ON EMPIRICAL BAYES SELECTION RULES FOR NEGATIVE BINOMIAL POPULATIONS*

by

Shanti S. Gupta, Purdue University and
TaChen Liang, Wayne State University

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Department of Statistics
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ON EMPIRICAL BAYES SELECTION RULES FOR NEGATIVE BINOMIAL POPULATIONS*

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Shanti S. Gupta, Purdue University

TaChen Liang, Wayne State University

Abstract

This paper deals with the problem of selecting good negative binomial populations as compared with a standard or a control. The main results are based on the use of the empirical Bayes approach. First we derive the monotone empirical Bayes estimators of the concerned parameters. Based on these estimators, we construct monotone empirical Bayes selection rules. Asymptotic optimality properties of the monotone empirical Bayes estimators and the monotone empirical Bayes selection rules are investigated. The respective convergence rates for the estimation problem and for the selection problem are studied, under some conditions.

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1. Introduction

The empirical Bayes approach in statistical decision theory is appropriate when one is confronted repeatedly and independently with the same decision problem. In such instances, it is reasonable to formulate the component problem in the sequence as Bayes decision problems with respect to an unknown prior distribution on the parameter space, and then use the accumulated observations to improve the decision rule at each stage. This approach is due to Robbins (1956, 1964). During the last twenty-five years, empirical Bayes methods have been studied extensively. Many such empirical Bayes rules have been shown to be asymptotically optimal in the sense that the risk for the nth decision problem converges to the minimum Bayes risk which would have been obtained if the prior distribution was known and the Bayes rule with respect to this prior distribution was used.

Empirical Bayes rules have been derived for subset selection goals by Deely (1965). Recently, Gupta and Hsiao (1983) and Gupta and Leu (1989) have studied empirical Bayes rules for selecting good populations with respect to a standard or a control with the underlying distributions being uniformly distributed. Gupta and Liang (1986, 1988) studied empirical Bayes rules for selecting binomial populations better than a standard or a control and for selecting the best among several binomial populations. In the above-mentioned papers, the authors have assumed that the form of the prior distribution is completely unknown. Hence, those approaches are referred to as nonparametric empirical Bayes. Gupta and Liang (1989a, 1989b) have also studied some other empirical Bayes selection rules, in which they assumed that the form of the prior distributions is known but the distributions depend on certain unknown hyperparameters. This approach is therefore referred to as parametric empirical Bayes. For a general formulation of multiple statistical decision procedures using empirical Bayes approach, reference should be made to Gupta and Liang (1987).

In this paper, we are concerned with the problem of selecting good negative binomial populations with respect to a standard or a control through the nonparametric empirical Bayes approach. The framework of the empirical Bayes selection problem is formulated in Section 2. Monotone empirical Bayes selection rules are proposed in Section 3. The monotone empirical Bayes selection rules are derived based on certain monotone empirical Bayes estimators of the posterior means of the concerned parameters which are also derived in Section 3. Asymptotic optimality properties of the monotone empirical Bayes estimators and the monotone empirical Bayes selection rules are studied in Section 4 and Section 5, respectively.

2. Formulation of the Empirical Bayes Approach

Consider $k+1$ independent negative binomial populations π_i, $i = 0, 1, \ldots, k$. For each i, $i = 0, 1, \ldots, k$, let p_i denote the probability of success for each independent trial from π_i and let X_i denote the number of successes before attaining the rth failure. Then, conditional on $p_i, X_i|p_i$ has a negative binomial distribution with probability function $f_i(x|p_i)$ of the form

$$f_i(x|p_i) = \binom{x+r-1}{r-1} p_i^x (1-p_i)^r, \; x = 0, 1, 2, \ldots \tag{2.1}$$

Let π_0 be the control population. For each $i = 1, \ldots, k$, population π_i is said to be good if $p_i \geq p_0$ and to be bad if $p_i < p_0$, where the control parameter p_0 is either known or unknown. Our goal is to derive empirical Bayes rules to select all good populations and exclude all bad populations. When the control parameter p_0 is known, the empirical Bayes framework can be formulated as follows:

(1) Let $\Omega = \{\underset{\sim}{p}|\underset{\sim}{p} = (p_1, \ldots, p_k),\ p_i \in (0,1) \text{ for } i = 1, \ldots, k\}$ be the parameter space. For each $\underset{\sim}{p} \in \Omega$, define $A(\underset{\sim}{p}) = \{i|p_i \geq p_0\}$ and $B(\underset{\sim}{p}) = \{i|p_i < p_0\}$. That is, $A(\underset{\sim}{p})(B(\underset{\sim}{p}))$ is the set of indices of good (bad) populations. Let $G(\underset{\sim}{p}) = \prod_{i=1}^{k} G_i(p_i)$ be the prior distribution on the parameter space Ω, where $G_i(\cdot)$ are unknown for all $i = 1, \ldots, k$.

(2) Let $\mathcal{A} = \{a|a \subset \{1, 2, \ldots, k\}\}$ be the action space. When action a is taken, it means that population π_i is selected as good if $i \epsilon a$ and excluded as bad if $i \notin a$.

(3) For each (fixed) parameter $\underset{\sim}{p}$ and action a, the loss function $L(\underset{\sim}{p}, a)$ is defined as:

$$L(\underset{\sim}{p}, a) = \sum_{i \in A(\underset{\sim}{p}) \setminus a} (p_i - p_0) + \sum_{i \in a \setminus A(\underset{\sim}{p})} (p_0 - p_i), \tag{2.2}$$

where the first summation is the loss due to not selecting good populations and the second summation is the loss due to selecting bad populations.

(4) For each i, $i = 1, \ldots, k$, let $(X_{ij},\ P_{ij})$, $j = 1, 2, \ldots, n$, be independent random vectors associated with population π_i, where X_{ij} is observable but P_{ij} is not. P_{ij} has prior distribution G_i. Conditional on $P_{ij} = p_{ij}$, $X_{ij}|p_{ij}$ has a negative binomial distribution with parameters r and p_{ij}. Let the jth stage observations be denoted by $\underset{\sim}{X}_j$. That is, $\underset{\sim}{X}_j = (X_{1j}, X_{2j}, \ldots, X_{kj})$. From the assumptions, $\underset{\sim}{X}_1, \underset{\sim}{X}_2, \ldots \underset{\sim}{X}_n$, are mutually independent and identically distributed.

(5) Let $\underset{\sim}{X} = (X_1, \ldots, X_k)$ denote the present observation. Conditional on $\underset{\sim}{p}$, $\underset{\sim}{X}$ has a joint probability function $f(\underset{\sim}{x}|\underset{\sim}{p}) = \prod_{i=1}^{k} f_i(x_i|p_i)$, where $\underset{\sim}{x} = (x_1, \ldots, x_k)$.

Finally, since we are interested in Bayes rules, we can restrict our attention to the nonrandomized selection rules.

(6) Let $D = \{d | d : \mathcal{X} \to \mathcal{A}$, being measurable$\}$ be the set of nonrandomized selection rules, where $\mathcal{X}$ is the sample space of $\underset{\sim}{X}$. For each $d \in D$, let $r(G, d)$ denote the associated Bayes risk. Then $r(G) \equiv \inf_{d \in D} r(G, d)$ is the minimum Bayes risk and a rule, say d_G, is called a Bayes selection rule if $r(G, d_G) = r(G)$.

When p_0 is unknown, the indices in the associated notations should begin at 0 instead of at 1. In the sequel, (0) will be used to show this additional fact.

We now consider decision rules $d_n(\underset{\sim}{x}) \equiv d_n(\underset{\sim}{x}; \underset{\sim}{X}_1, \ldots, \underset{\sim}{X}_n)$. Let $r(G, d_n)$ denote the overall Bayes risk associated with the selection rule $d_n(\underset{\sim}{x})$. That is,

$$r(G, d_n) \equiv \sum_{\underset{\sim}{x} \in \mathcal{X}} E \int_{\Omega} L(\underset{\sim}{p}, d_n(\underset{\sim}{x})) f(\underset{\sim}{x} | \underset{\sim}{p}) dG(\underset{\sim}{p}) \tag{2.3}$$

where the expectation E is taken with respect to $(\underset{\sim}{X}_1, \ldots, \underset{\sim}{X}_n)$.

Definition 2.1. A sequence of selection rules $\{d_n(\underset{\sim}{x})\}_{n=1}^{\infty}$ is said to be asymptotically optimal (a.o.) relative to the prior distribution G if $r(G, d_n) \to r(G)$ as $n \to \infty$.

For constructing a sequence of a.o. empirical Bayes rules, we first need to find the minimum Bayes risk and the associated Bayes rule d_G. From (2.2), the Bayes risk associated with the selection rule d is:

$$r(G, d) = \begin{cases} \sum\limits_{\underset{\sim}{x} \in \mathcal{X}} \sum\limits_{i \in d(\underset{\sim}{x})} [p_0 - \varphi_i(x_i)] f(\underset{\sim}{x}) + C & \text{if } p_0 \text{ is known,} \\ \sum\limits_{\underset{\sim}{x} \in \mathcal{X}} \sum\limits_{i \in d(\underset{\sim}{x})} [\varphi_0(x_0) - \varphi_i(x_i)] f(\underset{\sim}{x}) + C & \text{if } p_0 \text{ is unknown,} \end{cases} \tag{2.4}$$

where $f(\underset{\sim}{x}) = \prod_{i=m}^{k} f_i(x_i)$, $m = 0(1)$ if p_0 is unknown (known), $f_i(x_i) = \int_0^1 f_i(x_i|p) dG_i(p) = \beta(x_i) h_i(x_i)$, $\beta(x_i) = \binom{x_i + r - 1}{r - 1}$, $h_i(x_i) = \int_0^1 p^{x_i}(1-p)^r dG_i(p)$, $\varphi_i(x_i) = \frac{h_i(x_i+1)}{h_i(x_i)}$ (note that $0 < \varphi_i(x_i) < 1$), and $C = \sum_{\underset{\sim}{x} \in \mathcal{X}} \sum_{i=1}^{k} \int_{\Omega} (p_i - p_0) I_{(p_0, 1)}(p_i) f(\underset{\sim}{x} | \underset{\sim}{p}) dG(\underset{\sim}{p})$.

Note that in (2.4), C is a constant and does not affect the determination of the Bayes rule. Thus, the nonrandomized Bayes selection rule d_G can be obtained as follows:

$$d_G(\underset{\sim}{x}) = \{i | \Delta_{iG}(\underset{\sim}{x}) \leq 0\}, \tag{2.5}$$

where

$$\Delta_{iG}(\underset{\sim}{x}) = \begin{cases} p_0 - \varphi_i(x_i) & \text{if } p_0 \text{ is known,} \\ \varphi_0(x_0) - \varphi_i(x_i) & \text{if } p_0 \text{ is unknown.} \end{cases} \tag{2.6}$$

Now, for each $i = (0), 1, \ldots, k$, based on the past data $X_{i1}, \ldots, X_{in}$, and the present observation x_i, let $\varphi_{in}(x_i) \equiv \varphi_{in}(x_i; X_{i1}, \ldots, X_{in})$ be an estimator of $\varphi_i(x_i)$, and let

$$\Delta_{in}(\underset{\sim}{x}) = \begin{cases} p_0 - \varphi_{in}(x_i) & \text{if } p_0 \text{ is known,} \\ \varphi_{0n}(x_0) - \varphi_{in}(x_i) & \text{if } p_0 \text{ is unknown.} \end{cases} \tag{2.7}$$

We then define an empirical Bayes selection rule $d_n(\underset{\sim}{x})$ as follows:

$$d_n(\underset{\sim}{x}) = \{i | \Delta_{in}(\underset{\sim}{x}) \leq 0\}. \tag{2.8}$$

If $\varphi_{in}(x) \xrightarrow{\text{P}} \varphi_i(x)$ for all $x = 0, 1, 2, \ldots$ and $i = (0), 1, \ldots, k$, then $\Delta_{in}(\underset{\sim}{x}) \xrightarrow{\text{P}} \Delta_{iG}(\underset{\sim}{x})$ for all $\underset{\sim}{x} \in \mathcal{X}$. Therefore, from Corollary 2 of Robbins (1964), it follows that $r(G, d_n) \to r(G)$ as $n \to \infty$. So, the sequence of empirical Bayes selection rules $\{d_n(\underset{\sim}{x})\}$ defined in (2.8) is asymptotically optimal. Hence, we have only to find sequences of estimators $\{\varphi_{in}(x)\}$, $i = (0), 1, \ldots, k$, possessing the above mentioned convergence property.

3. The Proposed Empirical Bayes Selection Rules

Before we proceed to construct empirical Bayes estimators $\{\varphi_{in}(x)\}$, we describe some properties of the Bayes selection rule d_G defined in (2.5) and (2.6).

Definition 3.1. Let $\underset{\sim}{x}, \underset{\sim}{y} \in \mathcal{X}$ such that $x_i \leq y_i$ for $i = (0), 1, \ldots, k$.

a) When p_0 is known, a selection rule d is said to be monotone if $d(\underset{\sim}{x}) \subseteq d(\underset{\sim}{y})$.

b) When p_0 is unknown, a selection rule d is said to be monotone if the following two conditions are satisfied: (b1) If $x_0 = y_0$, then $d(\underset{\sim}{x}) \subseteq d(\underset{\sim}{y})$. (b2) If $x_0 < y_0$ and $x_i = y_i$ for all $i = 1, \ldots, k$, then $d(\underset{\sim}{x}) \supseteq d(\underset{\sim}{y})$.

Note that for each $i = (0), 1, \ldots, k$, $\varphi_i(x) = \frac{h_i(x+1)}{h_i(x)}$ is increasing in x. Also, note that $\varphi_i(x)$ is the posterior mean of p_i given $X_i = x$, and the Bayes estimator of p_i given $X_i = x$ for squared error loss. By the monotone property of $\varphi_i(x)$, $i = (0), 1, \ldots, k$, the Bayes selection rule d_G is a monotone selection rule for both cases, where the control parameter p_0 is either known or unknown.

Under the squared error loss, the problem of estimating the probabiltiy of success in a negative binomial distribution is a monotone estimation problem. From Berger (1985), for a monotone estimation problem, the class of monotone estimators form an essentially complete class. Also, under the linear loss given in (2.2), the concerned selection problem is a monotone decision problem. Again, from Berger (1985), the class of monotone selection rules form an essentially complete class. Now, one can see that if the estimators $\varphi_{in}(x)$, $i = (0), 1, \ldots, k$, are monotone, then the empirical Bayes selection rule given in (2.7) and (2.8) is also monotone. From these considerations, it is reasonable to require that the concerned estimators $\{\varphi_{in}(x)\}$ possess the above-mentioned monotone property.

<u>Proposed Monotone Empirical Bayes Selection Rules</u>

Let $F_i(x)$ denote the marginal distribution function of the random variable X_i. Then,

$$F_i(x) = \sum_{y=0}^{x} f_i(y) = \sum_{y=0}^{x} \beta(y) h_i(y), \; x = 0, 1, 2, \ldots \; . \tag{3.1}$$

The form of (3.1) will be used to construct our empirical Bayes estimators.

For each $i = (0), 1, \ldots, k$, and $x = 0, 1, 2, \ldots$ based on the past data $X_{i1}, \ldots, X_{in}$, let

$$f_{in}(x) = \frac{1}{n} \sum_{j=1}^{n} I_{\{x\}}(X_{ij}), \tag{3.2}$$

$$h_{in}(x) = \frac{f_{in}(x)}{\beta(x)}. \tag{3.3}$$

In view of the decreasing property of the function $h_i(y)$, it is desirable that the corresponding estimator possess the same property. To meet this requirement, we let $\{h_{in}^*(x)\}_{x=0}^{\infty}$ be the antitonic regression of $\{h_{in}(x)\}_{x=0}^{\infty}$ with weight $\{\beta(x)\}_{x=0}^{\infty}$. Then, let

$$F_{in}^*(x) = \sum_{y=0}^{x} f_{in}^*(y), \tag{3.4}$$

where

$$f_{in}^*(y) = \beta(y) h_{in}^*(y), y = 0, 1, 2 \ldots. \tag{3.5}$$

Note that $h_{in}^*(x)$ is nonincreasing in x, and $h_{in}^*(x) = 0$ if $x > \max(X_{i1}, \ldots, X_{in})$. By the decreasing property of the function $h_i(x)$, from Barlow, et al (1972),

$$F_{in}^*(x) \geq F_{in}(x) \text{ and } \sup_{x \geq 0} |F_{in}^*(x) - F_i(x)| \leq \sup_{x \geq 0} |F_{in}(x) - F_i(x)| \tag{3.6}$$

where $F_{in}(x) = \frac{1}{n} \sum_{j=1}^{n} I_{[0,x]}(X_{ij}) = \sum_{y=0}^{[x]} \beta(y) h_{in}(y)$.

Let $N_{in} = \max(X_{i1}, \ldots, X_{in}) - 1$. For each $x = 0, 1, 2, \ldots, N_{in}$, define

$$\varphi_{in}(x) = \frac{h_{in}^*(x+1)}{h_{in}^*(x)}. \tag{3.7}$$

Note that $0 \leq \varphi_{in}(x) \leq 1$ for $x = 0, 1, \ldots, N_{in}$. However, $\varphi_{in}(x)$ does not possess the monotone property. A smoothed version of $\varphi_{in}(x)$ is given as follows: Let

$$\varphi_{in}^*(x) = \begin{cases} \max_{y \leq x} \varphi_{in}(y) & \text{if } x \leq N_{in}, \\ \varphi_{in}^*(N_{in}) & \text{if } x > N_{in}. \end{cases} \tag{3.8}$$

From (3.8), it is easy to see that $\varphi_{in}^*(x)$ possesses the monotone property. Now, for each $\underset{\sim}{x} \in \mathcal{X}$, define

$$\Delta_{in}^*(\underset{\sim}{x}) = \begin{cases} p_0 - \varphi_{in}^*(x_i) & \text{if } p_0 \text{ is known,} \\ \varphi_{0n}^*(x_0) - \varphi_{in}^*(x_i) & \text{if } p_0 \text{ is unknown.} \end{cases} \tag{3.9}$$

We then propose a monotone empirical Bayes selection rule, say d_n^*, as follows:

$$d_n^*(\underset{\sim}{x}) = \{i | \Delta_{in}^*(\underset{\sim}{x}) \leq 0\}. \tag{3.10}$$

<u>Asymptotic Optimality of the Selection Rules</u> $\{d_n^*\}$

As mentioned above, to prove the asymptotic optimality of the sequence of empirical Bayes selection rules $\{d_n^*\}$, it suffices to prove that $\varphi_{in}^*(x) \xrightarrow{P} \varphi_i(x)$ for all $x = 0, 1, 2, \ldots$, and $i = (0), 1, \ldots, k$. For each $x = 0, 1, 2, \ldots$, let t be a number such that $0 < t < \min(\varphi_i(x), 1 - \varphi_i(x))$. We need to prove that $P\{|\varphi_{in}^*(x) - \varphi_i(x)| > t\} \to 0$ as $n \to \infty$. Now,

$$\begin{aligned} &P\{|\varphi_{in}^*(x) - \varphi_i(x)| > t\} \\ =&P\{\varphi_{in}^*(x) - \varphi_i(x) < -t, N_{in} \geq x\} + P\{\varphi_{in}^*(x) - \varphi_i(x) > t, N_{in} \geq x\} \\ &\quad + P\{|\varphi_{in}^*(x) - \varphi_i(x)| > t, N_{in} < x\}. \end{aligned} \tag{3.11}$$

By the definition of $N_{in}, P\{|\varphi_{in}^*(x) - \varphi_i(x)| > t, N_{in} < x\} \leq [F_i(x)]^n$ which tends to 0 as $n \to \infty$.

Note by its definition, $\varphi_{in}^*(x) \geq \varphi_{in}(x)$. Also, note that $h_{in}^*(x) > 0$ as $x \leq N_{in} + 1$. Let $p(x,t) = \frac{f_i(x+1)}{f_i(x)} - \frac{t\beta(x+1)}{\beta(x)}$ and $q(x,t) = \frac{t\beta(x+1)f_i(x)}{\beta(x)}$. Then $p(x,t) > 0$ since t is such that $0 < t < \varphi_i(x)$. From (3.1), (3.4) – (3.7), letting $\Delta_i(x,n) = F^*_{in}(x) - F_i(x)$, we have

$$
\begin{aligned}
&P\{\varphi^*_{in}(x) - \varphi_i(x) < -t \text{ and } N_{in} \geq x\} \\
&\leq P\{h^*_{in}(x+1) - h^*_{in}(x)[\varphi_i(x) - t] < 0\} \\
&= P\{F^*_{in}(x+1) - F^*_{in}(x)[1 + p(x,t)] + F^*_{in}(x-1)p(x,t) < 0\} \qquad (3.12) \\
&= P\{\Delta_i(x+1,n) - \Delta_i(x,n)[1 + p(x,t)] + \Delta_i(x-1,n)p(x,t) < -q(x,t)\} \\
&\leq 3P\{\sup_{y \geq 0} |F^*_{in}(y) - F_i(y)| \geq \frac{q(x,t)}{3[1 + p(x,t)]}\} \\
&\leq 3P\{\sup_{y \geq 0} |F_{in}(y) - F_i(y)| \geq \frac{q(x,t)}{3[1 + p(x,t)]}\}
\end{aligned}
$$

which tends to zero as n tends to infinity, by the Glivenko-Cantelli Theorem. Similarly,

$$
\begin{aligned}
&P\{\varphi^*_{in}(x) - \varphi_i(x) > t \text{ and } N_{in} \geq x\} \\
&= P\{\varphi_{in}(y) - \varphi_i(x) > t \text{ for some } y \leq x \text{ and } N_{in} \geq x\} \\
&\leq P\{\varphi_{in}(y) - \varphi_i(y) > t \text{ for some } y \leq x, f^*_{in}(y) > 0\} \qquad (3.13) \\
&\leq P\{f^*_{in}(y+1) - f^*_{in}(y)\frac{\beta(y+1)}{\beta(y)}[\varphi_i(y) + t] > 0 \text{ for some } y \leq x\} \\
&= P\{[F^*_{in}(y+1) - F^*_{in}(y)] - [F^*_{in}(y) - F^*_{in}(y-1)]\frac{\beta(y+1)}{\beta(y)}[\varphi_i(y) + t] > 0 \text{ for some } y \leq x\} \\
&\leq 3P\{\sup_{y \geq 0} |F^*_{in}(y) - F_i(y)| \geq \min_{y \leq x} \frac{|f_i(y+1) - f_i(y)\frac{\beta(y+1)}{\beta(y)}[\varphi_i(y) + t]|}{3[1 + \frac{\beta(y+1)}{\beta(y)}[\varphi_i(y) + t]]}\},
\end{aligned}
$$

which tends to zero as n tends to infinity.

Based on the above discussions, we have shown that $\varphi^*_{in}(x) \xrightarrow{P} \varphi_i(x)$ for each $x = 0, 1, \ldots$ and each $i = (0), 1, \ldots, k$. Therefore, the sequence of empirical Bayes selection rules $\{d^*_n\}$ is asymptotically optimal.

4. Asymptotic Optimality of the Monotone Estimators

In this section, we study the asymptotic optimality property of the estimators $\varphi_{in}^*(x)$. It is known that under the squared error loss, $\varphi_i(x)$ is the Bayes estimator of p_i given $X_i = x$. The associated Bayes risk is $R_i(G_i) = E[(P_i - \varphi_i(X_i))^2]$. Let $\psi_i(\cdot)$ be any estimator of p_i with the associated Bayes risk $R_i(G_i, \psi_i)$. Then

$$R_i(G_i, \psi_i) - R_i(G_i) = E[(\psi_i(X_i) - \varphi_i(X_i))^2]. \tag{4.1}$$

Let $\{\psi_{in}(x; X_{i1}, \ldots, X_{in}) \equiv \psi_{in}(x)\}$ be a sequence of empirical Bayes estimators based on $(x; X_{i1}, \ldots, X_{in})$.

Definition 4.1.

a) A sequence of empirical Bayes estimators $\{\psi_{in}\}_{n=1}^{\infty}$ is asymptotically optimal relative to the prior distribution G_i if $R_i(G_i,\ \psi_{in}) \to R_i(G_i)$ as $n \to \infty$.

b) A sequence of empirical Bayes estimators $\{\psi_{in}\}_{n=1}^{\infty}$ is asymptotically optimal of order α_n relative to the prior distribution G_i if $R_i(G_i,\ \psi_{in}) - R_i(G_i) \le O(\alpha_n)$ as $n \to \infty$, where $\{\alpha_n\}$ is a sequence of positive values such that $\lim_{n\to\infty} \alpha_n = 0$.

From Section 3, $\varphi_{in}^*(x) \xrightarrow{P} \varphi_i(x)$ for all $x = 0, 1, 2, \ldots$. Thus, $\{\varphi_{in}^*\}_{n=1}^{\infty}$ is asymptotically optimal. However, the usefulness of empirical Bayes estimators in practical applications clearly depends on the convergence rates with which the risks for the successive estimation problems approach the optimal Bayes risk. Hence in the following, we study the convergence rates of the sequence of empirical Bayes estimators $\{\varphi_{in}^*\}$.

For $0 < \varepsilon < 1$, let $A_i(\varepsilon) = \{x | f_i(x) < \varepsilon\},\ B_i(\varepsilon) = \{x | f_i(x) \ge \varepsilon\}$.

<u>Assumption A</u>: A1. There exist $t_i \epsilon (0, 1]$ and a positive constant c_i such that $P(A_i(\varepsilon)) \le c_i \varepsilon^{t_i}$ for all $\varepsilon \in (0, 1)$. A2. There exists a positive integer N_i such that $f_i(x)$ is decreasing in x for $x \ge N_i$.

Remark 4.1.: An example where Assumption A holds is given in Lin (1972).

Theorem 4.1. Let $\{\varphi_{in}^*\}_{n=1}^{\infty}$ be the sequence of empirical Bayes estimators defined in (3.8). Then, under Assumption A, $R_i(G_i, \varphi_{in}^*) - R_i(G_i) \leq O(n^{-t_i/(2+t_i)})$.

The proof of Theorem 4.1 can be obtained based on the following arguments.

For the empirical Bayes estimator φ_{in}^*, straight computation leads to

$$\begin{aligned} 0 &\leq R_i(G_i, \varphi_{in}^*) - R_i(G_i) \\ &= \sum_{x \in A_i(\delta_n)} E[(\varphi_{in}^*(x) - \varphi_i(x))^2 | X_i = x] f_i(x) \\ &\quad + \sum_{x \in B_i(\delta_n)} E[(\varphi_{in}^*(x) - \varphi_i(x))^2 | X_i = x] f_i(x), \end{aligned} \tag{4.2}$$

where $\delta_n = n^{-\alpha_i}$ and $\alpha_i = \frac{1}{2+t_i}$.

Lemma 4.1. Under Assumption A,

$$\sum_{x \in A_i(\delta_n)} E[(\varphi_{in}^*(x) - \varphi_i(x))^2 | X_i = x] f_i(x) \leq O(n^{-t_i/(2+t_i)}).$$

Proof: Note that $0 \leq \varphi_{in}^*(x)$, $\varphi_i(x) \leq 1$. Thus, by the definition of $A_i(\delta_n)$ and Assumption A1,

$$\sum_{x \in A_i(\delta_n)} E[(\varphi_{in}^*(x) - \varphi_i(x))^2 | X_i = x] f_i(x) \leq P(A_i(\delta_n)) \leq c_i n^{-\alpha_i t_i} = O(n^{-t_i/(2+t_i)}). \quad \square$$

Now, for each $x \in B_i(\delta_n)$, by letting $I_{in}(x) = 1(0)$ if $f_{in}(x) > (=)0$, we obtain

$$\begin{aligned} &E[(\varphi_{in}^*(x) - \varphi_i(x))^2 | X_i = x] \\ &= E[(\varphi_{in}^*(x) - \varphi_i(x))^2 I_{in}(x) | X_i = x] + E[(\varphi_{in}^*(x) - \varphi_i(x))^2 (1 - I_{in}(x)) | X_i = x], \end{aligned} \tag{4.3}$$

Lemma 4.2. For $x \in B_i(\delta_n)$,

$$E[(\varphi_{in}^*(x) - \varphi_i(x))^2 (1 - I_{in}(x)) | X_i = x] \leq O(n^{-t_i/(2+t_i)}).$$

Proof:

$$
\begin{aligned}
&E[(\varphi^*_{in}(x)-\varphi_i(x))^2(1-I_{in}(x))|X_i=x] \\
&\le P\{f_{in}(x)-f_i(x)\le -f_i(x)\} \\
&\le \exp\{-2nf_i^2(x)\} \quad \text{(by Theorem 1 of Hoeffding (1963))} \\
&\le \exp\{-2n\delta_n^2\} \quad (\text{since } x\in B_i(\delta_n)) \\
&\le O(n^{-t_i/(2+t_i)}). \qquad \square
\end{aligned}
$$

Next, a straight computation leads to:

$$
\begin{aligned}
&E[(\varphi^*_{in}(x)-\varphi_i(x))^2 I_{in}(x)|X_i=x] \\
&=\int_0^{\varphi_i(x)} 2sP\{\varphi^*_{in}(x)-\varphi_i(x)<-s, f_{in}(x)>0\}ds \\
&\quad+\int_0^{1-\varphi_i(x)} 2sP\{\varphi^*_{in}(x)-\varphi_i(x)>s, f_{in}(x)>0\}ds.
\end{aligned} \tag{4.4}
$$

Lemma 4.3. a) For $x\in B_i(\delta_n)$, and $s\in(0,\varphi_i(x))$,

$$P\{\varphi^*_{in}(x)-\varphi_i(x)<-s \text{ and } f_{in}(x)>0\}\le 3d_i\exp\{-2n[\frac{q(x,s)}{3(1+p(x,s))}]^2\}$$

for some positive constant d_i, where $p(x,s)=\frac{f_i(x+1)}{f_i(x)}-\frac{s\beta(x+1)}{\beta(x)}>0$ and $q(x,s)=\frac{s\beta(x+1)f_i(x)}{\beta(x)}>0$.

b) For $x\in B_i(\delta_n)$, $\int_0^{\varphi_i(x)}2sP\{\varphi^*_{in}(x)-\varphi_i(x)<-s \text{ and } f_{in}(x)>0\}ds\le O(n^{-t_i/(2+t_i)})$ and the upper bound is independent of x for all $x\in B_i(\delta_n)$.

Proof: a) First, it is trivial that $p(x,s)>0$ and $q(x,s)>0$. Next, from the definitions of $f^*_{in}(x)$ and $h^*_{in}(x)$ and that $f_{in}(x)>0$, it follows that $f^*_{in}(x)>0$ and $h^*_{in}(x)>0$. Also, $\varphi^*_{in}(x)\ge\varphi_{in}(x)$, by the definition of $\varphi^*_{in}(x)$. Then, following (3.12), we obtain: For $x\in B_i(\delta_n)$ and $s\in(0,\varphi_i(x))$,

$$
\begin{aligned}
P\{\varphi^*_{in}(x)-\varphi_i(x)<-s \text{ and } f_{in}(x)>0\}&\le 3P\{\sup_{y\ge 0}|F_{in}(y)-F_i(y)|\ge\frac{q(x,s)}{3[1+p(x,s)]}\} \\
&\le 3d_i\exp\{-2n[\frac{q(x,s)}{3[1+p(x,s)]}]^2\},
\end{aligned}
$$

where the last inequality follows from Lemma 2.1 of Schuster (1969).

b) By using the fact that $0 < \frac{h_i(x+1)}{h_i(x)} < 1$ and $\frac{\beta(x+1)}{\beta(x)} \leq r$ for all $x \geq 0$, we have $1 + p(x,s) \leq 1 + r$. Then, from the result of part a) of this Lemma, we obtain

$$\begin{aligned}
&\int_0^{\varphi_i(x)} 2sP\{\varphi_{in}^*(x) - \varphi_i(x) < -s, f_{in}(x) > 0\}ds \\
&\leq \int_0^{\varphi_i(x)} 6d_i\ s \exp\{-\frac{2n}{9}[\frac{sf_i(x)\frac{\beta(x+1)}{\beta(x)}}{1+r}]^2\}ds \\
&\leq 13.5d_i\frac{(1+r)^2\beta^2(x)}{\beta^2(x+1)} \times \frac{1}{nf_i^2(x)} \\
&\leq 13.5d_i(1+r)^2\frac{1}{n\delta_n^2} \quad \text{(since } x \in B_i(\delta_n)) \text{ and therefore, } f_i(x) \geq \delta_n) \\
&= O(n^{-t_i/(2+t_i)}).
\end{aligned}$$

Note this upper bound is independent of $x \in B_i(\delta_n)$. □

By Assumption A2, $f_i(x)$ is decreasing in x for all $x \geq N_i$. In the following, we only consider the case where n is large enough such that $\delta_n \equiv n^{-\frac{1}{2+t_i}} \leq f_i(y)$ for all $y \leq N_i$. Thus as $x \in B_i(\delta_n)$, then $f_i(y) \geq \delta_n$ for all $y \leq x$ (this holds true for either $x \leq N_i$ or $x > N_i$). Therefore, analogous to (3.13), we obtain: For $s \in (0, 1 - \varphi_i(x))$,

$$\begin{aligned}
&P\{\varphi_{in}^*(x) - \varphi_i(x) > s, f_{in}(x) > 0\} \\
&\leq 3P\{\sup_{y \geq 0} |F_{in}^*(y) - F_i(y)| \geq -\min_{y \leq x} H(y)\} \\
&\leq 3d_i\ \exp\{-2n[-\min_{y \leq x} H(y)]^2\} \text{ (by Lemma 2.1 of Schuster (1969)),} \qquad (4.5)
\end{aligned}$$

where

$$H(y) = \frac{-sf_i(y)\frac{\beta(y+1)}{\beta(y)}}{3[1+\frac{\beta(y+1)}{\beta(y)}]} \leq \frac{-sf_i(y)}{3[1+r]} \leq \frac{-s\delta_n}{3[1+r]} \ \text{(since } f_i(y) \geq \delta_n) < 0. \qquad (4.6)$$

Lemma 4.4. For n sufficiently large, and $x \in B_i(\delta_n)$,

$$\int_0^{1-\varphi_i(x)} 2sP\{\varphi_{in}^*(x) - \varphi_i(x) > s, f_{in}(x) > 0\}ds \leq O(n^{-t_i/(2+t_i)}).$$

Proof: From (4.5) and (4.6), for n sufficiently large, as $x \in B_i(\delta_n)$,

$$\begin{aligned}
&\int_0^{1-\varphi_i(x)} 2sP\{\varphi_{in}^*(x) - \varphi_i(x) > s, f_{in}(x) > 0\}ds \\
&\leq \int_0^{1-\varphi_i(x)} 6sd_i \exp\{-\frac{2n\delta_n^2 s^2}{9(1+r)^2}\}ds \\
&= O(n^{-t_i/(2+t_i)}).
\end{aligned}$$

□

From Lemmas 4.2, 4.3, 4.4 and (4.4), we have: For $x \in B_i(\delta_n)$,

$$E[(\varphi_{in}^*(x) - \varphi_i(x))^2 | X_i = x] \leq O(n^{-t_i/(2+t_i)}).$$

This upper bound is independent of $x \in B_i(\delta_n)$. Therefore, we conclude that

$$\sum_{x \in B_i(\delta_n)} E[(\varphi_{in}^*(x) - \varphi_i(x))^2 | X_i = x] f_i(x) \leq O(n^{-t_i/(2+t_i)}). \tag{4.7}$$

Then, Lemma 4.1, (4.2) and (4.7) together complete the proof of Theorem 4.1.

5. Asymptotic Optimality of the Empirical Bayes Selection Rules

Let $\{d_n\}_{n=1}^{\infty}$ be a sequence of empirical Bayes selection rules relative to the prior distribution G. Let $r(G, d_n)$ denote the expected Bayes risk of the selection rule d_n. Since the Bayes rule d_G achieves the minimum Bayes risk $r(G)$, $r(G, d_n) - r(G) \geq 0$ for all $n = 1, 2, \ldots$. Thus, the nonnegative difference $r(G, d_n) - r(G)$ is used as a measure of the optimality of the sequence of empirical Bayes selection rules $\{d_n\}$.

Definition 5.1. The sequence of empirical Bayes selection rules $\{d_n\}_{n=1}^{\infty}$ is asymptotically optimal of order α_n relative to the prior distribution G if $r(G, d_n) - r(G) \leq O(\alpha_n)$ as $n \to \infty$, where $\{\alpha_n\}$ is a sequence of positive numbers such that $\lim_{n \to \infty} \alpha_n = 0$.

In the following, we evaluate the asymptotic behavior of the sequence of empirical Bayes rules $\{d_n^*\}$ according to whether the control parameter p_0 is known or unknown.

Convergence Rates of $\{d_n^*\}$ for p_0 Known Case

For each $i = 1, \ldots, k$, let $S_i = \{x | \varphi_i(x) < p_0\}$ and $T_i = \{x | \varphi_i(x) > p_0\}$. Define

$$M_i = \begin{cases} \min T_i & \text{if } T_i \neq \phi, \\ \infty & \text{if } T_i = \phi, \end{cases} \tag{5.1}$$

$$m_i = \begin{cases} \max S_i & \text{if } S_i \neq \phi, \\ -1 & \text{if } S_i = \phi. \end{cases} \tag{5.2}$$

By the increasing property of $\varphi_i(x)$ with respect to the variable $x = 0, 1, \ldots, m_i \leq M_i$; also $m_i < M_i$ if $T_i \neq \phi$. Furthermore, $x \leq m_i$ iff $\varphi_i(x) < p_0$ and $y \geq M_i$ iff $\varphi_i(y) > p_0$.

Note that when $T_i = \phi$, it means that in terms of its quality, the population π_i is bad. Also, for $S_i = \phi$, it means that in terms of its quality population π_i is as good as the control. We exclude these extreme cases, in the following, and study the asymptotic behavior of the sequence of empirical Bayes selection rules $\{d_n^*\}$ under Assumption B.

Assumption B: $T_i \neq \phi$ and $S_i \neq \phi$ for all $i = 1, \ldots, k$.

Now, for the empirical Bayes selection rule d_n^*, we have

$$0 \leq r(G, d_n^*) - r(G) = \sum_{i=1}^{k} D_i(G, d_n^*), \tag{5.3}$$

where

$$D_i(G, d_n^*) = \sum_{x=0}^{m_i} [p_0 - \varphi_i(x)] P\{\varphi_{in}^*(x) \geq p_0\} f_i(x) + \sum_{x=M_i}^{\infty} [\varphi_i(x) - p_0] P\{\varphi_{in}^*(x) < p_0\} f_i(x). \tag{5.4}$$

By the nondecreasing property of the estimator φ_{in}^*, we have

$$\begin{cases} P\{\varphi_{in}^*(x) \geq p_0\} \leq P\{\varphi_{in}^*(m_i) \geq p_0\} & \text{for all } x \leq m_i, \text{ and} \\ P\{\varphi_{in}^*(y) < p_0\} \leq P\{\varphi_{in}^*(M_i) < p_0\} & \text{for all } y \geq M_i. \end{cases} \tag{5.5}$$

Thus, from (5.4) and (5.5),

$$D_i(G, d_n^*) \leq P\{\varphi_{in}^*(m_i) \geq p_0\} b_{i1} + P\{\varphi_{in}^*(M_i) < p_0\} b_{i2}, \tag{5.6}$$

where $b_{i1} = \sum_{x=0}^{m_i} [p_0 - \varphi_i(x)] f_i(x)$, and $b_{i2} = \sum_{x=M_i}^{\infty} [\varphi_i(x) - p_0] f_i(x)$ and $0 \le b_{i1},\ b_{i2} \le 1$.

Now,

$$
\begin{aligned}
&P\{\varphi^*_{in}(M_i) < p_0\} \\
&= P\{\varphi^*_{in}(M_i) < p_0, N_{in} < M_i\} + P\{\varphi^*_{in}(M_i) < p_0, N_{in} \ge M_i\} \qquad (5.7)\\
&\le [F_i(M_i)]^n + P\{\varphi^*_{in}(M_i) < p_0, N_{in} \ge M_i\}
\end{aligned}
$$

Analogous to Lemma 4.3. a), we can obtain

$$
\begin{aligned}
&P\{\varphi^*_{in}(M_i) < p_0 \text{ and } N_{in} \ge M_i\} \\
&\le P\{h^*_{in}(M_i+1) < h^*_{in}(M_i) p_0 \text{ and } N_{in} \ge M_i\} \\
&\le 3P\{\sup_{y \ge 0} |F^*_{in}(y) - F_i(y)| > \frac{\Delta(M_i, p_0)}{3[1 + \frac{\beta(M_i+1)}{\beta(M_i)} p_0]}\} \\
&\le 3d_i \exp\{-\frac{2}{9} n \Delta^2(M_i, p_0) / [1 + \frac{\beta(M_i+1)}{\beta(M_i)} p_0]^2\}, \qquad (5.8)
\end{aligned}
$$

where $\Delta(M_i, p_0) = f_i(M_i+1) - f_i(M_i) \frac{\beta(M_i+1)}{\beta(M_i)} p_0 > 0$.

From (5.7) and (5.8), by letting $\tau_{i1} = \min\left(\frac{2\Delta^2(M_i,p_0)}{9[1+\frac{\beta(M_i+1)}{\beta(M_i)} p_0]^2},\ \ell n \frac{1}{F_i(M_i)} \right)$, we have

$$
P\{\varphi^*_{in}(M_i) < p_0\} \le O(\exp(-\tau_{i1} n)). \qquad (5.9)
$$

Also,

$$
\begin{aligned}
&P\{\varphi^*_{in}(m_i) \ge p_0\} \\
&= P\{\varphi^*_{in}(m_i) \ge p_0, N_{in} < m_i\} + P\{\varphi^*_{in}(m_i) \ge p_0, N_{in} \ge m_i\} \\
&\le [F_i(m_i)]^n + P\{\varphi^*_{in}(m_i) \ge p_0, N_{in} \ge m_i\} \qquad (5.10)
\end{aligned}
$$

Analogous to (3.13), we obtain

$$\begin{aligned}
&P\{\varphi_{in}^*(m_i) \geq p_0,\ N_{in} \geq m_i\} \\
&= P\{\varphi_{in}(y) \geq p_0 \text{ for some } y \leq m_i, N_{in} \geq m_i\} \\
&\leq 3P\{\sup_{x\geq 0} |F_{in}^*(x) - F_i(x)| \geq \min_{y\leq m_i} \frac{|\Delta(y,p_0)|}{3[1+\frac{\beta(y+1)}{\beta(y)}p_0]}\} \\
&\leq 3d_i \exp\left\{-\frac{2n}{9}\min_{y\leq m_i}\left[\frac{|\Delta(y,p_0)|}{1+\frac{\beta(y+1)}{\beta(y)}p_0}\right]^2\right\},
\end{aligned} \tag{5.11}$$

where $\Delta(y,p_0) = f_i(y+1) - f_i(y)\frac{\beta(y+1)}{\beta(y)}p_0$. Note that $\Delta(y,p_0) < 0$ for all $y \leq m_i$.

Thus, by letting $\tau_{i2} = \min\left(\frac{2}{9}\min_{y\leq m_i}\left\{\frac{|\Delta(y,p_0)|^2}{[1+\frac{\beta(y+1)}{\beta(y)}p_0]^2}\right\}, \ell n\frac{1}{F_i(m_i)}\right)$, we have

$$P\{\varphi_{in}^*(m_i) \geq p_0\} \leq O(\exp\{-\tau_{i2}n\}). \tag{5.12}$$

Let $\tau_i = \min(\tau_{i1}, \tau_{i2})$, and $\tau = \min(\tau_1, \ldots, \tau_k)$. Note that $\tau > 0$, since $\tau_i > 0$ for each $i = 1, \ldots, k$. From the above results the following theorem follows:

Theorem 5.1. Under Assumption B, we have:

a) $D_i(G, d_n^*) \leq O(\exp(-\tau_i n))$ for each $i = 1 \ldots k$, and

b) $r(G, d_n^*) - r(G) \leq O(\exp(-\tau n))$.

Convergence Rates of $\{d_n^*\}$ for p_0 Unknown Case

When the parameter p_0 is unknown, the convergence rates of the sequence of empirical Bayes selection rules $\{d_n^*\}_{n=1}^\infty$ is evaluated under Assumption A. Without loss of generality, in this section, we assume that $c_i = c > 0$, and $t_i = t \in (0,1]$ for all $i = 0, 1, \ldots, k$, where the parameters $c_i,\ t_i,\ i = 0, 1 \ldots k$ are given in Assumption A1.

For each $i = 1, \ldots, k$, let $\psi(x_o, x_i) = \varphi_i(x_i) - \varphi_0(x_0)$ and let $S_i = \{(x_0, x_i) | \psi(x_o, x_i) < 0\}, T_i = \{(x_0, x_i) | \psi(x_0, x_i) \geq 0\}, E_{in} = \{(x_0, x_i) || \psi(x_0, x_i)| \leq \varepsilon_n\}$ and $E_{in}^c = \{(x_0, x_i)|$

$|\psi(x_0,x_i)| > \varepsilon_n\}$ where $\varepsilon_n > 0$. Also, let $I_1 = T_i \cap E_{in}, I_2 = T_i \cap E_{in}^c \cap A_0(\delta_n) \cap A_i(\delta_n), I_3 = T_i \cap E_{in}^c \cap A_0(\delta_n) \cap B_i(\delta_n), I_4 = T_i \cap E_{in}^c \cap B_0(\delta_n) \cap A_i(\delta_n), I_5 = T_i \cap E_{in}^c \cap B_0(\delta_n) \cap B_i(\delta_n), J_1 = S_i \cap E_{in}, J_2 = S_i \cap E_{in}^c \cap A_0(\delta_n) \cap A_i(\delta_n), J_3 = S_i \cap E_{in}^c \cap A_0(\delta_n) \cap B_i(\delta_n), J_4 = S_i \cap E_{in}^c \cap B_0(\delta_n) \cap A_i(\delta_n)$ and $J_5 = S_i \cap E_{in}^c \cap B_0(\delta_n) \cap B_i(\delta_n)$. Let

$$\begin{cases} Q_i(x_0,x_i) = [\varphi_i(x_i) - \varphi_0(x_0)]P\{\varphi_{in}^*(x_i) < \varphi_{0n}^*(x_0)\}f_i(x_i)f_0(x_0) \\ R_i(x_0,x_i) = [\varphi_0(x_0) - \varphi_i(x_i)]P\{\varphi_{in}^*(x_i) \geq \varphi_{0n}^*(x_0)\}f_i(x_i)f_0(x_0) \end{cases}$$

and let the summations $\sum\limits_{I_j} Q_i(x_0,x_i)$ and $\sum\limits_{J_j} R_i(x_0,x_i)$ be denoted by $Q_i(x_0,x_i,I_j)$ and $R_i(x_0,x_i,J_j)$, respectively. Thus,

$$0 \leq r(G,d_n^*) - r(G) = \sum_{i=1}^{k} D_i^*(G,d_n^*) \tag{5.13}$$

where $D_i^*(G,d_n^*) = \sum\limits_{j=1}^{5} Q_i(x_0,x_i,I_j) + \sum\limits_{j=1}^{5} R_i(x_0,x_i,J_j)$.

Careful examination leads to the following results: $Q_i(x_0,x_i,I_1) \leq O(\varepsilon_n), R_i(x_0,x_i,J_1) \leq O(\varepsilon_n), Q_i(x_0,x_i,I_2) \leq O(\delta_n^{2t}), R_i(x_0,x_i,J_2) \leq O(\delta_n^{2t}), Q_i(x_0,x_i,I_3) \leq O(\delta_n^t), R_i(x_0,x_i,I_3) \leq O(\delta_n^t), Q_i(x_0,x_i,I_4) \leq O(\delta_n^t)$ and $R_i(x_0,x_i,J_4) \leq O(\delta_n^t)$.

Now, for $(x_0,x_i) \in T_i \cap E_{in}^c \cap B_0(\delta_n) \cap B_i(\delta_n)$, $f_i(x_i) \geq \delta_n$, $f_0(x_0) \geq \delta_n$, and $\varepsilon_n < \varphi_i(x_i) - \varphi_0(x_0) < 1$. Thus,

$$\begin{aligned} & P\{\varphi_{in}^*(x_i) < \varphi_{0n}^*(x_0)\} \\ & = P\{[\varphi_{in}^*(x_i) - \varphi_i(x_i)] - [\varphi_{0n}^*(x_0) - \varphi_0(x_0)] < \varphi_0(x_0) - \varphi_i(x_i)\} \\ & \leq P\{\varphi_{in}^*(x_i) - \varphi_i(x_i) < -\frac{\varepsilon_n}{2}\} + P\{\varphi_{0n}^*(x_0) - \varphi_0(x_0) > \frac{\varepsilon_n}{2}\}. \end{aligned} \tag{5.14}$$

Now,

$$\begin{aligned} & P\{\varphi_{in}^*(x_i) - \varphi_i(x_i) < -\frac{\varepsilon_n}{2}\} \\ & = P\{\varphi_{in}^*(x_i) - \varphi_i(x_i) < -\frac{\varepsilon_n}{2}, f_{in}(x_i) = 0\} + P\{\varphi_{in}^*(x_i) - \varphi_i(x_i) < -\frac{\varepsilon_n}{2}, f_{in}(x_i) > 0\}, \end{aligned}$$

where

$$P\{\varphi_{in}^*(x_i) - \varphi_i(x_i) < -\frac{\varepsilon_n}{2}, f_{in}(x_i) = 0\} \le P\{f_{in}(x_i) = 0\} \le \exp\{-2n\delta_n^2\}$$

(since $x_i \in B_i(\delta_n)$; and see Lemma 4.2),

and

$$\begin{aligned} &P\{\varphi_{in}^*(x_i) - \varphi_i(x_i) < -\frac{\varepsilon_n}{2}, f_{in}(x_i) > 0\} \\ &\le 3d_i \exp\{-\frac{\beta^2(x_i+1)}{18(1+r)^2\beta^2(x_i)} n\varepsilon_n^2\delta_n^2\} \\ &\le O(\exp\{-\frac{n\varepsilon_n^2\delta_n^2}{18(1+r)^2}\}). \text{ (The proof is analogous to that of Lemma 4.3.a).} \end{aligned}$$

Therefore,

$$P\{\varphi_{in}^*(x_i) - \varphi_i(x_i) < -\frac{\varepsilon_n}{2}\} \le O(\exp\{-\frac{n\varepsilon_n^2\delta_n^2}{18(1+r)^2}\}). \tag{5.15}$$

Next, under Assumption A2, along the line of (4.5) and the argument given there, for n sufficiently large, we have

$$P\{\varphi_{0n}^*(x_0) - \varphi_0(x_0) > \frac{\varepsilon_n}{2}\} \le O(\exp\{-\frac{n\varepsilon_n^2\delta_n^2}{18(1+r)^2}\}). \tag{5.16}$$

Note that the convergence rates obtained at (5.15) and (5.16) are independent of $(x_0, x_i) \in T_i \cap E_{in}^c \cap B_0(\delta_n) \cap B_i(\delta_n)$. Therefore, from (5.14)

$$\sum_{T_i \cap E_{in}^c \cap B_0(\delta_n) \cap B_i(\delta_n)} Q_i(x_0, x_i) \le O(\exp\{-\frac{n\varepsilon_n^2\delta_n^2}{18(1+r)^2}\}). \tag{5.17}$$

Similarly, we can also conclude that

$$\sum_{S_i \cap E_{in}^c \cap B_0(\delta_n) \cap B_i(\delta_n)} R_i(x_0, x_i) \le O(\exp\{-\frac{n\varepsilon_n^2\delta_n^2}{18(1+r)^2}\}). \tag{5.18}$$

By letting $\delta_n = [\frac{18(1+r)^2 \ell n\ n}{(2+2t)n}]^{\frac{1}{2+2t}}$ and $\varepsilon_n = \delta_n^t$, based on the preceding discussions we have the following theorem.

Theorem 5.2. Under Assumption A, we have

a) $D_i^*(G, d_n^*) \leq O(\delta_n^t)$ for $i = 1, \ldots, k$, and

b) $r(G, d_n^*) - r(G) \leq O(\delta_n^t)$.

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This paper deals with the problem of selecting good negative binomial populations as compared with a standard or a control. The main results are based on the use of the empirical Bayes approach. First we derive the monotone empirical Bayes estimators of the concerned parameters. Based on these estimators, we construct monotone empirical Bayes selection rules. Asymptotic optimality properties of the monotone empirical Bayes estimators and the monotone empirical Bayes selection rules are investigated. The respective convergence rates for the estimation problem and for the selection problem are studied, under some conditions.

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