = \{(z_{m,j})_{1 \leq m \leq M_n, 1 \leq j \leq N_n}\}$ we define the following function

$$S_{z,n}(x) = \sum_{m=1}^{M_n} \sum_{j=1}^{N_n} z_{m,j} D_{m,j}(x), \quad (6.9)$$

where $D_{m,j}(x) = e_j(v_m(x)) \chi_\eta(v_m(x))$,

$$v_m(x) = (x - \tilde{x}_m)/h_n, \quad \tilde{x}_m = 2mh_n \quad \text{and} \quad M_n = [1/(2h_n)] - 1.$$

We assume that the sequences $(N_n)_{n \geq 1}$ and $(h_n)_{n \geq 1}$, satisfy the following conditions.

A₁) The sequence $N_n \rightarrow \infty$ as $n \rightarrow \infty$ and for any $\nu > 0$

$$\lim_{n \rightarrow \infty} N_n^\nu / n = 0.$$

Moreover, there exist $0 < \delta_1 < 1$ and $\delta_2 > 0$ such that

$$h_n = O(n^{-\delta_1}) \quad \text{and} \quad h_n^{-1} = O(n^{\delta_2}) \quad \text{as} \quad n \rightarrow \infty.$$

To define a prior distribution on the family of arrays, we choose the following random array $\vartheta = \{(\vartheta_{m,j})_{1 \leq m \leq M_n, 1 \leq j \leq N_n}\}$ with

$$\vartheta_{m,j} = t_{m,j} \zeta_{m,j}, \quad (6.10)$$

where $(\zeta_{m,j})$ are i.i.d. $\mathcal{N}(0, 1)$ random variables and $(t_{m,j})$ are some non-random positive coefficients. We make use of gaussian variables since they possess the minimal Fisher information and therefore maximize the lower bound (6.5). We set

$$t_n^* = \max_{1 \leq m \leq M_n} \sum_{j=1}^{N_n} t_{m,j}. \quad (6.11)$$

We assume that the coefficients $(t_{m,j})_{1 \leq m \leq M_n, 1 \leq j \leq N_n}$ satisfy the following conditions.

A₂) *There exists a sequence of positive numbers $(d_n)_{n \geq 1}$ such that*

$$\lim_{n \rightarrow \infty} \frac{d_n}{h_n^{2k-1}} \sum_{m=1}^{M_n} \sum_{j=1}^{N_n} t_{m,j}^2 j^{2(k-1)} = 0, \quad \lim_{n \rightarrow \infty} \sqrt{d_n} t_n^* = 0, \quad (6.12)$$

moreover, for any $\nu > 0$,

$$\lim_{n \rightarrow \infty} n^\nu \exp\{-d_n/2\} = 0.$$

A₃) *For some $0 < \epsilon < 1$*

$$\limsup_{n \rightarrow \infty} \frac{1}{h_n^{2k-1}} \sum_{m=1}^{M_n} \sum_{j=1}^{N_n} t_{m,j}^2 j^{2k} \leq (1 - \epsilon)r \left(\frac{2}{\pi}\right)^{2k}.$$

A₄) *There exists $\epsilon_0 > 0$ such that*

$$\lim_{n \rightarrow \infty} \frac{1}{h_n^{4k-2+\epsilon_0}} \sum_{m=1}^{M_n} \sum_{j=1}^{N_n} t_{m,j}^4 j^{4k} = 0.$$

Proposition 6.2. *Let the conditions **A₁**)–**A₂**). Then, for any $\nu > 0$ and for any $\delta > 0$,*

$$\lim_{n \rightarrow \infty} n^\nu \max_{0 \leq l \leq k-1} \mathbf{P} \left(\|S_{\vartheta,n}^{(l)}\| > \delta \right) = 0.$$

Proposition 6.3. *Let the conditions **A₁**)–**A₄**). Then, for any $\nu > 0$,*

$$\lim_{n \rightarrow \infty} n^\nu \mathbf{P}(S_{\vartheta,n} \notin W_r^k) = 0.$$

Proposition 6.4. *Let the conditions **A₁**)–**A₄**). Then, for any $\nu > 0$,*

$$\lim_{n \rightarrow \infty} n^\nu \mathbf{E} \|S_{\vartheta,n}\|^2 \left(\mathbf{1}_{\{S_{\vartheta,n} \notin W_r^k\}} + \mathbf{1}_{\Xi_n^c} \right) = 0.$$

Proposition 6.5. *Let the conditions **A₁**)–**A₄**). Then for any function g satisfying the conditions (3.7) and **H₄**)*

$$\lim_{n \rightarrow \infty} \sup_{0 \leq x \leq 1} \mathbf{E} |g^{-2}(x, S_{\vartheta,n}) - g_0^{-2}(x)| = 0.$$

Proofs of Propositions 6.2–6.5 are given in Appendix.

6.3 Bayes risk

Now we will obtain the lower bound for the bayesian risk that yields the lower bound (4.11) for the minimax risk.

We make use of the sequence of random functions $(S_{\vartheta,n})_{n \geq 1}$ defined in (6.9)-(6.10) with the coefficients $(t_{m,j})$ satisfying the conditions \mathbf{A}_1 - \mathbf{A}_4) which will be chosen later.

For any estimator \widehat{S}_n we introduce now the corresponding Bayes risk

$$\mathcal{E}_n(\widehat{S}_n) = \int_{\mathbb{R}^l} \mathbf{E}_{S_{z,n},q} \|\widehat{S}_n - S_{z,n}\|^2 \mu_{\vartheta}(\mathrm{d}z), \quad (6.13)$$

where the kernel family $(S_{z,n})$ is defined in (6.9), μ_{ϑ} denotes the distribution of the random array ϑ defined by (6.10) in \mathbb{R}^l with $l = M_n N_n$.

We remember that q is a centered gaussian distribution in \mathbb{R}^n with unit covariation matrix.

First of all, we replace the functions \widehat{S}_n and S by their Fourier series with respect to the basis

$$\tilde{e}_{m,i}(x) = (1/\sqrt{h}) e_i(v_m(x)) \mathbf{1}_{(|v_m(x)| \leq 1)}.$$

By making use of this basis we can estimate the norm $\|\widehat{S}_n - S_{z,n}\|^2$ from below as

$$\|\widehat{S}_n - S_{z,n}\|^2 \geq \sum_{m=1}^{M_n} \sum_{j=1}^{N_n} (\widehat{\tau}_{m,j} - \tau_{m,j}(z))^2,$$

where

$$\widehat{\tau}_{m,j} = \int_0^1 \widehat{S}_n(x) \tilde{e}_{m,j}(x) \mathrm{d}x \quad \text{and} \quad \tau_{m,j}(z) = \int_0^1 S_{z,n}(x) \tilde{e}_{m,j}(x) \mathrm{d}x.$$

Moreover, from the definition (6.9) one gets

$$\tau_{m,j}(z) = \sqrt{h} \sum_{i=1}^{N_n} z_{m,i} \int_{-1}^1 e_i(u) e_j(u) \chi_{\eta}(u) \mathrm{d}u.$$

It is easy to see that the functions $\tau_{m,j}(\cdot)$ satisfy the condition (6.2) for gaussian prior densities. In this case (see the definition in (6.5)) we have

$$\bar{\tau}_{m,j} = (\partial/\partial z_{m,j}) \tau_{m,j}(z) = \sqrt{h} \bar{e}_j(\chi_{\eta}),$$

where

$$\bar{e}_j(f) = \int_{-1}^1 e_j^2(v) f(v) dv. \quad (6.14)$$

Now to obtain a lower bound for the Bayes risk $\mathcal{E}_n(\hat{S}_n)$ we make use of Theorem 6.1 which implies that

$$\mathcal{E}_n(\hat{S}_n) \geq \sum_{m=1}^{M_n} \sum_{j=1}^{N_n} \frac{h \bar{e}_j^2(\chi_\eta)}{F_{m,j} + B_{m,j} + t_{m,j}^{-2}}, \quad (6.15)$$

where $F_{m,j} = \sum_{i=1}^n D_{m,j}^2(x_i) \mathbf{E} g^{-2}(x_i, S_{\vartheta,n})$ and

$$B_{m,j} = \frac{1}{2} \sum_{i=1}^n \mathbf{E} \frac{\tilde{\mathbf{L}}_{m,j}^2(x_i, S_{\vartheta,n})}{g^4(x_i, S_{\vartheta,n})}$$

with $\tilde{\mathbf{L}}_{m,j}(x, S) = \mathbf{L}_{x,S}(D_{m,j})$. In the Appendix we show that

$$\lim_{n \rightarrow \infty} \sup_{1 \leq m \leq M_n} \sup_{1 \leq j \leq N_n} \left| \frac{1}{nh} F_{m,j} - \bar{e}_j(\chi_\eta^2) g_0^{-2}(\tilde{x}_m) \right| = 0 \quad (6.16)$$

and

$$\lim_{n \rightarrow \infty} \frac{1}{nh} \sup_{1 \leq m \leq M_n} \sup_{1 \leq j \leq N_n} B_{m,j} = 0. \quad (6.17)$$

This means that, for any $\nu > 0$ and for sufficiently large n ,

$$\sup_{1 \leq m \leq M_n} \sup_{1 \leq j \leq N_n} \frac{F_{m,j} + B_{m,j} + t_{m,j}^{-2}}{nh \bar{e}_j(\chi_\eta^2) g_0^{-2}(\tilde{x}_m) + t_{m,j}^{-2}} \leq 1 + \nu,$$

where \tilde{x}_m is defined in (6.9). Therefore, if we denote in (6.15)

$$\kappa_{m,j}^2 = nh g_0^{-2}(\tilde{x}_m) t_{m,j}^2 \quad \text{and} \quad \psi_j(\eta, y) = \frac{\bar{e}_j^2(\chi_\eta) y}{\bar{e}_j^2(\chi_\eta^2) y + 1}$$

we obtain, for sufficiently large n ,

$$n^{\frac{2k}{2k+1}} \mathcal{E}_n(\hat{S}_n) \geq \frac{n^{-\frac{1}{2k+1}}}{1 + \nu} \sum_{m=1}^{M_n} g_0^2(\tilde{x}_m) \sum_{j=1}^{N_n} \psi_j(\eta, \kappa_{m,j}^2).$$

In the Appendix we show that

$$\lim_{\eta \rightarrow 0} \sup_{N \geq 1} \sup_{(y_1, \dots, y_N) \in \mathbb{R}_+^N} \left| \frac{\sum_{j=1}^N \psi_j(\eta, y_j)}{\Psi_N(y_1, \dots, y_N)} - 1 \right| = 0, \quad (6.18)$$

where

$$\Psi_N(y_1, \dots, y_N) = \sum_{j=1}^N \frac{y_j}{y_j + 1}.$$

Therefore we can write that, for sufficiently large n ,

$$n^{\frac{2k}{2k+1}} \mathcal{E}_n(\widehat{S}_n) \geq \frac{1-\nu}{1+\nu} n^{-\frac{1}{2k+1}} \sum_{m=1}^{M_n} g_0^2(\tilde{x}_m) \Psi_{N_n}(\kappa_{m,1}^2, \dots, \kappa_{m,N_n}^2). \quad (6.19)$$

Obviously, to obtain a "good" lower bound for the risk $\mathcal{E}_n(\widehat{S}_n)$ one needs to maximize the right-hand side of the inequality (6.19). Hence we choose the coefficients $(\kappa_{m,j}^2)$ by maximization the function Ψ_N , i.e.

$$\max_{y_1, \dots, y_N} \Psi_N(y_1, \dots, y_N) \quad \text{subject to} \quad \sum_{j=1}^N y_j j^{2k} \leq R.$$

The parameter $R > 0$ will be chosen later to satisfy the condition \mathbf{A}_3). By the Lagrange multipliers method it is easy to find that the solution of this problem is given by

$$y_j^*(R) = a^*(R) j^{-k} - 1 \quad (6.20)$$

with

$$a^*(R) = \frac{1}{\sum_{i=1}^N i^k} \left(R + \sum_{i=1}^N i^{2k} \right) \quad \text{and} \quad 1 \leq j \leq N.$$

To obtain a positive solution in (6.20) we need to impose the following condition

$$R > N^k \sum_{i=1}^N i^k - \sum_{i=1}^N i^{2k}. \quad (6.21)$$

Moreover, from the condition \mathbf{A}_3) we obtain that

$$R \leq \frac{2^{2k+1}(1-\varepsilon) r n h_n^{2k+1}}{\pi^{2k} \widehat{g}_0} := R_n^*, \quad (6.22)$$

where

$$\widehat{g}_0 = 2h_n \sum_{m=1}^{M_n} g_0^2(\tilde{x}_m).$$

Note that by the condition \mathbf{H}_4) the function $g_0(\cdot) = g(\cdot, S_0)$ is continuous on the interval $[0, 1]$, therefore

$$\lim_{n \rightarrow \infty} \widehat{g}_0 = \int_0^1 g^2(x, S_0) dx = \varsigma(S_0) \quad \text{with} \quad S_0 \equiv 0. \quad (6.23)$$

Now we have to choose the sequence (h_n) . Note that if we put in (6.10)

$$t_{m,j} = g_0(\tilde{x}_m) \sqrt{y_j^*(R)} / \sqrt{nh_n} \quad \text{i.e.} \quad \kappa_{m,j} = y_j^*(R), \quad (6.24)$$

we can rewrite the inequality (6.19) as

$$n^{\frac{2k}{2k+1}} \mathcal{E}_n(\widehat{S}_n) \geq \frac{(1-\nu)}{(1+\nu)} \frac{\widehat{g}_0 \Psi_{N_n}^*(R)}{2h_n} n^{-\frac{1}{2k+1}}, \quad (6.25)$$

where

$$\Psi_N^*(R) = N - \frac{\left(\sum_{j=1}^N j^k\right)^2}{R + \sum_{j=1}^N j^{2k}}.$$

It is clear that

$$k^2/(k+1)^2 \leq \liminf_{N \rightarrow \infty} \inf_{R>0} \Psi_N^*(R)/N \leq \limsup_{N \rightarrow \infty} \sup_{R>0} \Psi_N^*(R)/N \leq 1.$$

Therefore to obtain a positive finite asymptotic lower bound in (6.25) we have to take the parameter h_n as

$$h_n = h_* n^{-1/(2k+1)} N_n \quad (6.26)$$

with some positive coefficient h_* . Moreover, the conditions (6.21)-(6.22) imply that, for sufficiently large n ,

$$(1-\varepsilon)r \frac{2^{2k+1}}{\pi^{2k}} \frac{1}{\widehat{g}_0} h_*^{2k+1} > \frac{1}{N_n^{k+1}} \sum_{j=1}^{N_n} j^k - \frac{1}{N_n^{2k+1}} \sum_{j=1}^{N_n} j^{2k}.$$

Moreover, taking into account that for sufficiently large n

$$\widehat{g}_0 \frac{1}{N_n} \sum_{j=1}^{N_n} ((j/N_n)^k - (j/N_n)^{2k}) < \frac{(1 + \varepsilon) \varsigma(S_0) k}{(k + 1)(2k + 1)},$$

we obtain the following condition on h_*

$$h_* \geq (v_\varepsilon^*)^{1/(2k+1)}, \quad (6.27)$$

where

$$v_\varepsilon^* = \frac{(1 + \varepsilon)k}{c_\varepsilon^*(k + 1)(2k + 1)} \quad \text{and} \quad c_\varepsilon^* = \frac{2^{2k+1}(1 - \varepsilon)r}{\pi^{2k} \varsigma(S_0)}.$$

To maximize the function $\Psi_{N_n}^*(R)$ on the right-hand side of the inequality (6.25) we take $R = R_n^*$ defined in (6.22). Therefore we obtain that

$$\liminf_{n \rightarrow \infty} \inf_{\widehat{S}_n} n^{2k/(2k+1)} \mathcal{E}_n(\widehat{S}_n) \geq \varsigma(S_0) F(h_*)/2, \quad (6.28)$$

where

$$F(x) = \frac{1}{x} - \frac{2k + 1}{(k + 1)^2 (c_\varepsilon^*(2k + 1)x^{2k+2} + x)}.$$

Furthermore, taking into account that

$$F'(x) = -\frac{(c_\varepsilon^*(2k + 1)(k + 1)x^{2k+1} - k)^2}{(k + 1)^2 (c_\varepsilon^*(2k + 1)x^{2k+2} + x)^2} \leq 0$$

we get

$$\max_{h_* \geq (v_\varepsilon^*)^{1/(2k+1)}} F(h_*) = F((v_\varepsilon^*)^{1/(2k+1)}) = \frac{(1 + \varepsilon')k}{k + 1} (v_\varepsilon^*)^{-1/(2k+1)},$$

where

$$\varepsilon' = \frac{\varepsilon}{2k + \varepsilon k + 1}. \quad (6.29)$$

This means that to obtain in (6.28) the maximal lower bound one has to take in (6.26)

$$h_* = (v_\varepsilon^*)^{1/(2k+1)}. \quad (6.30)$$

It is important to note that if one defines the prior distribution μ_ϑ in the bayesian risk (6.13) by formulas (6.10), (6.24), (6.26) and (6.30), then the bayesian risk would depend on a parameter $0 < \varepsilon < 1$, i.e. $\mathcal{E}_n = \mathcal{E}_{\varepsilon,n}$.

Therefore, the inequality (6.28) implies that, for any $0 < \varepsilon < 1$,

$$\liminf_{n \rightarrow \infty} \inf_{\widehat{S}_n} n^{2k/(2k+1)} \mathcal{E}_{\varepsilon, n}(\widehat{S}_n) \geq \frac{(1 + \varepsilon')(1 - \varepsilon)^{1/(2k+1)}}{(1 + \varepsilon)^{1/(2k+1)}} \gamma_k(S_0), \quad (6.31)$$

where the function $\gamma_k(S_0)$ is defined in (4.7) for $S_0 \equiv 0$.

Now to end the definition of the sequence of the random functions $(S_{\vartheta, n})$ defined by (6.9) and (6.10) one has to define the sequence (N_n) . Let us remember that we make use of the sequence $(S_{\vartheta, n})$ with the coefficients $(t_{m, j})$ constructed in (6.24) for $R = R_n^*$ given in (6.22) and for the sequence h_n given by (6.26) and (6.30) for some fixed arbitrary $0 < \varepsilon < 1$.

We will choose the sequence (N_n) to satisfy the conditions $\mathbf{A}_1)$ – $\mathbf{A}_4)$. One can take, for example, $N_n = \lceil \ln^4 n \rceil + 1$. Then the condition $\mathbf{A}_1)$ is trivial. Moreover, taking into account that in this case

$$R_n^* = \frac{2^{2k+1}(1 - \varepsilon)r}{\pi^{2k}\widehat{g}_0} \nu_\varepsilon^* N_n^{2k+1} = \frac{\varsigma(S_0)}{\widehat{g}_0} \frac{k}{(k+1)(2k+1)} N_n^{2k+1}$$

we find thanks to the convergence (6.23)

$$\lim_{n \rightarrow \infty} \frac{R_n^* + \sum_{j=1}^{N_n} j^{2k}}{N_n^k \sum_{j=1}^{N_n} j^k} = 1.$$

Therefore, the solution (6.20), for sufficiently large n , satisfies the following inequality

$$\max_{1 \leq j \leq N_n} y_j^*(R_n^*) j^k \leq 2N_n^k.$$

Now it is easy to see that the condition $\mathbf{A}_2)$ holds with $d_n = \sqrt{N_n}$ and the condition $\mathbf{A}_4)$ holds for arbitrary $0 < \epsilon_0 < 1$. As to the condition $\mathbf{A}_3)$, note that in view of the definition of $t_{m, j}$ in (6.24) we get

$$\begin{aligned} \frac{1}{h_n^{2k-1}} \sum_{m=1}^{M_n} \sum_{j=1}^{N_n} t_{m, j}^2 j^{2k} &= \frac{1}{2n h_n^{2k+1}} \widehat{g}_0 \sum_{j=1}^{N_n} y_j^*(R_n^*) j^{2k} \\ &= \frac{R_n^* \widehat{g}_0}{N_n^{2k+1} 2\nu_\varepsilon^*} = (1 - \varepsilon)r \left(\frac{2}{\pi} \right)^{2k}. \end{aligned}$$

Hence the condition $\mathbf{A}_3)$.

7 Estimation of non periodic function

Now we consider the estimation problem of the non periodic regression function S in the model (1.1). In this case we will estimate the function S on any interior interval $[a, b]$ of $[0, 1]$, i.e. for $0 < a < b < 1$.

It should be pointed out that at the boundary points $x = 0$ and $x = 1$, one must to make use of kernel estimators (see Brua, 2007).

Let now χ be a infinitely differentiable $[0, 1] \rightarrow \mathbb{R}_+$ function such that $\chi(x) = 1$ for $a \leq x \leq b$ and $\chi^{(k)}(0) = \chi^{(k)}(1) = 0$ for all $k \geq 0$, for example,

$$\chi(x) = \frac{1}{\eta} \int_{-\infty}^{\infty} V\left(\frac{x-z}{\eta}\right) \mathbf{1}_{[a', b']}(z) dz,$$

where V is some kernel function introduced in (6.6),

$$a' = \frac{a}{2}, \quad b' = \frac{b}{2} + \frac{1}{2} \quad \text{and} \quad \eta = \frac{1}{4} \min(a, 1-b).$$

Multiplying the equation (1.1) by the function $\chi(\cdot)$ and simulating the i.i.d. $\mathcal{N}(0, 1)$ sequence $(\zeta_j)_{1 \leq j \leq n}$ one comes to the estimation problem of the periodic regression function $\tilde{S}(x) = S(x)\chi(x)$, i.e.

$$\tilde{y}_j = \tilde{S}(x_j) + \tilde{\sigma}_j(S) \tilde{\xi}_j,$$

where $\tilde{\sigma}_j(S) = \sqrt{\sigma_j^2(S) + \epsilon^2}$,

$$\tilde{\xi}_j = \frac{\sigma_j(S)}{\tilde{\sigma}_j(S)} \xi_j + \frac{\epsilon}{\tilde{\sigma}_j(S)} \zeta_j.$$

and $\epsilon > 0$ is some sufficiently small parameter.

It is easy to see that if the sequence $(\sigma_j(S))_{1 \leq j \leq n}$ satisfies the conditions $\mathbf{H}_1) - \mathbf{H}_4)$, then the sequence $(\tilde{\sigma}_j(S))_{1 \leq j \leq n}$ satisfies these conditions as well with

$$\tilde{\sigma}_j(S) = \tilde{g}(x_j, S) = \sqrt{g^2(x_j, S)\chi^2(x_j) + \epsilon^2}.$$

8 Conclusions

In conclusion, it should be noted that this paper completes the investigation of the estimation problem of the nonparametric regression function for the heteroscedastic regression model (1.1) in the case of quadratic risk. It is proved that the adaptive procedure (2.12) satisfies the non asymptotic oracle inequality and it is asymptotically efficient for estimating a periodic regression function. Moreover, in Section 7 we have explained how to apply the procedure to the case of non periodic function. As far as we know, the procedure (2.12) is unique for estimating the regression function at the model (1.1). Let us remember once more the main steps of this investigation. The procedure (2.12) combines the both principal aspects of nonparametric estimation: non asymptotic and asymptotic. Non-asymptotic aspect is based on the selection model procedure with penalization (see e.g., Barron, Birgé and Massart, 1999, or Fourdrinier, Pergamenshchikov, 2007). Our selection model procedure differs from the commonly used one by a small coefficient in the penalty term going to zero that provides the sharp non-asymptotic oracle inequality. Moreover, the commonly used selection model procedure is based on the least-squares estimators whereas our procedure uses weighted least-squares estimators with the weights minimizing the asymptotic quadratic risk that provides the asymptotic efficiency, as the final result. From practical point of view, the procedure (2.12) gives an acceptable accuracy even for small samples as it is shown via simulations by Galtchouk, Pergamenshchikov, 2008.

9 Proofs

9.1 Proof of Theorem 5.1

To prove the theorem we will adapt to the heteroscedastic case the corresponding proof from Nussbaum, 1985.

First, from (2.4) we obtain that, for any $p \in \mathcal{P}_n$,

$$\mathbf{E}_{S,p} \|\tilde{S} - S\|_n^2 = \sum_{j=1}^n (1 - \tilde{\lambda}_j)^2 \theta_{j,n}^2 + \frac{1}{n} \sum_{j=1}^n \tilde{\lambda}_j^2 \zeta_{j,n}(S), \quad (9.1)$$

where

$$\varsigma_{j,n}(S) = \frac{1}{n} \sum_{l=1}^n \sigma_l^2(S) \phi_j^2(x_l).$$

Setting now $\tilde{\omega} = \omega(\tilde{\alpha})$ with the function ω defined in (2.8), the index $\tilde{\alpha}$ defined in (5.2), $\tilde{j}_0 = [\tilde{\omega}\varepsilon_n]$, $\tilde{j}_1 = [\tilde{\omega}/\varepsilon_n]$ and

$$\varsigma_n(S) = \frac{1}{n} \sum_{l=1}^n \sigma_l^2(S),$$

we rewrite (9.1) as follows

$$\begin{aligned} \mathbf{E}_{S,p} \|\tilde{S} - S\|_n^2 &= \sum_{j=\tilde{j}_0+1}^{\tilde{j}_1-1} (1 - \tilde{\lambda}_j)^2 \theta_{j,n}^2 \\ &\quad + \frac{\varsigma_n(S)}{n} \sum_{j=1}^n \tilde{\lambda}_j^2 + \tilde{\Delta}_{1,n} + \tilde{\Delta}_{2,n} \end{aligned} \quad (9.2)$$

with

$$\tilde{\Delta}_{1,n} = \sum_{j=\tilde{j}_1}^n \theta_{j,n}^2 \quad \text{and} \quad \tilde{\Delta}_{2,n} = \frac{1}{n} \sum_{j=1}^n \tilde{\lambda}_j^2 (\varsigma_{j,n}(S) - \varsigma_n(S)).$$

Note that we have decomposed the first term on the right-hand of (9.1) into the sum

$$\sum_{j=\tilde{j}_0+1}^{\tilde{j}_1-1} (1 - \tilde{\lambda}_j)^2 \theta_{j,n}^2 + \tilde{\Delta}_{1,n}.$$

This decomposition allows us to show that $\tilde{\Delta}_{1,n}$ is negligible and further to approximate the first term by a similar term in which the coefficients $\theta_{j,n}$ will be replaced by the Fourier coefficients θ_j of the function S .

Taking into account the definition of ω in (2.8) we can bound $\tilde{\omega}$ as

$$\tilde{\omega} \geq (A_k)^{\frac{1}{2k+1}} (n\varepsilon_n)^{\frac{1}{2k+1}}.$$

Therefore, by Lemma A.1 we obtain

$$\lim_{n \rightarrow \infty} \sup_{S \in W_r^k} n^{\frac{2k}{2k+1}} \tilde{\Delta}_{1,n} = 0.$$

Let us consider now the next term $\tilde{\Delta}_{2,n}$. We have

$$\begin{aligned} |\tilde{\Delta}_{2,n}| &= \frac{1}{n^2} \left| \sum_{i=1}^n \sigma_i^2(S) \sum_{j=1}^n \tilde{\lambda}_j^2 (\phi_j^2(x_i) - 1) \right| \\ &\leq \frac{\sigma_*}{n} \sup_{0 \leq x \leq 1} \left| \sum_{j=1}^n \tilde{\lambda}_j^2 (\phi_j^2(x) - 1) \right|. \end{aligned}$$

Now by Lemma A.2 and the definition (2.7) we obtain directly the same property for $\tilde{\Delta}_{2,n}$, i.e.

$$\lim_{n \rightarrow \infty} \sup_{S \in W_r^k} n^{\frac{2k}{2k+1}} |\tilde{\Delta}_{2,n}| = 0.$$

Setting

$$\hat{\gamma}_{k,n}(S) = n^{\frac{2k}{2k+1}} \sum_{j=\tilde{j}_0}^{\tilde{j}_1-1} (1 - \tilde{\lambda}_j)^2 \theta_j^2 + \varsigma_n(S) n^{-\frac{1}{2k+1}} \sum_{j=1}^n \tilde{\lambda}_j^2$$

and applying the well-known inequality

$$(a + b)^2 \leq (1 + \delta)a^2 + (1 + 1/\delta)b^2$$

to the first term on the right-hand side of the inequality (9.2) we obtain that, for any $\delta > 0$ and for any $p \in \mathcal{P}_n$,

$$\begin{aligned} \mathbf{E}_{S,p} \|\tilde{S} - S\|_n^2 &\leq (1 + \delta) \hat{\gamma}_{k,n}(S) n^{-2k/(2k+1)} \\ &\quad + \tilde{\Delta}_{1,n} + \tilde{\Delta}_{2,n} + (1 + 1/\delta) \tilde{\Delta}_{3,n}, \end{aligned} \quad (9.3)$$

where

$$\tilde{\Delta}_{3,n} = \sum_{j=\tilde{j}_0+1}^{\tilde{j}_1-1} (\theta_{j,n} - \theta_j)^2.$$

Taking into account that $k \geq 1$ and that

$$\tilde{j}_1 \leq \bar{\omega} \varepsilon_n^{-1} + (A_k)^{\frac{1}{2k+1}} n^{\frac{1}{2k+1}} (\varepsilon_n)^{-(2k+2)/(2k+1)},$$

we can show through Lemma A.3 that

$$\lim_{n \rightarrow \infty} \sup_{S \in W_r^k} n^{\frac{2k}{2k+1}} \tilde{\Delta}_{3,n} = 0.$$

Therefore, the inequality (9.3) yields

$$\limsup_{n \rightarrow \infty} n^{\frac{2k}{2k+1}} \sup_{S \in W_r^k} \mathcal{R}_n(\tilde{S}, S) / \gamma_k(S) \leq \limsup_{n \rightarrow \infty} \sup_{S \in W_r^k} \hat{\gamma}_{k,n}(S) / \gamma_k(S)$$

and to prove (5.4) it suffices to show that

$$\limsup_{n \rightarrow \infty} \sup_{S \in W_r^k} \hat{\gamma}_{k,n}(S) / \gamma_k(S) \leq 1. \quad (9.4)$$

First, it should be noted that the definition (5.1) and the inequalities (3.7)-(3.8) imply directly

$$\lim_{n \rightarrow \infty} \sup_{S \in W_r^k} |\tilde{t}_n / \bar{r}(S) - 1| = 0.$$

Moreover, by the definition of $(\tilde{\lambda}_j)_{1 \leq j \leq n}$ in (5.3), for sufficiently large n , for which $\tilde{t}_n \geq \bar{r}(S)$ we find

$$\sup_{j \geq 1} n^{\frac{2k}{2k+1}} \frac{(1 - \tilde{\lambda}_j)^2}{(\pi j)^{2k}} = \pi^{-2k} (A_k \tilde{t}_n)^{-2k/(2k+1)} \leq \pi^{-2k} (A_k \bar{r}(S))^{-2k/(2k+1)}.$$

Therefore, by the definition of the coefficients $(a_j)_{j \geq 1}$ in (3.5)

$$\limsup_{n \rightarrow \infty} n^{\frac{2k}{2k+1}} \sup_{S \in W_r^k} \sup_{j \geq \tilde{j}_0} \pi^{2k} (A_k \bar{r}(S))^{2k/(2k+1)} (1 - \tilde{\lambda}_j)^2 / a_j \leq 1.$$

Furthermore, in view of the definition (2.7) we calculate directly

$$\lim_{n \rightarrow \infty} \sup_{S \in W_r^k} \left| n^{-\frac{1}{2k+1}} \sum_{j=1}^n \tilde{\lambda}_j^2 - (A_k \bar{r}(S))^{\frac{1}{2k+1}} \int_0^1 (1 - z^k)^2 dz \right| = 0.$$

Now, the definition of W_r^k in (3.3) and the condition (3.6) imply the inequality (9.4). Hence Theorem 5.1. \square

9.2 Proof of Theorem 4.3

In this section we prove Theorem 4.3. Lemma 5.2 implies that to prove the lower bounds (4.10) and (4.11), it suffices to show

$$\liminf_{n \rightarrow \infty} \inf_{\hat{S}_n} n^{\frac{2k}{2k+1}} \mathcal{R}_0(\hat{S}_n) \geq 1, \quad (9.5)$$

where

$$\mathcal{R}_0(\widehat{S}_n) = \sup_{S \in W_r^k} \mathbf{E}_{S,q} \|\widehat{S}_n - S\|^2 / \gamma_k(S).$$

For any estimator \widehat{S}_n , we denote by \widehat{S}_n^0 its projection onto W_r^k , i.e. $\widehat{S}_n^0 = \Pr_{W_r^k}(\widehat{S}_n)$. Since W_r^k is a convex set, we get

$$\|\widehat{S}_n - S\|^2 \geq \|\widehat{S}_n^0 - S\|^2.$$

Now we introduce the following set

$$\Xi_n = \left\{ \max_{1 \leq m \leq M_n} \max_{1 \leq j \leq N} \zeta_{m,j}^2 \leq d_n \right\}, \quad (9.6)$$

where $(\zeta_{m,j})$ are i.i.d. $\mathcal{N}(0, 1)$ random variables from (6.10) and the sequence $(d_n)_{n \geq 1}$ is given in the condition \mathbf{A}_2 . Therefore, we can write that

$$\mathcal{R}_0(\widehat{S}_n) \geq \int_{\{z: S_{z,n} \in W_r^k\} \cap \Xi_n} \frac{\mathbf{E}_{S_{z,n},q} \|\widehat{S}_n^0 - S_{z,n}\|^2}{\gamma_k(S_{z,n})} \mu_\vartheta(dz).$$

Here the kernel function family $(S_{z,n})$ is given in (6.9) in which $N_n = [\ln^4 n] + 1$ and the parameter h is defined in (6.26) and (6.30); the measure μ_ϑ is defined in (6.13). Moreover, note that on the set Ξ the random function $S_{\vartheta,n}$ is uniformly bounded, i.e.

$$\|S_{\vartheta,n}\|_\infty = \sup_{0 \leq x \leq 1} |S_{\vartheta,n}(x)| \leq \sqrt{d_n} t_n^*, \quad (9.7)$$

where the coefficient t_n^* is defined in (6.11).

Thus, we estimate the risk $\mathcal{R}_0(\widehat{S}_n)$ from below as

$$\mathcal{R}_0(\widehat{S}_n) \geq \frac{1}{\gamma_n^*} \int_{\{z: S_{z,n} \in W_r^k\} \cap \Xi_n} \mathbf{E}_{S_{z,n},q} \|\widehat{S}_n^0 - S_{z,n}\|^2 \mu_\vartheta(dz)$$

with

$$\gamma_n^* = \sup_{\|S\|_\infty \leq \sqrt{d_n} t_n^*} \gamma_k(S). \quad (9.8)$$

By making use of the Bayes risk (6.13) with the prior distribution given by formulae (6.10), (6.24), (6.26) and (6.30) for any fixed parameter $0 < \varepsilon < 1$ we rewrite the lower bound for $\mathcal{R}_0(\widehat{S}_n)$ as

$$\mathcal{R}_0(\widehat{S}_n) \geq \mathcal{E}_{\varepsilon,n}(\widehat{S}_n^0) / \gamma_n^* - 2 \Omega_n / \gamma_n^* \quad (9.9)$$

with

$$\Omega_n = \mathbf{E}(\mathbf{1}_{\{S_{\vartheta,n} \notin W_r^k\}} + \mathbf{1}_{\Xi_n^c})(r + \|S_{\vartheta,n}\|^2).$$

In Section 6.3 we proved that the parameters in chosen prior distribution satisfy the conditions $\mathbf{A}_1)$ – $\mathbf{A}_4)$. Therefore Propositions 6.3–6.4 and the limit (A.4) imply that, for any $\nu > 0$,

$$\lim_{n \rightarrow \infty} n^\nu \Omega_n = 0.$$

Moreover, by the condition $\mathbf{H}_4)$ the sequence γ_n^* goes to $\gamma_k(S_0)$ as $n \rightarrow \infty$. Therefore, from this, (6.31) and (9.9) we get, for any $0 < \varepsilon < 1$,

$$\liminf_{n \rightarrow \infty} \inf_{\hat{S}_n} n^{\frac{2k}{2k+1}} \mathcal{R}_0(\hat{S}_n) \geq \frac{(1 + \varepsilon')(1 - \varepsilon)^{\frac{1}{2k+1}}}{(1 + \varepsilon)^{\frac{1}{2k+1}}}.$$

where ε' is defined in (6.29). Limiting here $\varepsilon \rightarrow 0$ implies inequality (9.5). Hence Theorem 4.3. \square

10 Appendix

A.1 Proof of Lemma 5.2

First notice that, for any $S \in W_r^k$, one has

$$\|\hat{S}_n - S\|_n^2 = \|T_n(\hat{S}) - S\|^2 + \Delta_{1,n}^* + \Delta_{2,n}^*,$$

where

$$\Delta_{1,n}^* = 2 \sum_{j=1}^n \int_{x_{j-1}}^{x_j} (\hat{S}_n(x_j) - S(x))(S(x) - S(x_j)) dx$$

and

$$\Delta_{2,n}^* = \sum_{j=1}^n \int_{x_{j-1}}^{x_j} (S(x) - S(x_j))^2 dx.$$

For any $0 < \delta < 1$, by making use of the elementary inequality

$$2ab \leq \delta a^2 + \delta^{-1} b^2,$$

one gets

$$\Delta_{1,n}^* \leq \delta \|T_n(\hat{S}) - S\|^2 + \delta^{-1} \Delta_{2,n}^*.$$

Moreover, for any $S \in W_r^k$ with $k \geq 1$, by the Bounyakovskii-Cauchy-Schwarz inequality we obtain that

$$\Delta_{2,n}^* \leq \frac{1}{n^2} \sum_{j=1}^n \int_{x_{j-1}}^{x_j} \dot{S}^2(t) dt = \frac{1}{n^2} \|\dot{S}\|^2 \leq \frac{r}{n^2}.$$

Hence Lemma 5.2. \square

A.2 Proof of Theorem 6.1

For any $z = (z_1, \dots, z_l)' \in \mathbb{R}^n$ we set

$$\varrho_i(v, z) = \frac{1}{f(v, z)\Phi(z)} \frac{\partial}{\partial z_i} (f(v, z)\Phi(z)).$$

Note that due to the condition (3.7), the density (6.3) is bounded, i.e.

$$f(v, z) \leq (2\pi g_*)^{-n/2}.$$

So through (6.2) we obtain that

$$\lim_{|z_i| \rightarrow \infty} \tau(z) f(v, z) \varphi_i(z_i) = 0.$$

Therefore, integrating by parts yields

$$\begin{aligned} \tilde{\mathbf{E}}(\hat{\tau}_n - \tau(\vartheta))\varrho_i &= \int_{\mathbb{R}^{n+l}} (\hat{\tau}_n(v) - \tau(z)) \frac{\partial}{\partial z_i} (f(v, z)\Phi(z)) dz dv \\ &= \int_{\mathbb{R}^l} \left(\frac{\partial}{\partial z_i} \tau(z) \right) \Phi(z) \left(\int_{\mathbb{R}^n} f(v, z) dv \right) dz = \bar{\tau}_i. \end{aligned}$$

Now the Bounyakovskii-Cauchy-Schwarz inequality gives the following lower bound

$$\tilde{\mathbf{E}}(\hat{\tau}_n - \tau(\vartheta))^2 \geq \bar{\tau}_i^2 / \tilde{\mathbf{E}}\varrho_i^2.$$

To estimate the denominator in the last ratio, note that

$$\varrho_i(v, z) = \tilde{f}_i(v, z) + \frac{\dot{\varphi}_i(z_i)}{\varphi_i(z_i)} \quad \text{with} \quad \tilde{f}_i(v, z) = (\partial/\partial z_i) \ln f(v, z).$$

From (6.1) it follows that

$$\tilde{f}_i(v, z) = \sum_{j=1}^n (\xi_j^2 - 1) \frac{1}{2\sigma_j^2(z)} \frac{\partial}{\partial z_i} \sigma_j^2(z) + \sum_{j=1}^n \xi_j \frac{S'_j(x_j)}{\sigma_j(z)}.$$

Moreover, the conditions \mathbf{H}_2) and (6.4) imply

$$(\partial/\partial z_i) \sigma_j^2(z) = (\partial/\partial z_i) g^2(x_j, S_z) = \tilde{\mathbf{L}}_i(x_j, z)$$

from which it follows

$$\tilde{\mathbf{E}} \left(\tilde{f}_i(Y, \vartheta) \right)^2 = F_i + B_i.$$

This implies inequality (6.5). Hence Theorem 6.1. \square

A.3 Proof of Proposition 6.2

First note that, for $0 \leq x \leq 1$, we can represent the l th derivative as

$$S_{z,n}^{(l)}(x) = \frac{1}{h^l} \sum_{m=1}^{M_n} \sum_{i=0}^l \binom{l}{i} \chi_\eta^{(l-i)}(v_m(x)) Q_{i,m}(z, v_m(x)), \quad (\text{A.1})$$

where

$$Q_{i,m}(z, v) = \sum_{j=1}^{N_n} z_{m,j} e_j^{(i)}(v).$$

Therefore

$$\|S_{z,n}^{(l)}\|^2 = \frac{1}{h_n^{2l-1}} \sum_{m=1}^{M_n} \int_{-1}^1 \left(\sum_{i=0}^l \binom{l}{i} \chi_\eta^{(l-i)}(v) Q_{i,m}(z, v) \right)^2 dv$$

and by the Bounyakovskii-Cauchy-Schwarz inequality we obtain that

$$\|S_{z,n}^{(l)}\|^2 \leq \frac{C^{*}(l, \eta)}{h_n^{2l-1}} \sum_{i=0}^l \bar{Q}_i(z) \quad (\text{A.2})$$

with $C^{*}(l, \eta) = \max_{-1 \leq v \leq 1} \sum_{i=0}^l \left(\binom{l}{i} \chi_\eta^{(l-i)}(v) \right)^2$ and

$$\bar{Q}_i(z) = \sum_{m=1}^{M_n} \int_{-1}^1 Q_{i,m}^2(z, v) dv.$$

Now we show that, for any $0 \leq i \leq k-1$ and $\delta > 0$,

$$\lim_{n \rightarrow \infty} n^\nu \mathbf{P}(\overline{Q}_i(\vartheta) > \delta h_n^{2k-1}) = 0. \quad (\text{A.3})$$

To this end note that

$$\begin{aligned} \int_{-1}^1 Q_{i,m}^2(\vartheta, v) dv &= \sum_{j=1}^{N_n} \vartheta_{m,j}^2 \int_{-1}^1 (e_j^{(i)}(v))^2 dv \\ &\leq \left(\frac{\pi}{2}\right)^{2i} \sum_{j=1}^{N_n} t_{m,j}^2 j^{2i} \zeta_{m,j}^2. \end{aligned}$$

Therefore, taking into account the definition of the set Ξ_n in (9.6), the functions $\overline{Q}_i(\vartheta)$ with $0 \leq i \leq k-1$ can be estimated on this set as

$$\overline{Q}_i(\vartheta) \leq \left(\frac{\pi}{2}\right)^{2(k-1)} d_n \sum_{m=1}^{M_n} \sum_{j=1}^{N_n} t_{m,j}^2 j^{2(k-1)}$$

and by (6.12) we get, for any $\delta > 0$ and sufficiently large n ,

$$\mathbf{P}(\overline{Q}_i(\vartheta) > \delta h_n^{2k-1}) \leq \mathbf{P}(\Xi_n^c).$$

Moreover, for sufficiently large n ,

$$\mathbf{P}(\Xi_n^c) \leq M_n N_n e^{-d_n/2}.$$

Therefore, the condition \mathbf{A}_1) implies

$$\limsup_{n \rightarrow \infty} n^\nu \mathbf{P}(\Xi_n^c) = 0, \quad (\text{A.4})$$

for any $\nu > 0$. Hence Proposition 6.2. \square

A.4 Proof of Proposition 6.3

First of all we prove that for ϵ from the condition \mathbf{A}_3)

$$\lim_{n \rightarrow \infty} n^\nu \mathbf{P}\left(\|S_{\vartheta,n}^{(k)}\| > \sqrt{(1-\epsilon/4)r}\right) = 0. \quad (\text{A.5})$$

Indeed, putting in (A.1) $l = k$ we can represent the k th derivative of $S_{z,n}$ as follows

$$S_{z,n}^{(k)}(x) = S'_{z,n}(x) + S''_{z,n}(x) \quad (\text{A.6})$$

with

$$S'_{z,n}(x) = \frac{1}{h^k} \sum_{m=1}^{M_n} \sum_{i=0}^{k-1} \binom{k}{i} \chi_\eta^{(k-i)}(v_m(x)) Q_{i,m}(z, v_m(x))$$

and

$$S''_{z,n}(x) = \frac{1}{h^k} \sum_{m=1}^{M_n} \chi_\eta(v_m(x)) Q_{k,m}(z, v_m(x)).$$

First, note that, we can estimate the norm of $S'_{z,n}(x)$ by the same way as in the inequality (A.2), i.e.

$$\|S'_{z,n}\|^2 \leq \frac{C^*(k, \eta)}{h_n^{2k-1}} \sum_{i=0}^{k-1} \bar{Q}_i(z).$$

By making use of (A.3) we obtain that, for any $p > 0$ and for any $\delta > 0$,

$$\lim_{n \rightarrow \infty} n^\nu \mathbf{P} \left(\|S'_{z,n}\| > \delta \right) = 0. \quad (\text{A.7})$$

Let us consider now the last term in (A.6). Taking into account that $0 \leq \chi_\eta(v) \leq 1$ we get

$$\begin{aligned} \|S''_{z,n}\|^2 &= \frac{1}{h_n^{2k-1}} \sum_{m=1}^{M_n} \int_{-1}^1 \chi_\eta^2(v) Q_{k,m}^2(z, v) dv \\ &\leq \left(\frac{\pi}{2}\right)^{2k} \frac{1}{h_n^{2k-1}} \sum_{m=1}^{M_n} \sum_{j=1}^{N_n} t_{m,j}^2 j^{2k} \zeta_{m,j}^2. \end{aligned}$$

Therefore from the condition **A₃**) we get for sufficiently large n

$$\|S''_{z,n}\|^2 \leq (1 - \epsilon/2)r + \left(\frac{\pi}{2}\right)^{2k} \sum_{m=1}^{M_n} \bar{\zeta}_m := (1 - \epsilon/2)r + \left(\frac{\pi}{2}\right)^{2k} Z_n$$

with

$$\bar{\zeta}_m = \frac{1}{h_n^{2k-1}} \sum_{j=1}^{N_n} t_{m,j}^2 j^{2k} \tilde{\zeta}_{m,j} \quad \text{and} \quad \tilde{\zeta}_{m,j} = \zeta_{m,j}^2 - 1.$$

We show that for any $\nu > 0$ and for any $\delta > 0$

$$\lim_{n \rightarrow \infty} n^\nu \mathbf{P}(|Z_n| > \delta) = 0. \quad (\text{A.8})$$

Indeed, by the Chebychev inequality for any $\iota > 0$

$$\mathbf{P}(|Z_n| > \delta) \leq \mathbf{E}(Z_n)^{2\iota} / \delta^{2\iota}. \quad (\text{A.9})$$

Note now that according to the Burkholder-Davis-Gundy inequality for any $\iota > 1$ there exists a constant $B^*(\iota) > 0$ such that

$$\mathbf{E}(Z_n)^{2\iota} \leq B^*(\iota) \mathbf{E} \left(\sum_{m=1}^{M_n} \bar{\zeta}_m^2 \right)^\iota.$$

Moreover, by putting

$$\tilde{\zeta}_* = \max_{1 \leq m \leq M_n} \max_{1 \leq j \leq N_n} \tilde{\zeta}_{m,j}^2$$

we can estimate the random variable $\bar{\zeta}_m$ as

$$\bar{\zeta}_m^2 \leq \frac{N_n}{h_n^{4k-2}} \sum_{j=1}^{N_n} t_{m,j}^4 j^{4k} \tilde{\zeta}_*.$$

Therefore, by the condition \mathbf{A}_4), for sufficiently large n ,

$$\begin{aligned} \mathbf{E}(Z_n)^{2\iota} &\leq B^*(\iota) N_n^\iota h_n^{\iota\epsilon_0} \mathbf{E} \tilde{\zeta}_*^\iota \\ &\leq B^*(\iota) \mathbf{E}(\zeta^2 - 1)^{2\iota} M_n N_n^{\iota+1} h_n^{\iota\epsilon_0}, \end{aligned}$$

where $\zeta \sim \mathcal{N}(0, 1)$. Now the condition \mathbf{A}_1) implies, for sufficiently large n ,

$$\mathbf{E}(Z_n)^{2\iota} \leq n^{-\delta_1(\iota\epsilon_0 - 2)}.$$

Thus, choosing in (A.9)

$$\iota > \nu / (\epsilon_0 \delta_1) + 2 / \epsilon_0$$

we obtain the limiting equality (A.8) which together with (A.6)-(A.7) implies (A.5). Now it is easy to deduce that Proposition 6.2 yields Proposition 6.3.

□

A.5 Proof of Proposition 6.4

First of all, we recall that, due to the condition \mathbf{A}_2 ,

$$\lim_{n \rightarrow \infty} \sum_{m=1}^{M_n} \sum_{j=1}^{N_n} t_{m,j}^2 \leq \lim_{n \rightarrow \infty} \frac{d_n}{h_n^{2k-1}} \sum_{m=1}^{M_n} \sum_{j=1}^{N_n} t_{m,j}^2 j^{2(k-1)} = 0.$$

Therefore, taking into account that

$$\|S_{\vartheta,n}\|^2 \leq h_n \sum_{m=1}^{M_n} \sum_{j=1}^{N_n} t_{m,j}^2 \zeta_{m,j}^2, \quad (\text{A.10})$$

we obtain, for sufficiently large n ,

$$\mathbf{E} \|S_{\vartheta,n}\|^2 \left(\mathbf{1}_{\{S_{\vartheta,n} \notin W_r^k\}} + \mathbf{1}_{\Xi_n^c} \right) \leq \max_{m,j} \mathbf{E} \zeta_{m,j}^2 \left(\mathbf{1}_{\{S_{\vartheta,n} \notin W_r^k\}} + \mathbf{1}_{\Xi_n^c} \right).$$

Moreover, for any $1 \leq m \leq M_n$ and $1 \leq j \leq N_n$, we estimate the last term as

$$\begin{aligned} \mathbf{E} \zeta_{m,j}^2 \left(\mathbf{1}_{\{S_{\vartheta,n} \notin W_r^k\}} + \mathbf{1}_{\Xi_n^c} \right) &\leq n \mathbf{P}(S_{\vartheta,n} \notin W_r^k) \\ &\quad + n \mathbf{P}(\Xi_n^c) + 2\mathbf{E} \zeta^2 \mathbf{1}_{\{\zeta^2 \geq n\}}, \end{aligned}$$

where $\zeta \sim \mathcal{N}(0, 1)$. By applying now Proposition 6.3 and the limit (A.4) we come to Proposition 6.4. \square

A.6 Proof of Proposition 6.5

Taking into account the inequality (9.7) and the condition \mathbf{H}_1) we obtain

$$\begin{aligned} \mathbf{E} |g^{-2}(x, S_{\vartheta,n}) - g_0^{-2}(x)| &\leq \max_{\|S\|_\infty \leq \sqrt{d_n} t_n^*} |g^{-2}(x, S) - g_0^{-2}(x)| \\ &\quad + (2/g_*) \mathbf{P}(\Xi_n^c). \end{aligned}$$

Conditions \mathbf{A}_2) and \mathbf{H}_4) together with the limit relation (A.4) imply Proposition 6.5. \square

A.7 Properties of the trigonometric basis

Lemma A.1. For any function $S \in W_r^k$,

$$\sup_{n \geq 1} \sup_{1 \leq m \leq n-1} m^{2k} \left(\sum_{j=m+1}^n \theta_{j,n}^2 \right) \leq \frac{4r}{\pi^{2(k-1)}}. \quad (\text{A.11})$$

Lemma A.2. For any $m \geq 0$,

$$\sup_{N \geq 2} \sup_{x \in [0,1]} N^{-m} \left| \sum_{l=2}^N l^m (\phi_l^2(x) - 1) \right| \leq 2^m. \quad (\text{A.12})$$

Proofs of Lemma A.1 and Lemma A.2 are given in Galtchouk, Pergamenshchikov, 2007.

Lemma A.3. Let $\theta_{j,n}$ and θ_j be the Fourier coefficients defined in (2.3) and (3.4), respectively. Then, for $1 \leq j \leq n$ and $n \geq 2$,

$$\sup_{S \in W_r^1} |\theta_{j,n} - \theta_j| \leq 2\pi \sqrt{r} j/n. \quad (\text{A.13})$$

Proof. Indeed, we have

$$\begin{aligned} |\theta_{j,n} - \theta_j| &= \left| \sum_{l=1}^n \int_{x_{l-1}}^{x_l} (S(x_l)\phi_j(x_l) - S(x)\phi_j(x)) dx \right| \\ &\leq n^{-1} \sum_{l=1}^n \int_{x_{l-1}}^{x_l} (|\dot{S}(z)\phi_j(z)| + |S(z)\dot{\phi}_j(z)|) dz \\ &= n^{-1} \int_0^1 (|\dot{S}(z)| |\phi_j(z)| + |S(z)| |\dot{\phi}_j(z)|) dz. \end{aligned}$$

By making use of the Bounyakovskii-Cauchy-Schwarz inequality we get

$$\begin{aligned} |\theta_{j,n} - \theta_j| &\leq n^{-1} \left(\|\dot{S}\| \|\phi\| + \|\dot{\phi}\| \|S\| \right) \\ &\leq n^{-1} \left(\|\dot{S}\| + \pi j \|S\| \right). \end{aligned}$$

The definition of class W_r^1 implies (A.13). Hence Lemma A.1. \square

A.8 Proofs of (6.16) and (6.17)

First of all, note that Proposition 6.5, the condition (3.7) and the condition \mathbf{H}_4) imply that

$$\lim_{n \rightarrow \infty} \max_{1 \leq m \leq M_n} \sup_{0 \leq x \leq 1} \mathbf{1}_{\{|v_m(x)| \leq 1\}} \mathbf{E} |g^{-2}(x, S_{\vartheta, n}) - g_0^{-2}(\tilde{x}_m)| = 0. \quad (\text{A.14})$$

Let us show now that for any continuously differentiable function f on $[-1, 1]$

$$\lim_{n \rightarrow \infty} \sup_{1 \leq m \leq M_n} \left| \frac{1}{nh} \sum_{i=1}^n f(v_m(x_i)) \mathbf{1}_{\{|v_m(x_i)| \leq 1\}} - \int_{-1}^1 f(v) dv \right| = 0. \quad (\text{A.15})$$

Indeed, setting

$$\Delta_{n,m} = \frac{1}{nh} \sum_{i=1}^n f(v_m(x_i)) \mathbf{1}_{\{|v_m(x_i)| \leq 1\}} - \int_{-1}^1 f(v) dv,$$

we deduce

$$\begin{aligned} |\Delta_{n,m}| &= \left| \frac{1}{nh} \sum_{i=i_*}^{i^*} f(v_m(x_i)) - \int_{-1}^1 f(v) dv \right| \\ &\leq \sum_{i=i_*}^{i^*} \int_{v_m(x_{i-1})}^{v_m(x_i)} |f(v_m(x_i)) - f(z)| dz + \max_{|z| \leq 1} |f(z)| (2 - v^* + v_*), \end{aligned}$$

where $i_* = [n\tilde{x}_m - nh] + 1$, $i^* = [n\tilde{x}_m + nh]$,

$$v_* = ([n\tilde{x}_m - nh] + 1 - n\tilde{x}_m)/(nh) \quad \text{and} \quad v^* = ([n\tilde{x}_m + nh] - n\tilde{x}_m)/(nh).$$

Therefore, taking into account that the derivative of the function f is bounded on the interval $[-1, 1]$ we obtain that

$$|\Delta_{n,m}| \leq \frac{3 \max_{|z| \leq 1} |f'(z)| + 2 \max_{|z| \leq 1} |f(z)|}{nh_n}.$$

Taking into account the conditions on the sequence $(h_n)_{n \geq 1}$ given in \mathbf{A}_1) we obtain limiting equality (A.15) which together with (A.14) implies (6.16).

Now we study the behavior of $B_{m,j}$. Due to the inequality (3.9) we estimate the Fréchet derivative as

$$|\tilde{\mathbf{L}}_{m,j}(x, S_{\vartheta,n})| \leq C^* (|S_{\vartheta,n}(x)D_{m,j}(x)| + |D_{m,j}|_1 + \|S_{\vartheta,n}\| \|D_{m,j}\|) . \quad (\text{A.16})$$

Consider now the first term on the right-hand side of this inequality. We have

$$\begin{aligned} \mathbf{E}(S_{\vartheta,n}(x)D_{m,j}(x))^2 &= \mathbf{E} \left(\sum_{l=1}^{N_n} \vartheta_{m,l} e_l(v_m(x)) \right)^2 e_j^2(v_m(x)) I_\eta^4(v_m(x)) \\ &\leq \sum_{l=1}^{N_n} t_{m,l}^2 \mathbf{1}_{\{|v_m(x)| \leq 1\}} \leq (t_n^*)^2 \mathbf{1}_{\{|v_m(x)| \leq 1\}} . \end{aligned}$$

We recall that the sequence t_n^* is defined in (6.11). Therefore, property (A.15) implies

$$\max_{1 \leq m \leq M_n} \max_{1 \leq j \leq N_n} \frac{1}{nh} \sum_{i=1}^n \mathbf{E}(S_{\vartheta,n}(x_i)D_{m,j}(x_i))^2 = O((t_n^*)^2) .$$

As to the second term on the right-hand side of (A.16), we get

$$|D_{m,j}|_1 = \int_0^1 |e_j(v_m(x)) \chi_\eta(v_m(x))| dx = h \int_{-1}^1 |e_j(v) \chi_\eta(v)| dv \leq 2h .$$

Similarly, $\|D_{m,j}\|^2 \leq h$ and, by (A.10)

$$\mathbf{E}\|S_{\vartheta,n}\|^2 \leq h \sum_{m=1}^{M_n} \sum_{j=1}^{N_n} t_{m,j}^2 \leq (t_n^*)^2 .$$

Therefore,

$$\frac{1}{nh} \max_{1 \leq m \leq M_n} \max_{1 \leq j \leq N_n} |B_{m,j}| = O((t_n^*)^2 + h_n)$$

and the condition \mathbf{A}_1) implies (6.17). \square

A.9 Proof of (6.18)

Indeed, by the direct calculation it easy to see that, for any $N \geq 1$ and for any vector $(y_1, \dots, y_N)' \in \mathbb{R}_+^N$,

$$\left| \frac{\sum_{j=1}^N \psi_j(\eta, y_j)}{\Psi_N(y_1, \dots, y_N)} - 1 \right| \leq \frac{\max_{j \geq 1} \left(|\bar{e}_j^2(\chi_\eta) - \bar{e}_j(\chi_\eta^2)| + |\bar{e}_j^2(\chi_\eta) - 1| \right)}{\min_{j \geq 1} \bar{e}_j(\chi_\eta^2)},$$

where the operator $\bar{e}_j(f)$ is defined in (6.14). Moreover, we remember that $\int_{-1}^1 e_j^2(v)dv = 1$. Therefore, taking into account the property (6.7) we obtain (6.18). \square

Acknowledgements

We are grateful for the comments and stimulating questions by two anonymous Referees which have led to considerable improvements.

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