



# PURDUE UNIVERSITY



**CENTER FOR STATISTICAL  
DECISION SCIENCES AND  
DEPARTMENT OF STATISTICS**

DATE RECEIVED 10/16/2000

20000707 147

EMPIRICAL BAYES TESTS WITH  $n^{-1+\epsilon}$  CONVERGENCE RATE  
IN CONTINUOUS ONE-PARAMETER EXPONENTIAL FAMILY

by

Shanti S. Gupta and Jianjun Li  
Purdue University

Technical Report #99-09C

Department of Statistics  
Purdue University  
West Lafayette, IN USA

June 1999

◇

# Empirical Bayes Tests With $n^{-1+\varepsilon}$ Convergence Rate In Continuous One-Parameter Exponential Family

Shanti S. Gupta  
Department of Statistics  
Purdue University  
W. Lafayette, IN 47907

and

Jianjun Li  
Department of Statistics  
Purdue University  
W. Lafayette, IN 47907

## Abstract

Empirical Bayes tests for testing  $H_0 : \theta \leq \theta_0$  against  $H_1 : \theta > \theta_0$  in the continuous one-parameter family with density  $c(\theta)\exp\{\theta x\}h(x)$ ,  $\infty \leq \alpha < x < \beta \leq \infty$ , are considered under the linear loss. Using the assumptions that  $\int_{\Omega} |\theta| dG(\theta) < \infty$  and the critical point  $b_0$  of a Bayes test falls in some known interval  $[C_1, C_2]$ , where  $\alpha < C_1 < C_2 < \beta$ , we show that, for any  $0 < \varepsilon < 1$ , the empirical Bayes tests can be constructed such that they have a convergence rate of order  $o(n^{-1+\varepsilon})$ , which generalizes the result of Liang (1999) from the positive (one-parameter) exponential family to any continuous one-parameter exponential family.

AMS 1991 Subject classification: 62c12.

Keywords: Empirical Bayes; exponential family; rate of convergence; regret.

---

\*This research was supported in part by US Army Research Office, Grant DAAH04-95-0165 at Purdue University.

## § 1 Introduction

Let  $X$  denote a random variable from the exponential family having density function

$$f(x|\theta) = c(\theta)\exp\{\theta x\}h(x), \quad -\infty \leq \alpha < x < \beta \leq +\infty, \quad (1.1)$$

where  $h(x)$  is continuous, positive for  $x \in (\alpha, \beta)$ ,  $\theta$  is the natural parameter, distributed according to an unknown prior distribution  $G$  on the parameter space  $\Omega = \{\theta : c(\theta) > 0\}$ .

We consider the problem of testing the hypotheses  $H_0 : \theta \leq \theta_0$  versus  $H_1 : \theta > \theta_0$ , where  $\theta_0 \in \Omega$ .

Let  $a = i$  be the action in favor of  $H_i$ . For the parameter  $\theta$  and action  $a$ , we use the loss function

$$l(\theta, a) = a(\theta_0 - \theta)I_{[\theta \leq \theta_0]} + (1 - a)(\theta - \theta_0)I_{[\theta > \theta_0]}, \quad (1.2)$$

where  $I_{[\ ]}$  is the indicator function, which equals 1 or 0 if the statement inside [ ] is true or not. Assume that  $\int_{\Omega} |\theta| dG(\theta) < \infty$ . Define

$$\alpha_G(x) = \int_{\Omega} c(\theta)e^{\theta x} dG(\theta),$$

and

$$\psi_G(x) = \int_{\Omega} \theta c(\theta)e^{\theta x} dG(\theta).$$

By Fubini Theorem,

$$\begin{aligned} & \int_{\alpha}^{\beta} h(x)\alpha_G(x)dx \quad (1.3) \\ &= \int_{\alpha}^{\beta} h(x) \int_{\Omega} c(\theta)e^{\theta x} dG(\theta)dx \\ &= \int_{\Omega} \int_{\alpha}^{\beta} c(\theta)e^{\theta x} h(x) dx dG(\theta) = 1 \end{aligned}$$

and

$$\begin{aligned} & \int_{\alpha}^{\beta} h(x)|\psi_G(x)|dx \quad (1.4) \\ &\leq \int_{\alpha}^{\beta} h(x) \int_{\Omega} |\theta|c(\theta)e^{\theta x} dG(\theta)dx \\ &= \int_{\Omega} |\theta| \int_{\alpha}^{\beta} c(\theta)e^{\theta x} dx dG(\theta) \\ &= \int_{\Omega} |\theta| dG(\theta) < \infty. \end{aligned}$$

Let  $W(x) = \theta_0\alpha_G(x) - \psi_G(x)$ . Then  $W(x)$  is a continuous function. By (1.3) and (1.4),

$$\begin{aligned} & \int_{\alpha}^{\beta} |W(x)|h(x)dx \\ &< \int_{\alpha}^{\beta} [|\theta_0|\alpha_G(x) + |\psi_G(x)|]h(x)dx < \infty. \quad (1.5) \end{aligned}$$

A test  $\delta(x)$  is defined to be a measurable mapping from  $(\alpha, \beta)$  into  $[0, 1]$  so that  $\delta(x) = P\{\text{accepting } H_1 | X = x\}$ , i.e.,  $\delta(x)$  is the probability of accepting  $H_1$  when  $X = x$  is observed.

Let  $R(G, \delta)$  denote the Bayes risk of the test  $\delta$  when  $G$  is the prior distribution. Then Bayes risk  $R(G, \delta)$  can be expressed as

$$\begin{aligned}
R(G, \delta) &= C_G + \int_{\Omega} \int_{\alpha}^{\beta} \delta(x)(\theta_0 - \theta)c(\theta)e^{\theta x}h(x)dx dG(\theta) \quad (1.6) \\
&= C_G + \int_{\alpha}^{\beta} \delta(x) \left[ \int_{\Omega} (\theta_0 - \theta)c(\theta)e^{\theta x} dG(\theta) \right] h(x) dx \\
&= C_G + \int_{\alpha}^{\beta} \delta(x) [\theta_0 \alpha_G(x) - \psi_G(x)] h(x) dx \\
&= C_G + \int_{\alpha}^{\beta} \delta(x) W(x) h(x) dx \\
&= C_G + \int_{\alpha}^{\beta} \delta(x) [\theta_0 - \phi_G(x)] \alpha_G(x) h(x) dx,
\end{aligned}$$

where

$$C_G = \int_{\Omega} (\theta - \theta_0) I_{[\theta > \theta_0]} dG(\theta),$$

and

$$\phi_G(x) = E[\theta | X = x] = \frac{\psi_G(x)}{\alpha_G(x)}.$$

Here,  $\phi_G(x)$  is the posterior mean of  $\theta$  given  $X = x$ .  $\phi_G(x)$  is continuous and increasing in  $x$ .

From (1.6), we see that a Bayes test  $\delta_G$  is determined by

$$\delta_G(x) = \begin{cases} 1 & \text{if } W(x) \leq 0 \\ 0 & \text{if } W(x) > 0 \end{cases} \quad (1.7)$$

$$= \begin{cases} 1 & \text{if } \phi_G(x) \geq \theta_0 \\ 0 & \text{if } \phi_G(x) < \theta_0. \end{cases} \quad (1.8)$$

The minimum Bayes risk is

$$R(G, \delta_G) = C_G + \int_{\alpha}^{\beta} \delta_G(x) W(x) h(x) dx. \quad (1.9)$$

To exclude trivial cases, we assume that

$$\lim_{x \rightarrow \alpha} \phi_G(x) < \theta_0 < \lim_{x \rightarrow \beta} \phi_G(x). \quad (1.10)$$

From (1.10), we get that  $\phi_G(x)$  is strictly increasing and there exists the unique point  $b_0$  (critical value) such that  $\phi_G(b_0) = \theta_0$ .  $\phi_G(x) < \theta_0$  for  $x < b_0$ , and  $\phi_G(x) > \theta_0$  for  $x > b_0$ . Therefore, the Bayes test  $\delta_G$  can be represented as

$$\delta_G(x) = \begin{cases} 1 & \text{if } x \geq b_0 \\ 0 & \text{if } x < b_0. \end{cases}$$

Furthermore we assume that there exist two known constants  $C_1, C_2, \alpha < C_1 < C_2 < \beta$  such that

$$C_1 \leq b_0 \leq C_2. \quad (1.11)$$

We will deal with this testing problem via the empirical Bayes approach. The empirical Bayes approach was introduced first by Robbins (1956, 1964). Let  $X_1, X_2, \dots, X_n$  denote the observations from  $n$  independent past experiences. Denote  $\widetilde{X}_n = (X_1, X_2, \dots, X_n)$ . Let  $X$  be the present observation. An empirical Bayes test  $\delta_n(X, \widetilde{X}_n)$  is defined to be the probability of accepting  $H_1$  when  $X$  and  $\widetilde{X}_n$  are observed. Let  $R(G, \delta_n | \widetilde{X}_n)$  denote the Bayes risk of  $\delta_n$  conditioning on  $\widetilde{X}_n$  and  $R(G, \delta) = E[R(G, \delta | \widetilde{X}_n)]$  the overall (unconditional) Bayes risk of  $\delta_n$ .

Since  $R(G, \delta_G)$  is the minimum Bayes risk,  $R(G, \delta_n | \widetilde{X}_n) - R(G, \delta_G) \geq 0$  for all  $\widetilde{X}_n$  and for all  $n$ . Thus, the regret  $R(G, \delta_n) - R(G, \delta_G) \geq 0$  for all  $n$ . The nonnegative regret  $R(G, \delta_n) - R(G, \delta_G)$  is often used as a measure of performance of the empirical Bayes test of  $\delta_n$ .

Johns and Van Ryzin (1972) constructed empirical Bayes tests for the continuous one-parameter exponential family and studied the rate of convergence for their associated regrets. Van Houwelingen (1976) improved Johns and Van Ryzin's result by using the monotonicity of the testing problem and showed that the empirical Bayes tests there have a convergence rate of order  $O(n^{-2r/(2r+3)} \log^2(n))$ , where  $r \geq 1$  is an integer, associated with the moment condition that  $\int_{\Omega} |\theta|^{r+1} dG(\theta) < \infty$  and the  $r$ -times differentiability condition on  $m(x)$ . Under the assumptions  $\int_{\Omega} |\theta|^{r+1} dG(\theta) < \infty$ , (1.11), and a few others, Karunamuni and Yang (1995) claimed that the empirical Bayes tests, constructed by them, achieve an exact rate of convergence of order  $O(n^{-2r/(2r+3)})$ . Note that if  $r$  is small, these rates are still slow. Recently, Liang (1999) investigated the empirical Bayes test for the positive exponential family and much improved the previous results. His paper shows that the empirical Bayes tests there have a rate of convergence of order  $O(n^{-s/(s+3)})$ , where  $s > 0$  is any prespecified number, under the (weaker) condition  $\int_0^{\infty} \theta dG(\theta) < \infty$  and (1.11).

Our research interest on empirical Bayes tests is motivated by Liang (1999). Making full use of properties of  $W(x)$ , with the help of classical result about sum of i.i.d. random variables, under the assumption that  $\int_{\Omega} |\theta| dG(\theta) < \infty$  and (1.11), we show that, for any  $0 < \varepsilon < 1$ , the empirical Bayes tests can be constructed such that they have a convergence rate of order  $o(n^{-1+\varepsilon})$ . Thus our result generalizes the result of Liang(1999) from the positive (one-parameter) exponential family to any continuous one-parameter exponential family.

The paper is organized as follows: §1 gives the introduction; §2 constructs a empirical Bayes test  $\delta_n$ ; §3 proves that the empirical Bayes test has a convergence rate of order  $o(n^{-1+\varepsilon})$ . §4 proves the lemmas stated in §3.

## § 2 Construction of Empirical Bayes Tests

We use the kernel method to construct the empirical Bayes tests. The idea here is similar to that of Stijnen (1985) and Liang (1999).

For any  $0 < \varepsilon < 1$ , take integer  $m$  such that  $m\varepsilon > 4$ . Suppose  $[a, b]$  is a finite closed interval inside  $(\alpha, \beta)$ . For each  $i = 0, 1$ , let  $K_i(y)$  be a Borel-measurable, bounded function vanishing outside the interval  $[a, b]$  and for  $K_0(y)$

$$\int_a^b y^j K_0(y) dy = \begin{cases} 1 & \text{if } j = 0, \\ 0 & \text{if } j = 1, 2, \dots, m-1, \end{cases} \quad (2.1)$$

and for  $K_1(y)$

$$\int_a^b y^j K_1(y) dy = \begin{cases} 0 & \text{if } j = 0, 2, 3, \dots, m, \\ 1 & \text{if } j = 1. \end{cases} \quad (2.2)$$

We may let  $B_2$  be a positive constant such that  $|K_i(y)| \leq B_2$  for all  $y \in [a, b]$  and  $i = 0$  or  $1$ .

Let  $u = u(n) = n^{-\frac{\varepsilon}{4}}$ . Then  $u \rightarrow 0$  as  $n \rightarrow \infty$ . For any  $x \in (\alpha, \beta)$ , define

$$\alpha_n(x) = \frac{1}{nu} \sum_{j=1}^n K_0\left(\frac{X_j - x}{u}\right) / h(X_j), \quad (2.3)$$

and

$$\psi_n(x) = \frac{1}{nu^2} \sum_{j=1}^n K_1\left(\frac{X_j - x}{u}\right) / h(X_j). \quad (2.4)$$

Let  $W_n(x) = \theta_0 \alpha_n(x) - \psi_n(x)$ . We shall show later that  $W_n(x)$  is an asymptotically unbiased and consistent estimator of  $W(x)$  (Lemma 3.2). Recalling that the critical value  $b_0$  is inside  $[C_1, C_2]$ , then an empirical Bayes test  $\delta_n(x, \tilde{X}_n)$  can be proposed by

$$\delta_n = \begin{cases} 1 & \text{if } (x > C_2) \text{ or } (C_1 \leq x \leq C_2 \text{ and } W_n(x) \leq 0), \\ 0 & \text{if } (x < C_1) \text{ or } (C_1 \leq x \leq C_2 \text{ and } W_n(x) > 0). \end{cases} \quad (2.5)$$

The conditional Bayes risk of the empirical Bayes test  $\delta_n$  is:

$$R(G, \delta_n | \tilde{X}_n) = C_G + \int_{\alpha}^{\beta} \delta_n(x) W(x) h(x) dx. \quad (2.6)$$

Note that  $W(x) \leq 0$  if  $x \in [C_1, b_0]$ ;  $W(x) \geq 0$  if  $x \in [b_0, C_2]$ . Then the conditional regret can be expressed as

$$\begin{aligned} R(G, \delta_n | \tilde{X}_n) - R(G, \delta) &= \int_{\alpha}^{\beta} (\delta_n - \delta) W(x) h(x) dx & (2.7) \\ &= \int_{C_1}^{b_0} I_{[W_n(x) \leq 0]} W(x) h(x) dx \\ &\quad + \int_{b_0}^{C_2} I_{[W_n(x) > 0]} |W(x)| h(x) dx \end{aligned}$$

and the unconditional regret becomes

$$\begin{aligned} R(G, \delta_n) - R(G, \delta) &= \int_{C_1}^{b_0} P(W_n(x) \leq 0)W(x)h(x)dx \quad (2.8) \\ &+ \int_{b_0}^{C_2} P(W_n(x) > 0)|W(x)|h(x)dx. \end{aligned}$$

### §3 Asymptotic Optimality of $\delta_n(x)$

In this section, we shall prove that  $R(G, \delta_n) - R(G, \delta_G) = o(n^{-1+\epsilon})$ . The convergence rate of  $R(G, \delta_n) - R(G, \delta_G)$  depends on the properties of  $W(x)$  and  $W_n(x)$ . The more information about  $W(x)$  and  $W_n(x)$  (including  $h(x)$ ) is used, the more accurate rate we will get. So firstly, we dig out a few properties of  $W(x)$  and  $W_n(x)$ . That is a few lemmas, whose proofs are left to §4. Then we state two well-known facts. Following that, a desired convergence rate of  $R(G, \delta_n) - R(G, \delta_G)$  is given as a theorem.

The first lemma is concerned about  $W(x)$ , which gives us a solution to deal with the case that  $W(x)$  is small, but not small enough.

**Lemma 3.1** *For any  $\eta > 0$ , define  $l(\eta) = \int_{C_1}^{C_2} I_{\{|W(x)| \leq \eta\}} dx$ , the Lebesgue measure of  $\{x : |W(x)| \leq \eta\} \cap [C_1, C_2]$ . Then there exists an  $\eta_0 > 0$  and some positive constant  $B_3 > 0$  such that, for any  $\eta \leq \eta_0$ ,*

$$l(\eta) \leq B_3\eta. \quad (3.1)$$

Next we consider  $W_n(x)$ . We have two lemmas, which are direct results of computations. Note that

$$W_n(x) = \theta_0\alpha_n(x) - \psi_n(x) = \frac{1}{n} \sum_{j=1}^n V(X_j, x, n), \quad (3.2)$$

where

$$V(X_j, x, n) = \frac{\theta_0}{u} \times \frac{K_0\left(\frac{X_j-x}{u}\right)}{h(X_j)} - \frac{1}{u^2} \times \frac{K_1\left(\frac{X_j-x}{u}\right)}{h(X_j)}. \quad (3.3)$$

Let  $\bar{W}(x, n) = E[V(X_j, x, n)]$  and  $Z_{jn} = V(X_j, x, n) - \bar{W}(x, n)$ . Then we have

**Lemma 3.2**  $\bar{W}(x, n)$  can be expressed as

$$\bar{W}(x, n) = W(x) + u^m W(x, n),$$

where  $W(x, n)$  is some function such that  $|W(x, n)| \leq B_4$  for all  $x \in [C_1, C_2]$  and  $n > N_1$ , and where  $B_4$  is some positive number and  $N_1$  is some integer.

Also, we have

**Lemma 3.3** *For any fixed  $n$ ,  $Z_{jn}$  are i.i.d. and*

$$EZ_{jn} = 0, \quad EZ_{jn}^2 = \frac{1}{u^3}D_2(x, n), \quad E|Z_{jn}|^3 = \frac{1}{u^5}D_3(x, n),$$

where  $D_2(x, n)$  and  $D_3(x, n)$  are some functions such that  $D_2(x, n) \leq B_5$ ,  $D_3(x, n) \leq B_5$ ,  $\frac{D_3(x, n)}{D_2(x, n)} \leq B_5$  for all  $x \in [C_1, C_2]$  and  $n > N_1$ , and where  $B_5$  is some positive number and  $N_1$  is same as in Lemma 3.2.

From Lemma 3.3, we see that

$$P(W_n(x) > 0) = P\left(\frac{1}{\sqrt{nEZ_{jn}^2}} \sum_{j=1}^n Z_{jn} > -\sqrt{nu^3D_2^{-1}(x, n)}\bar{W}(x, n)\right), \quad (3.4)$$

and

$$P(W_n(x) \leq 0) = P\left(\frac{1}{\sqrt{nEZ_{jn}^2}} \sum_{j=1}^n Z_{jn} \leq -\sqrt{nu^3D_2^{-1}(x, n)}\bar{W}(x, n)\right). \quad (3.5)$$

Comparing  $W(x)$  and  $\bar{W}(x, n)$ , we get the following useful result:

**Lemma 3.4** *There exists an integer  $N_2(> N_1)$  such that, for any  $n > N_2$  and  $x \in [C_1, C_2]$ ,*

$$W(x) > \frac{1}{n} \implies \bar{W}(x, n) \geq 0 \text{ and } \frac{W(x)}{\bar{W}(x, n)} \leq 2, \quad (3.6)$$

and

$$W(x) < -\frac{1}{n} \implies \bar{W}(x, n) \leq 0 \text{ and } \left|\frac{W(x)}{\bar{W}(x, n)}\right| \leq 2. \quad (3.7)$$

Lemma 3.4 allows us to replace  $\bar{W}(x, n)$  with  $W(x)$  in (3.4) and (3.5). That makes things a little easier since  $W(x)$  does not depend on  $n$  and has a few good properties.

Next we state two general well-known results. One is about the non-uniform estimate of the distance between the distribution of a sum of i.i.d. random variables and the normal distribution; the other is about the normal quantile bounds.

**Result A** *Let  $X_1, X_2, \dots, X_n$  be i.i.d random variables,  $EX_1 = 0$ ,  $EX_1^2 = \sigma^2 > 0$ ,  $E|X_1|^3 < \infty$ . Then for all  $x$*

$$|F_n(x) - \Psi(x)| \leq A \frac{\rho}{\sqrt{n}(1 + |x|)^3}. \quad (3.8)$$

Here  $\Psi(x)$  is the c.d.f. of  $N(0, 1)$ ,  $F_n(x)$  and  $\rho$  are given by

$$F_n(x) = P\left(\frac{1}{\sigma\sqrt{n}} \sum_{j=1}^n X_j \leq x\right), \quad \rho = \frac{E|X_1|^3}{\sigma^3}.$$

**Remark** Result A can be found in Petrov (1975, pp125 Theorem 14) or Michel (1981). Here  $A$  is independent of  $n$ . Michel proved  $A < 30.54$ .

**Result B** Let  $\Psi(x)$  be the c.d.f. of  $N(0, 1)$ . Then for some constant  $B_6 > 0$ ,

$$x > 0 \implies 1 - \Psi(x) = \int_x^\infty \frac{1}{\sqrt{2\pi}} e^{-\frac{x^2}{2}} dx \leq \frac{B_6}{x} e^{-\frac{x^2}{2}}, \quad (3.9)$$

and

$$x < 0 \implies \Psi(x) = \int_{-\infty}^x \frac{1}{\sqrt{2\pi}} e^{-\frac{x^2}{2}} dx \leq \frac{B_6}{|x|} e^{-\frac{x^2}{2}}. \quad (3.10)$$

Combining (3.8), (3.9) and (3.10), we see that, for any fixed  $n$ , if  $x > 0$ ,

$$\begin{aligned} & P\left(\frac{1}{\sigma\sqrt{n}} \sum_{j=1}^n X_j > x\right) \quad (3.11) \\ &= 1 - F_n(x) \\ &\leq 1 - \Psi(x) + A \frac{\rho}{\sqrt{n}(1+|x|)^3} \\ &\leq \frac{B_6}{|x|} e^{-\frac{x^2}{2}} + \frac{A\rho}{\sqrt{n}|x|(1+|x|)^2}, \end{aligned}$$

if  $x < 0$ ,

$$\begin{aligned} & P\left(\frac{1}{\sigma\sqrt{n}} \sum_{j=1}^n X_j \leq x\right) \quad (3.12) \\ &= F_n(x) \\ &\leq \Psi(x) + A \frac{\rho}{\sqrt{n}(1+|x|)^3} \\ &\leq \frac{B_6}{|x|} e^{-\frac{x^2}{2}} + \frac{A\rho}{\sqrt{n}|x|(1+|x|)^2}. \end{aligned}$$

Now, we prove our main result:

**Theorem 3.5** Let  $\delta_n$  be the empirical Bayes test constructed in Section 2. Then we have, as  $n \rightarrow \infty$ ,

$$n^{1-\varepsilon} [R(G, \delta_n) - R(G, \delta)] \rightarrow 0.$$

**Proof.** For convinience, let  $B_7 = \max_{C_1 \leq x \leq C_2} h(x)$  and  $B_8 = \int_{C_1}^{C_2} h(x) dx < \infty$ . From (2.8),

$$\begin{aligned}
R(G, \delta_n) - R(G, \delta) &= \int_{C_1}^{b_0} P_n(W_n(x) \leq 0) W(x) h(x) I_{[0 < W(x) < \frac{1}{n}]} dx & (3.13) \\
&+ \int_{b_0}^{C_2} P_n(W_n(x) > 0) |W(x)| h(x) I_{[-\frac{1}{n} < W(x) < 0]} dx \\
&+ \int_{C_1}^{b_0} P_n(W_n(x) \leq 0) W(x) h(x) I_{[W(x) > \frac{1}{n}]} dx \\
&+ \int_{b_0}^{C_2} P_n(W_n(x) > 0) |W(x)| h(x) I_{[W(x) < -\frac{1}{n}]} dx \\
&\equiv I + II + III + IV.
\end{aligned}$$

Part I and Part II are trivial. Since we have  $|W(x)| \leq \frac{1}{n}$  in both,

$$I \leq \frac{1}{n} \int_{C_1}^{b_0} h(x) dx \leq \frac{1}{n} B_8 \quad (3.14)$$

and

$$II \leq \frac{1}{n} \int_{b_0}^{C_2} h(x) dx \leq \frac{1}{n} B_8. \quad (3.15)$$

Part III and Part IV are a little more complicated. We treat Part III first. Using (3.6) and (3.12), we have

$$\begin{aligned}
III &\leq \int_{C_1}^{b_0} P\left(\frac{1}{\sqrt{nEZ_{jn}^2}} \sum_{j=1}^n Z_{jn} \leq -\sqrt{nu^3 D_2^{-1}(x, n)} \bar{W}(x, n)\right) I_{[W(x) > \frac{1}{n}]} W(x) h(x) dx \\
&\leq \int_{C_1}^{b_0} P\left(\frac{1}{\sqrt{nEZ_{jn}^2}} \sum_{j=1}^n Z_{jn} \leq -\frac{1}{2} \sqrt{nu^3 D_2^{-1}(x, n)} W(x)\right) I_{[W(x) > \frac{1}{n}]} W(x) h(x) dx \\
&\leq \int_{C_1}^{b_0} \frac{2B_6 e^{-nu^3 D_2^{-1}(x, n) W^2(x)/8}}{\sqrt{nu^3 D_2^{-1}(x, n)} W(x)} I_{[W(x) > \frac{1}{n}]} W(x) h(x) dx \\
&\quad + \int_{C_1}^{b_0} \frac{2Au^{-5} D_3(x, n) W(x) h(x) I_{[W(x) > \frac{1}{n}]} dx}{[u^{-3} D_2(x, n)]^{3/2} \sqrt{n} \sqrt{nu^3 D_2^{-1}(x, n)} W(x) [1 + \frac{1}{2} \sqrt{nu^3 D_2^{-1}(x, n)} W(x)]^2} \\
&\leq \int_{C_1}^{b_0} \frac{2B_6 e^{-nu^3 D_2^{-1}(x, n) W^2(x)/8}}{\sqrt{nu^3 D_2^{-1}(x, n)}} I_{[W(x) > \frac{1}{n}]} h(x) dx \\
&\quad + \int_{C_1}^{b_0} \frac{2AD_3(x, n)}{nu^2 D_2(x, n) [1 + \frac{1}{2} \sqrt{nu^3 D_2(x, n)} W(x)]^2} I_{[W(x) > \frac{1}{n}]} h(x) dx \\
&\leq \int_{C_1}^{b_0} \frac{2B_6 e^{-nu^3 D_2^{-1}(x, n) W^2(x)/8}}{\sqrt{nu^3 D_2^{-1}(x, n)}} I_S h(x) dx
\end{aligned}$$

$$\begin{aligned}
& + \int_{C_1}^{b_0} \frac{2B_6}{\sqrt{nu^3 D_2^{-1}(x, n)}} I_{S^c} h(x) dx \\
& + \int_{C_1}^{b_0} \frac{2AD_3(x, n)}{nu^2 D_2(x, n)} h(x) dx \\
& \equiv V + VI + VII,
\end{aligned}$$

where  $S = \{x : e^{-nu^3 D_2^{-1}(x, n) W^2(x)/8} \leq \frac{1}{\sqrt{n}}\}$  and  $S^c = \{x : e^{-nu^3 D_2^{-1}(x, n) W^2(x)/8} > \frac{1}{\sqrt{n}}\}$ . Obviously,

$$VII \leq \frac{2A}{nu^2} \int_{C_1}^{b_0} \frac{D_3(x, n)}{D_2(x, n)} h(x) dx \leq \frac{2A}{nu^2} B_5 B_8. \quad (3.16)$$

By our definition of  $S$ ,

$$V \leq \frac{2B_6}{n\sqrt{u^3}} \int_{C_1}^{b_0} \sqrt{D_2(x, n)} h(x) dx \leq \frac{2}{n\sqrt{u^3}} \sqrt{B_5 B_6 B_8}. \quad (3.17)$$

As for Part VI, we have firstly that

$$\begin{aligned}
e^{-nu^3 D_2^{-1}(x, n) W^2(x)/8} > \frac{1}{\sqrt{n}} & \implies W^2(x) < \frac{4D_2(x, n) \log n}{nu^3} \\
& \implies W^2(x) < \frac{4B_5 \log n}{nu^3} \\
& \implies |W(x)| < \sqrt{\frac{4B_5 \log n}{nu^3}}.
\end{aligned}$$

So  $I_{S^c} \leq I_{\{|W(x)| \leq \eta\}}$ , where  $\eta = \sqrt{\frac{4B_5 \log n}{nu^3}}$ . Note that  $\sqrt{\frac{4B_5 \log n}{nu^3}} \rightarrow 0$  as  $n \rightarrow \infty$  since  $u = n^{-\frac{5}{4}}$ . Let  $N_3 > N_2 (> N_1)$  be an integer and such that for  $n \geq N_3$ ,  $\sqrt{\frac{4B_5 \log n}{nu^3}} < \eta_0$ . From Lemma 3.1, when  $n \geq N_3$ ,

$$\begin{aligned}
VI & \leq \frac{2B_6}{\sqrt{nu^3}} \sqrt{B_5 B_7} \int_{C_1}^{b_0} I_{S^c} dx \\
& \leq \frac{2B_6}{\sqrt{nu^3}} \sqrt{B_5 B_7 B_3} \sqrt{\frac{4B_5 \log n}{nu^3}} \\
& \leq \frac{\sqrt{\log n}}{nu^3} 4B_3 B_5 B_6 B_7.
\end{aligned} \quad (3.18)$$

Combining (3.16), (3.17) and (3.18), we get that when  $n \geq N_3$ ,

$$III \leq \frac{2A}{nu^2} B_5 B_8 + \frac{2}{n\sqrt{u^3}} \sqrt{B_5 B_6 B_8} + \frac{\sqrt{\log n}}{nu^3} 4B_3 B_5 B_6 B_7. \quad (3.19)$$

Now we deal with  $IV$ . Using (3.7) and (3.11), we get

$$IV \leq \int_{b_0}^{C_2} P\left(\frac{1}{\sqrt{n}EZ_{jn}^2} \sum_{j=1}^n Z_{jn} > -\sqrt{nu^3 D_2^{-1}(x, n) \bar{W}(x, n)}\right) I_{[W(x) < -\frac{1}{n}]} |W(x)| h(x) dx$$

$$\begin{aligned}
&\leq \int_{b_0}^{C_2} P\left(\frac{1}{\sqrt{nEZ_{jn}^2}} \sum_{j=1}^n Z_{jn} > \frac{1}{2} \sqrt{nu^3 D_2^{-1}(x, n)} |W(x)|\right) I_{\{|W(x)| < -\frac{1}{n}\}} |W(x)| h(x) dx \\
&\leq \int_{b_0}^{C_2} \frac{2B_6 e^{-nu^3 D_2^{-1}(x, n) W^2(x)/8}}{\sqrt{nu^3 D_2^{-1}(x, n)} |W(x)|} I_{\{|W(x)| < -\frac{1}{n}\}} |W(x)| h(x) dx \\
&\quad + \int_{b_0}^{C_2} \frac{2AD_3(x, n) |W(x)| h(x)}{nu^2 D_2(x, n) |W(x)| \left[1 + \frac{1}{2} \sqrt{nu^3 D_2^{-1}(x, n)} |W(x)|\right]^2} I_{\{|W(x)| < -\frac{1}{n}\}} dx \\
&\leq \int_{b_0}^{C_2} \frac{2B_6 e^{-nu^3 D_2^{-1}(x, n) W^2(x)/8}}{\sqrt{nu^3 D_2^{-1}(x, n)}} I_S h(x) dx \\
&\quad + \int_{b_0}^{C_2} \frac{2B_6}{\sqrt{nu^3 D_2^{-1}(x, n)}} I_{S^c} h(x) dx \\
&\quad + \int_{b_0}^{C_2} \frac{2AD_3(x, n)}{nu^2 D_2(x, n)} h(x) dx \\
&\equiv VIII + IX + X.
\end{aligned}$$

Recalling  $S = \{x : e^{-nu^3 D_2^{-1}(x, n) W^2(x)/8} \leq \frac{1}{\sqrt{n}}\}$  and  $S^c = \{x : e^{-nu^3 D_2^{-1}(x, n) W^2(x)/8} > \frac{1}{\sqrt{n}}\}$ , obviously,

$$VIII \leq \frac{2B_6}{n\sqrt{u^3}} \int_{b_0}^{C_2} \sqrt{D_2(x, n)} h(x) dx \leq \frac{2B_6}{n\sqrt{u^3}} \sqrt{B_5} B_8. \quad (3.20)$$

Similarly to (3.16),

$$X \leq \frac{2A}{nu^2} \int_{b_0}^{C_2} \frac{D_3(x, n)}{D_2(x, n)} h(x) dx \leq \frac{2A}{nu^2} B_5 B_8. \quad (3.21)$$

As for Part IX, similarly to Part VI, we have

$$\begin{aligned}
e^{-nu^3 D_2^{-1}(x, n) W^2(x)/8} > \frac{1}{\sqrt{n}} &\implies W^2(x) < \frac{4D_2(x, n) \log n}{nu^3} \\
&\implies W^2(x) < \frac{4B_5 \log n}{nu^3} \\
&\implies |W(x)| < \sqrt{\frac{4B_5 \log n}{nu^3}}.
\end{aligned}$$

When  $n > N_3$ ,  $\frac{\sqrt{4B_5 \log n}}{\sqrt{nu^3}} \leq \eta_0$  and  $I_{S^c} \leq I_{\{|W(x)| \leq \frac{\sqrt{4B_5 \log n}}{\sqrt{nu^3}}\}}$ . From Lemma 3.1,

$$\begin{aligned}
IX &\leq \frac{2B_6}{\sqrt{nu^3}} \sqrt{B_5} B_7 \int_{b_0}^{C_2} I_{S^c} dx & (3.22) \\
&\leq \frac{2B_6}{\sqrt{nu^3}} \sqrt{B_5} B_7 B_3 \frac{\sqrt{4B_5 \log n}}{\sqrt{nu^3}} \\
&\leq \frac{\sqrt{\log n}}{nu^3} 4B_3 B_5 B_6 B_7.
\end{aligned}$$

Combining (3.20), (3.21) and (3.22), we get that when  $n \geq N_3$ ,

$$IV \leq \frac{2}{n\sqrt{u^3}}\sqrt{B_5B_6B_8} + \frac{2A}{nu^2}B_5B_8 + \frac{\sqrt{\log n}}{nu^3}4B_3B_5B_6B_7. \quad (3.23)$$

From (3.14), (3.15) (3.19) and (3.23), we get

$$R(G, \delta_n) - R(G, \delta) \leq \frac{4}{n\sqrt{u^3}}\sqrt{B_5B_6B_8} + \frac{4A}{nu^2}B_5B_8 + \frac{\sqrt{\log n}}{nu^3}8B_3B_5B_6B_7. \quad (3.24)$$

As  $n \rightarrow \infty$ ,

$$\left\{ \begin{array}{l} n^{1-\varepsilon} \frac{1}{n\sqrt{u^3}} = n^{1-\varepsilon} \frac{1}{nn^{-3\varepsilon/8}} = n^{-5\varepsilon/8} \rightarrow 0; \\ n^{1-\varepsilon} \frac{1}{u^2} = n^{1-\varepsilon} \frac{1}{nn^{-\varepsilon/2}} = n^{-\varepsilon/2} \rightarrow 0; \\ n^{1-\varepsilon} \frac{\sqrt{\log n}}{nu^3} = n^{1-\varepsilon} \frac{\sqrt{\log n}}{nn^{-3\varepsilon/4}} = n^{-\varepsilon/4} \sqrt{\log n} \rightarrow 0. \end{array} \right.$$

We get that  $n^{1-\varepsilon}[R(G, \delta_n) - R(G, \delta)] \rightarrow 0$ .

The proof is completed.

#### § 4 Proofs of Lemmas

**Proof of Lemma 3.1** Since  $\alpha(x) < \infty$  for all  $x$ ,  $\alpha^{(l)}(x)$  exists for all  $x$  and all  $l \geq 1$ . Then  $W^{(l)}(x)$  exists for all  $x$  and all  $l \geq 1$ , since  $\psi(x) = \alpha'(x)$ . These will be used in the proof of Lemma 3.2. Now, we need that

$$W'(x) = \theta_0 \int_{\Omega} \theta c(\theta) e^{\theta x} dG(\theta) - \int_{\Omega} \theta^2 c(\theta) e^{\theta x} dG(\theta). \quad (4.1)$$

First, we prove that

$$W'(b_0) < 0. \quad (4.2)$$

If  $\psi_G(b_0) = 0$ , then  $W'(b_0) = - \int_{\Omega} \theta^2 c(\theta) e^{\theta b_0} dG(\theta) < 0$ . If  $\psi_G(b_0) > 0$ , then

$$\frac{\psi'_G(b_0)}{\psi_G(b_0)} = \frac{\int_{\Omega} \theta^2 c(\theta) e^{\theta b_0} dG(\theta)}{\int_{\Omega} \theta c(\theta) e^{\theta b_0} dG(\theta)} > \frac{\int_{\Omega} \theta c(\theta) e^{\theta b_0} dG(\theta)}{\int_{\Omega} c(\theta) e^{\theta b_0} dG(\theta)} = \theta_0.$$

Thus  $W'(b_0) = \psi_G(b_0)[\theta_0 - \frac{\psi'_G(b_0)}{\psi_G(b_0)}] < 0$ . If  $\psi_G(b_0) < 0$ , then

$$\frac{\psi'_G(b_0)}{\psi_G(b_0)} = \frac{\int_{\Omega} \theta^2 c(\theta) e^{\theta b_0} dG(\theta)}{\int_{\Omega} \theta c(\theta) e^{\theta b_0} dG(\theta)} < \frac{\int_{\Omega} \theta c(\theta) e^{\theta b_0} dG(\theta)}{\int_{\Omega} c(\theta) e^{\theta b_0} dG(\theta)} = \theta_0.$$

Thus  $W'(b_0) = \psi_G(b_0)[\theta_0 - \frac{\psi'_G(b_0)}{\psi_G(b_0)}] < 0$ . Then (4.2) is proved.

From (4.1), we see that  $W'(x)$  is continuous. Then we can find  $b'_0 > 0$  such that, for any  $x \in [b_0 - b'_0, b_0 + b'_0] \subset [C_1, C_2]$ ,

$$-W'(x) \geq \frac{1}{2}[-W'(b_0)] \equiv 2B_3^{-1}. \quad (4.3)$$

Note that  $W(x) > 0$  for  $x \in [C_1, b_0]$ ;  $W(x) < 0$  for  $x \in (b_0, C_2]$ . Then

$$\eta_0 \equiv \left[ \min_{C_1 \leq x \leq b_0 - b'_0} W(x) \right] \wedge \left[ \min_{b_0 + b'_0 \leq x \leq C_2} W(x) \right] > 0,$$

where  $a \wedge b = \min\{a, b\}$ . For  $\eta < \eta_0$ , let

$$\eta_L = \{x : W(x) = \eta, C_1 \leq x \leq C_2\},$$

and

$$\eta_R = \{x : W(x) = -\eta, C_1 \leq x \leq C_2\}.$$

Since  $\eta < \eta_0$ ,  $\eta_L$  and  $\eta_R$  are unique, and

$$[\eta_L, \eta_R] \subset [b_0 - b'_0, b_0 + b'_0].$$

Recall  $W'(x) < 0$  for  $x \in [b_0 - b'_0, b_0 + b'_0]$ . Then  $l(\eta) = \eta_R - \eta_L$ . Using the slope formula and (4.3), we have

$$-\frac{-\eta - \eta}{\eta_R - \eta_L} \geq 2B_3^{-1}.$$

Thus

$$l(\eta) \leq B_3\eta. \quad (4.4)$$

**Proof of Lemma 3.2** Using Taylor's Theorem, (2.1) and (2.2), a straight-forward computation shows that

$$\begin{aligned} & E\left[\frac{\theta_0 K_0\left(\frac{X_j - x}{u}\right)}{uh(X_j)}\right] \\ &= \int_{\Omega} \int_{\alpha}^{\beta} \frac{\theta_0 K_0\left(\frac{y-x}{u}\right)}{uh(y)} c(\theta) e^{\theta y} h(y) dy dG(\theta) \\ &= \theta_0 \int_{\Omega} \int_a^b K_0(t) c(\theta) e^{\theta x} e^{\theta ut} dt dG(\theta) \\ &= \theta_0 \int_{\Omega} c(\theta) e^{\theta x} \left[ \int_a^b K_0(t) e^{\theta ut} dt \right] dG(\theta) \\ &= \theta_0 \int_{\Omega} c(\theta) e^{\theta x} \left[ 1 + \frac{u^m}{m!} \theta^m \int_a^b K_0(t) t^m e^{\theta ut} dt \right] dG(\theta) \\ &= \theta_0 \int_{\Omega} c(\theta) e^{\theta x} dG(\theta) + u^m \int_{\Omega} \theta_0 \theta^m c(\theta) e^{\theta x} \left[ \frac{1}{m!} \int_a^b K_0(t) t^m e^{\theta ut} dt \right] dG(\theta), \end{aligned} \quad (4.5)$$

where  $|t^*| \leq \max\{|a|, |b|\} \equiv c$ . Also,

$$\begin{aligned}
& E\left[\frac{K_1\left(\frac{X_j-x}{u}\right)}{u^2 h(X_j)}\right] \tag{4.6} \\
&= \int_{\Omega} \int_{\alpha}^{\beta} \frac{K_1\left(\frac{y-x}{u}\right)}{u^2 h(y)} c(\theta) e^{\theta y} h(y) dy dG(\theta) \\
&= \frac{1}{u} \int_{\Omega} \int_a^b K_1(t) c(\theta) e^{\theta x} e^{\theta u t} dt dG(\theta) \\
&= \frac{1}{u} \int_{\Omega} c(\theta) e^{\theta x} \left[ \int_a^b K_1(t) e^{\theta u t} dt \right] dG(\theta) \\
&= \frac{1}{u} \int_{\Omega} c(\theta) e^{\theta x} \left[ u\theta + \frac{u^{m+1}}{(m+1)!} \theta^{m+1} \int_a^b K_1(t) t^{m+1} e^{\theta u t^{**}} dt \right] dG(\theta) \\
&= \int_{\Omega} \theta c(\theta) e^{\theta x} dG(\theta) + u^m \int_{\Omega} \theta^{m+1} c(\theta) e^{\theta x} \left[ \frac{1}{(m+1)!} \int_a^b K_1(t) t^{m+1} e^{\theta u t^{**}} dt \right] dG(\theta),
\end{aligned}$$

where  $|t^{**}| \leq \max\{|a|, |b|\} = c$ . From (4.5) and (4.6), we get that

$$E[V(X_j, x, n)] = W(x) + u^m W(x, n),$$

where

$$\begin{aligned}
W(x, n) &= \theta_0 \int_{\Omega} \theta^m c(\theta) e^{\theta x} \left[ \frac{1}{m!} \int_a^b K_0(t) t^m e^{\theta u t^*} dt \right] dG(\theta) \\
&\quad - \int_{\Omega} \theta^{m+1} c(\theta) e^{\theta x} \left[ \frac{1}{(m+1)!} \int_a^b K_1(t) t^{m+1} e^{\theta u t^{**}} dt \right] dG(\theta).
\end{aligned}$$

Choose  $\zeta > 0$  such that  $\alpha < C_1 - \zeta < C_2 + \zeta < \beta$ . Then, we can find an  $N_1$  such that for  $n > N_1$ ,

$$uc \leq \zeta. \tag{4.7}$$

Then

$$\begin{aligned}
|W(x, n)| &\leq B_2 |\theta_0| c^m (b-a) \int_{\Omega} |\theta|^m c(\theta) e^{\theta x} (e^{\theta \zeta} + e^{-\theta \zeta}) dG(\theta) \\
&\quad + B_2 c^{m+1} (b-a) \int_{\Omega} |\theta|^{m+1} c(\theta) e^{\theta x} (e^{\theta \zeta} + e^{-\theta \zeta}) dG(\theta).
\end{aligned}$$

Let  $m_0$  be an even number such that  $m+1 \leq m_0$ . Then

$$\begin{aligned}
& \int_{\Omega} |\theta|^m c(\theta) e^{\theta x} (e^{\theta \zeta} + e^{-\theta \zeta}) dG(\theta) \\
&\leq \int_{\Omega[|\theta| \leq 1]} c(\theta) e^{\theta x} (e^{\theta \zeta} + e^{-\theta \zeta}) dG(\theta) + \int_{\Omega[|\theta| > 1]} \theta^{m_0} c(\theta) e^{\theta x} (e^{\theta \zeta} + e^{-\theta \zeta}) dG(\theta) \\
&\leq \int_{\Omega} c(\theta) e^{\theta x} (e^{\theta \zeta} + e^{-\theta \zeta}) dG(\theta) + \int_{\Omega} \theta^{m_0} c(\theta) e^{\theta x} (e^{\theta \zeta} + e^{-\theta \zeta}) dG(\theta) \\
&= \alpha_G(x + \zeta) + \alpha_G(x - \zeta) + \alpha_G^{(m_0)}(x + \zeta) + \alpha_G^{(m_0)}(x - \zeta).
\end{aligned}$$

Similarly

$$\int_{\Omega} |\theta|^{m+1} c(\theta) e^{\theta x} (e^{\theta \zeta} + e^{-\theta \zeta}) dG(\theta) \leq \alpha_G(x + \zeta) + \alpha_G(x - \zeta) + \alpha_G^{(m_0)}(x + \zeta) + \alpha_G^{(m_0)}(x - \zeta).$$

Since  $\alpha_G^{(l)}$  is continuous, it follows that

$$B_9 = \max_{C_1 - \zeta \leq x \leq C_2 + \zeta} \alpha_G(x) < \infty,$$

and

$$B_{10} = \max_{C_1 - \zeta \leq x \leq C_2 + \zeta} \alpha_G^{(m_0)}(x) < \infty.$$

Thus

$$|W(x, n)| \leq B_2 |\theta_0| c^m (b - a) \times 2(B_9 + B_{10}) + B_2 c^{m+1} (b - a) \times 2(B_9 + B_{10}) \equiv B_4 < \infty.$$

**Proof of Lemma 3.3** Obviously,  $Z_{jn}$  are i.i.d. for fixed  $n$ . A few computations show that

$$\begin{aligned} E Z_{jn}^2 &= \frac{1}{u^4} \int_{\Omega} \int_{\alpha}^{\beta} \left[ \theta_0 u \frac{K_0\left(\frac{y-x}{u}\right)}{h(y)} - \frac{K_1\left(\frac{y-x}{u}\right)}{h(y)} - u^2 \bar{W}(x, n) \right]^2 c(\theta) e^{\theta y} h(y) dy dG(\theta) \\ &= \frac{1}{u^3} \int_{\Omega} \int_a^b \left[ \frac{\theta_0 u K_0(t) - K_1(t)}{h(x+ut)} - u^2 \bar{W}(x, n) \right]^2 c(\theta) e^{\theta x} e^{\theta ut} h(x+ut) dt dG(\theta) \\ &= \frac{1}{u^3} D_2(x, u), \end{aligned}$$

where

$$D_2(x, u) = \int_{\Omega} \int_a^b \left[ \frac{\theta_0 u K_0(t) - K_1(t)}{h(x+ut)} - u^2 \bar{W}(x, n) \right]^2 c(\theta) e^{\theta x} e^{\theta ut} h(x+ut) dt dG(\theta).$$

Also,

$$\begin{aligned} E |Z_{jn}|^3 &= \frac{1}{u^6} \int_{\Omega} \int_{\alpha}^{\beta} \left| \theta_0 u \frac{K_0\left(\frac{y-x}{u}\right)}{h(y)} - \frac{K_1\left(\frac{y-x}{u}\right)}{h(y)} - u^2 \bar{W}(x, n) \right|^3 c(\theta) e^{\theta y} h(y) dy dG(\theta) \\ &= \frac{1}{u^5} \int_{\Omega} \int_a^b \left| \frac{\theta_0 u K_0(t) - K_1(t)}{h(x+ut)} - u^2 \bar{W}(x, n) \right|^3 c(\theta) e^{\theta x} e^{\theta ut} h(x+ut) dt dG(\theta) \\ &= \frac{1}{u^5} D_3(x, u), \end{aligned}$$

where

$$D_3(x, u) = \int_{\Omega} \int_a^b \left| \frac{\theta_0 u K_0(t) - K_1(t)}{h(x+ut)} - u^2 \bar{W}(x, n) \right|^3 c(\theta) e^{\theta x} e^{\theta ut} h(x+ut) dt dG(\theta).$$

When  $n \geq N_1$ , for any  $x \in [C_1, C_2]$  and  $t \in [a, b]$ ,  $x + tu \in [C_1 - \zeta, C_2 + \zeta]$  and

$$\begin{aligned}\bar{W}(x, n) &= W(x) + u^m W(x, n) \\ &\leq \max_{C_1 \leq x \leq C_2} W(x) + u^m B_4 \\ &\leq \max_{C_1 \leq x \leq C_2} W(x) + B_4 \\ &\equiv B_{11}.\end{aligned}$$

Let  $B_{12} = \max_{x \in [C_1 - \zeta, C_2 + \zeta]} \frac{1}{h(x)} < \infty$  and  $B_{13} = \max_{x \in [C_1 - \zeta, C_2 + \zeta]} h(x) < \infty$ . Then, for any  $n \geq N_1$  and any  $x \in [C_1, C_2]$ ,

$$\begin{aligned}\left| \frac{\theta_0 u K_0(t) - K_1(t)}{h(x + ut)} - u^2 \bar{W}(x, n) \right| &\leq (|\theta_0| u B_2 + B_2) B_{12} + u^2 B_{11} \\ &\leq (|\theta_0| + 1) B_2 B_{12} + B_{11} \\ &\equiv B_{14},\end{aligned}$$

and

$$\begin{aligned}&\int_{\Omega} \int_a^b c(\theta) e^{\theta x} e^{\theta ut} h(x + ut) dt dG(\theta) \\ &\leq B_{13} (b - a) \int_{\Omega} c(\theta) e^{\theta x} (e^{\theta \zeta} + e^{-\theta \zeta}) dG(\theta) \\ &\leq 2 B_{13} (b - a) \max_{x \in [C_1 - \zeta, C_2 + \zeta]} \alpha_G(x) \\ &\equiv B_{15}.\end{aligned}$$

Therefore

$$D_2(x, n) \leq B_{14}^2 B_{15}, \quad D_3(x, n) \leq B_{14}^3 B_{15}, \quad \frac{D_3(x, n)}{D_2(x, n)} \leq B_{14}.$$

Letting  $B_5 = \max\{B_{14}, B_{14}^2 B_{15}, B_{14}^3 B_{15}\} < \infty$ , completes the proof.

**Proof of Lemma 3.4** Noting that  $u = n^{-\frac{\varepsilon}{4}}$  and  $m\varepsilon > 4$ ,  $nu^m = n^{-\frac{\varepsilon m - 4}{4}} \rightarrow 0$  as  $n \rightarrow \infty$ . Since  $W(x, n) \leq B_4$  for all  $x \in [C_1, C_2]$  if  $n > N_1$ , there exists an  $N_2 (\geq N_1)$  such that  $|nu^m W(x, n)| \leq \frac{1}{2}$  for all  $x \in [C_1, C_2]$ . If  $W(x) > \frac{1}{n}$ ,

$$n\bar{W}(x, n) = n[W(x) + u^m W(x, n)] = nW(x) + nu^m W(x, n) > 1 - \frac{1}{2} = \frac{1}{2} > 0$$

and

$$\begin{aligned}\frac{W(x)}{\bar{W}(x, n)} &= \frac{nW(x)}{nW(x) + nu^m W(x, n)} \\ &\leq \frac{nW(x) - 1 + 1}{nW(x) - 1 + \frac{1}{2}} \\ &\leq 2.\end{aligned}$$

Then (3.6) is proved. (3.7) can be proved in a similar way.

**Acknowledgment** We would like to thank Prof. T. Liang for his helpful comments and suggestions.

### References

- [1] Johns, M.V., Jr. and Van Ryzin, J.R. (1972). Convergence rates for empirical Bayes two-action problems, II. Continuous Case. **Ann. Math. Statist.** **43**, 934-947.
- [2] Karunamuni, R. J. and Yang, H. (1995). On convergence rates of monotone empirical Bayes tests for the continuous one-parameter exponential family. **Statistics and Decisions**, **13** 181-192.
- [3] Liang, T. (1999). On empirical Bayes tests in a positive exponential family. To appear in **J. Statist. Plann. Inference**.
- [4] Michel, R. (1981). On the constant in the nonuniform version of the Berry-Esseen Theorem. **Z.W.**, **55** 109-117.
- [5] Petrov, V.V. (1975). **Sums of Independent Random Variables**. Springer-verlag, Berlin Heidelberg New York.
- [6] Robbins, H.(1956). An empirical Bayes approach to statistics. **Proc. Third. Berkeley Symp. Math. Statist, Probab.** **1**, 157-163, Univ. California Press, Berkeley.
- [7] Robbins, H.(1964). The empirical Bayes approach to statistical decision problems. **Ann. Math. Statist.** **35** 1-20.
- [8] Stijnen, T.(1985). On the asymptotic behavior of empirical Bayes tests for the continuous one-parameter exponential family. **Ann. Statist.** **13**, 403-412
- [9] Van Houwelingen, J.C.(1976). Monotone empirical Bayes tests for the continuous one-parameter exponential family. **Ann. Statist.** **4**, 981-989.

## REPORT DOCUMENTATION PAGE

Form Approved  
OMB NO. 0704-0188

Public reporting burden for this collection of information is estimated to average 1 hour per response, including the time for reviewing instructions, searching existing data sources, gathering and maintaining the data needed, and completing and reviewing the collection of information. Send comment regarding this burden estimate or any other aspect of this collection of information, including suggestions for reducing this burden, to Washington Headquarters Services, Directorate for Information Operations and Reports, 1215 Jefferson Davis Highway, Suite 1204, Arlington, VA 22202-4302, and to the Office of Management and Budget, Paperwork Reduction Project (0704-0188), Washington, DC 20503.

1. AGENCY USE ONLY (Leave blank)		2. REPORT DATE June 1999	3. REPORT TYPE AND DATES COVERED Technical Report, June 1999	
4. TITLE AND SUBTITLE Empirical Bayes Tests with $n^{-1+\epsilon}$ Convergence Rate in Continuous One-Parameter Exponential Family			5. FUNDING NUMBERS DAAH04-95-1-0165	
6. AUTHOR(S) Shanti S. Gupta and Jianjun Li			8. PERFORMING ORGANIZATION REPORT NUMBER Technical Report #99-09C	
7. PERFORMING ORGANIZATION NAME(S) AND ADDRESS(ES) Purdue University Department of Statistics West Lafayette, IN 47907-1399			10. SPONSORING / MONITORING AGENCY REPORT NUMBER ARO 32922.25-MA	
9. SPONSORING / MONITORING AGENCY NAME(S) AND ADDRESS(ES) U.S. Army Research Office P.O. Box 12211 Research Triangle Park, NC 27709-2211			11. SUPPLEMENTARY NOTES The views, opinions and/or findings contained in this report are those of the author(s) and should not be construed as an official Department of the Army position, policy or decision, unless so designated by other documentation.	
12a. DISTRIBUTION / AVAILABILITY STATEMENT Approved for public release; distribution unlimited.			12 b. DISTRIBUTION CODE	
13. ABSTRACT (Maximum 200 words) Empirical Bayes tests for testing $H_0: \theta \leq \theta_0$ against $H_1: \theta > \theta_0$ in the continuous one-parameter family with density $c(\theta) \exp\{\theta x\}h(x)$ , $-\infty \leq \alpha < x < \beta \leq \infty$ , are considered under the linear loss. Using the assumptions that $\int_{\Omega}  \theta  dG(\theta) < \infty$ and the critical point $b_0$ of a Bayes test falls in some known interval $[C_1, C_2]$ , where $\alpha < C_1 < C_2 < \beta$ , we show that, for any $0 < \epsilon < 1$ , the empirical Bayes tests can be constructed such that they have a convergence rate of order $o(n^{-1+\epsilon})$ , which generalizes the result of Liang (1999) from the positive (one-parameter) exponential family to any continuous one-parameter exponential family.				
14. SUBJECT TERMS Empirical Bayes; exponential family; rate of convergence; regret			15. NUMBER IF PAGES 17	
			16. PRICE CODE	
17. SECURITY CLASSIFICATION OR REPORT UNCLASSIFIED	18. SECURITY CLASSIFICATION OF THIS PAGE UNCLASSIFIED	19. SECURITY CLASSIFICATION OF ABSTRACT UNCLASSIFIED	20. LIMITATION OF ABSTRACT UL	