To cite this version :
Chafaï, D and Concordet, D (2009) *Confidence regions for the multinomial parameter with small sample size.* Journal of the American Statistical Association, vol. 104 (n° 487). pp. 1071-1079. ISSN 0162-1459

Confidence regions for the multinomial parameter with small sample size

Djalil Chafaï and Didier Concordet

Preprint – March 2008 – Revised July 2008

Abstract

Consider the observation of n iid realizations of an experiment with $d \geq 2$ possible outcomes, which corresponds to a single observation of a multinomial distribution $\mathcal{M}_d(n,p)$ where p is an unknown discrete distribution on $\{1,\ldots,d\}$. In many applications, the construction of a confidence region for p when n is small is crucial. This concrete challenging problem has a long history. It is well known that the confidence regions built from asymptotic statistics do not have good coverage when n is small. On the other hand, most available methods providing non-asymptotic regions with controlled coverage are limited to the binomial case $d = 2$. In the present work, we propose a new method valid for any $d \geq 2$. This method provides confidence regions with controlled coverage and small volume, and consists in the inversion of the "covering collection" associated to level-sets of the likelihood. The behavior when d/n tends to infinity remains an interesting open problem beyond the scope of this work.

Keywords. Confidence regions, small samples, multinomial distribution.

Contents

1 Introduction

Consider the observation of n iid realizations $Y_1, \dots, Y_n$ of an experiment with $d \geq 2$ possible outcomes with common discrete distribution $p_1\delta_1 + \cdots + p_d\delta_d$ on $\{1, \dots, d\}$ (here δ_a denotes the Dirac mass at point a). This corresponds to a single observation $X = (X_1, \dots, X_d)$ of the multinomial distribution

$$\mathcal{M}_d(n, p) = \sum_{\substack{0 \leq k_1, \dots, k_n \leq n \\ k_1 + \cdots + k_d = n}} \mu_p(k)\delta_{(k_1, \dots, k_d)} \quad \text{where} \quad \mu_p(k) = p_1^{k_1} \cdots p_d^{k_d} \frac{n!}{k_1! \cdots k_d!}$$

where $p = (p_1, \dots, p_d)$ and $X_k = \mathrm{Card}\{1 \leq i \leq n \text{ such that } Y_i = k\}$ for every $1 \leq k \leq d$. Here d is known, X is observed, and p is unknown. The present article deals with the problem of constructing a confidence region for p from the single observation X of $\mathcal{M}_d(n, p)$, in the *non-asymptotic* situation where n is *small.* More precisely, let

$$\Lambda_d = \{(u_1, \dots, u_d) \in [0, 1]^d \text{ such that } u_1 + \cdots + u_d = 1\}$$

be the simplex of probability distributions on $\{1, \dots, d\}$. The observation $X \sim \mathcal{M}_d(n, p)$ lies in the discrete simplex

$$E_d = \left\{(x_1, \dots, x_d) \in \{0, \dots, n\}^d \;\text{ such that }\; x_1 + \cdots + x_d = n\right\}. \tag{1}$$

From the single observation X and for some prescribed level $\alpha \in (0,1)$, we are interested in the construction of a random region $R_\alpha(X) \subset \Lambda_d$ depending on X and α such that

- *the coverage probability has a prescribed lower bound*

$$\mathbb{P}(p \in R_\alpha(X)) \geq 1-\alpha \tag{2}$$

- *the volume of $R_\alpha(X)$ in $\mathbb{R}^d$ is as small as possible.*

These two properties are the most important in practice. We propose to solve this problem by defining the "level-set" confidence region $R_\alpha\left(X\right) \subset \Lambda_d$ given by

$$R_\alpha\left(X\right) = \{p \in \Lambda_d \quad \text{such that} \quad \mu_p(X) \geq u(p,\alpha)\} \tag{3}$$

where

$$u(p,\alpha) = \sup\left\{u \in [0,1] \ \text{ such that } \sum_{\substack{k \in E_d \\ \mu_p(k) \geq u}} \mu_p(k) \geq 1-\alpha\right\}.$$

It is immediate to check that this confidence region (3) contains always the maximum likelihood estimator $n^{-1}X$ of p. Moreover, this region can be easily computed numerically (i.e. for each value of p one may compute $u(p,\alpha)$ and compare it to $\mu_p(X)$). Furthermore, it fulfills (2), and the numerical computations presented in Section 3 show that it has small volume and actual coverage often close to $1-\alpha$ at least for $d=2$ and $d=3$. In fact, this region is a special case of a generic method of construction based on *covering collections*. The concept of covering collections is presented in Section 2 and encompasses also as another special case the Clopper-Pearson interval which is classical for the binomial case $d=2$. On the other hand, it is well known (see for instance Remark 2.6) that there exists a natural correspondence via inversion between confidence regions with prescribed coverage and families of tests with prescribed level. However, this correspondence is a simple translation and does not give any clue to construct regions with small volume.

One can find in the literature (see for instance [10, 11, 8, 6, 7, 23] for reviews) two kinds of methods for the construction of confidence region for p. On the first hand, methods that give confidence regions with small volume but that fail to control the prescribed coverage (e.g. Bayesian methods with Jeffrey prior, Wald or score methods based on the Central Limit Theorem, Bootstrapped regions,...), and on the second hand, methods that control the prescribed coverage but have a too large volume to be useful (e.g. concentration methods based on Hoeffding-Bernstein inequalities, Clopper-Pearson type methods, ...). Note that the discrete nature of the multinomial distribution produces a staircase effect which makes difficult the construction of non-asymptotic regions with coverage equal exactly to $1-\alpha$. For a discussion of such aspects, we refer for instance to Agresti et al. [3, 2, 1]. In general, it seems reasonable to ask for a coverage of at least $1-\alpha$, without being too conservative, while maintaining the volume as small as possible. Here the term conservative means that the coverage is greater than $1-\alpha$. It is also well known that even when $d=2$ and n is large but finite, the confidence regions built from the asymptotic approaches based on the Central Limit Theorem have a poor and uncontrolled coverage. For the same reasons, it is also the case for the bootstrapped versions which only improve asymptotically the coverage probability (see [28, 21, 17, 18, 16, 27]). For the binomial case $d=2$, one of the best known method is due to Blyth & Still [8] and mixes various approaches. To our knowledge, the available methods for the general multinomial case $d>2$ are unfortunately asymptotic or Bayesian, which explains their poor performances when n is small in terms of coverage or volume (see [26, 25, 4, 17, 21]).

Recall that the coverage of our region (3) is strictly controlled since it fulfills (2) whatever d and n are. However, this says nothing about the actual coverage and the actual mean volume. The comparisons presented in Section 3 suggests that our region for $d=2$ is comparable to the Blyth & Still region in terms of actual coverage and actual mean volume. For $d=3$, the Blyth & Still method is no longer available, and our region seems to have an actual coverage close to the nominal level while maintaining a volume comparable to the asymptotic region constructed with the score method based on the Central Limit Theorem. Section 3 provides two concrete examples, one for $d=3$ and another one for $d=4$ in relation with the χ^2-test. The article ends up with a final discussion.

2 Covering collections

The aim of this section is to introduce the notion of *covering collection*, which allows to build confidence regions in a general abstract space. Let us consider a random variable $X : (\Omega, \mathcal{A}) \to (E, \mathcal{B}_E)$ having a distribution μ_{θ^*} where $\theta^* \in \Theta$. For some $\alpha \in (0,1)$, we would like to construct a confidence region $R_\alpha(X)$ for θ^* with a coverage of at least $(1-\alpha)$, from a single realization of X. In other words,

$$\mathbb{P}\left(\theta^* \in R_\alpha\left(X\right)\right) \geq 1-\alpha. \tag{4}$$

Definition 2.1 (Covering collection)**.** *A* covering collection *of E is a collection of measurable events $(A_k)_{k\in\mathcal{K}} \subset \mathcal{B}_E$ such that*

- *$\mathcal{K}$ is totally ordered and admits a minimal element and a maximal element;*
- *if $k \leq k'$ then $A_k \subset A_{k'}$ with equality if and only if $k = k'$;*
- *$A_{\min(\mathcal{K})} = \emptyset$ and $A_{\max(\mathcal{K})} = E$.*

For instance, for $E = \{0, 1, \dots, n\}$, the sequence of sets

$$\emptyset, \{\sigma(0)\}, \{\sigma(0), \sigma(1)\}, \dots, \{\sigma(0), \sigma(1), \dots, \sigma(n)\} = E$$

is a covering collection of E for any permutation σ of E. For $E = \mathbb{R}$, the collection $(A_t)_{t\in\overline{\mathbb{R}}}$ where $\overline{\mathbb{R}} = \mathbb{R} \cup \{-\infty, +\infty\}$ defined by $A_{-\infty} = \emptyset$, $A_t = (-\infty, t]$ for every $t \in \mathbb{R}$, and $A_{+\infty} = \mathbb{R}$ is a covering collection of E. Many other choices are possible, like $A_t = [-t, +t]$ or $A_t = [t, +\infty)$. We recognize the usual shapes of the confidence regions used in univariate Statistics.

Theorem 2.2 (Confidence region associated with a covering collection)**.** *Let $(A_k)_{k\in\mathcal{K}}$ be a covering collection of E, and k_X be the smallest $k \in \mathcal{K}$ such that $X \in A_k$. For every $\alpha \in (0,1)$, the region $R_\alpha(X)$ defined below satisfies to* (4)*.*

$$R_\alpha\left(X\right) = \left\{\theta \in \Theta \;\; \textit{such that} \;\; \mu_\theta(A_{k_X}) \geq \alpha\right\}. \tag{5}$$

Proof. For every $\theta \in \Theta$, let $k_\alpha(\theta)$ be the largest $k \in \mathcal{K}$ such that $\mu_\theta(A_k) < \alpha$. With this definition of $k_\alpha(\cdot)$, we have then

$$x \in A_{k_\alpha(\theta)} \text{ if and only if } \mu_\theta(A_{k_x}) < \alpha.$$

Thus we have

$$\begin{aligned}
\mathbb{P}\left(\theta^* \in R_\alpha(X)\right) &= \mathbb{P}\left(\mu_{\theta^*}(A_{k_X}) \geq \alpha\right) \\
&= \mathbb{P}\left(X \notin A_{k_\alpha(\theta^*)}\right) \\
&= 1 - \mu_{\theta^*}\left(A_{k_\alpha(\theta^*)}\right) \\
&\geq 1 - \alpha.
\end{aligned}$$

□

These confidence regions highly depend on the chosen covering collection $(A_k)_{k\in\mathcal{K}}$. Each choice of covering collection gives a particular region $R_\alpha(X)$. One can notice that a small value of k_X gives a small set A_{k_X} and thus leads to a confidence region with a small volume. For instance, assume that we have two realizations x_1 and x_2 of X with $k_{x_1} < k_{x_2}$. For a given sequence $(A_k)_{k\in\mathcal{K}}$, we have $A_{k_{x_1}} \subset A_{k_{x_2}}$ and thus $R_\alpha(x_1) \subset R_\alpha(x_2)$. It is tempting to choose the covering collection $(A_k)_{k\in\mathcal{K}}$ in such a way that k_X is as small as possible. Unfortunately, with such as choice, the covering collection $(A_k)_{k\in\mathcal{K}}$ could be random and the coverage of the associated region could be less than the prescribed level $1-\alpha$.

Note that the set A_{k_X} can be empty, which means that a confidence region cannot be built with such a sequence $(A_k)_{k\in\mathcal{K}}$. In contrast, the case where $A_{k_X} = E$ leads to the trivial region $R_\alpha(X) = \Theta$. In the case where $A_{k_X} = \{X\}$, we have $\mu_\theta(A_{k_X}) = \mu_\theta(\{X\})$, which is the likelihood of X at point θ, and the region $R_\alpha(X)$ corresponds to the complement of a level-set of the likelihood.

The following symmetrization Lemma allows for instance the construction of two-sided confidence intervals from one-sided confidence intervals. We use it in Section 2.2 to interpret the Clopper-Pearson confidence interval as a special case of the covering collection method.

Lemma 2.3 (Symmetrization). *Consider a covering collection $(A_k)_{0\le k\le\kappa}$ of E. For every $0\le k\le\kappa$ let us define $A'_k = E\setminus A_{\kappa-k}$. For any $\theta\in\Theta$, any $X\sim\mu_\theta$, and any $\alpha\in(0,1)$, we construct*

$$R_{\frac{1}{2}\alpha} = \left\{\theta\in\Theta;\ \mu_\theta(A_{k_X}) > \frac{1}{2}\alpha\right\} \quad \text{and} \quad R'_{\frac{1}{2}\alpha} = \left\{\theta\in\Theta;\ \mu_\theta(A'_{k'_X}) > \frac{1}{2}\alpha\right\}$$

where k'_X is built from $(A'_k)_{0\le k\le\kappa}$ as k_X from $(A_k)_{0\le k\le\kappa}$ and $A'_{k'_X} = E\setminus A_{k_X-1}$. Then

$$R_{\frac{1}{2}\alpha}\cap R'_{\frac{1}{2}\alpha}$$

is a confidence region with coverage greater than or equal to $1-\alpha$.

Proof. We have $\mu_\theta(A_{k_X}) + \mu_\theta(A'_{k_X}) = 1+\mu_\theta(\{X\}) \ge 1$ and thus $R_{\frac{1}{2}\alpha}$ and $R'_{\frac{1}{2}\alpha}$ have disjoint complements. The conclusion follows now from a general fact: if R_1 and R_2 are two confidence regions with a coverage of at least $1-\frac{1}{2}\alpha$ such that $R_1\cup R_2 = E$ (equivalently $R_1^c = \Theta\setminus R_1$ and $R_2^c = \Theta\setminus R_2$ are disjoint), then R_1^c and R_2^c are disjoint and thus $R_1\cap R_2 = (R_1^c\cup R_2^c)^c$ is a confidence region with a coverage of at least $1-\alpha$. $\square$

Remark 2.4 (Discrete case and staircase effect). *Let $(A_k)_{k\in\mathcal{K}}$ be a covering collection of a finite set E. Due to staircase effects, the coverage of the confidence regions constructed from this covering collection cannot take arbitrary values in $(0,1)$. These staircase effects can be reduced by using a fully granular collection for which* $\mathrm{Card}(\mathcal{K}) = \mathrm{Card}(E)$. *The term* fully granular *means that the elements of the collection are obtained by adding the points of E one by one. It is impossible to remove completely the staircase effects when E is discrete, while maintaining a nominal lower bound on the coverage.*

Remark 2.5 (Reverse regions). *For the region $R_\alpha(X) = \{\theta\in\Theta; \mu_\theta(A_{k_X})\le 1-\alpha\}$ we have*

$$\mathbb{P}(R_\alpha) = \mathbb{P}(\mu_\theta(A_{k_X})\le 1-\alpha) = \mathbb{P}(X\in A_{k_{1-\alpha}}) = \mu_\theta(A_{k_{1-\alpha}})\le 1-\alpha.$$

Remark 2.6 (Link with tests)**.** *Let us recall briefly the correspondence between confidence regions and statistical tests (we refer to [9, Section 48] for further details). Consider a parametric model* $(\mu_\theta)_{\theta\in\Theta}$ *with data space* $\mathcal{X}$*. For any fixed* $\theta_0 \in \Theta$*, the test problem of* $H_0 : \theta = \theta_0$ *versus* $H_1 : \theta \neq \theta_0$ *with level* $\alpha \in (0, 1)$ *corresponds to the construction of an acceptance region* $C_\alpha(\theta_0) \subset \mathcal{X}$ *such that*

$$\mu_{\theta_0}(C_\alpha(\theta_0)) \geq 1 - \alpha.$$

The construction of a confidence region for θ_0 *can be done by inversion (i.e. by collecting the values of* θ_0 *for which* H_0 *is accepted). Namely, for every* $x \in \mathcal{X}$*, one can define the region* $R_\alpha(x) \subset \Theta$ *by*

$$R_\alpha(x) = \{\theta \in \Theta \text{ such that } x \in C_\alpha(\theta)\}.$$

Now if $X \sim \mu_{\theta_0}$ *then*

$$\mathbb{P}(\theta_0 \in R_\alpha(X)) = \mathbb{P}(X \in C_\alpha(\theta_0)) = \mu_{\theta_0}(C_\alpha(\theta_0)) \geq 1 - \alpha.$$

This shows that for any fixed $\theta_0 \in \Theta$*, the set* $R_\alpha(X) \subset \Theta$ *is a confidence region for* θ_0 *when* $X \sim \mu_{\theta_0}$*. Conversely, if for every* $\theta_0 \in \Theta$ *and every* $x \in \mathcal{X}$ *one has a region* $R_\alpha(x) \subset \Theta$ *such that* $\mathbb{P}(\theta_0 \in R_\alpha(X)) \geq 1 - \alpha$ *when* $X \sim \mu_{\theta_0}$*, then one can construct immediately a test for* $H_0 : \theta = \theta_0$ *versus* $H_1 : \theta \neq \theta_0$ *with acceptance region*

$$C_\alpha(\theta_0) = \{x \in \mathcal{X} \text{ such that } \theta_0 \in R_\alpha(x)\}.$$

Note that this correspondence between confidence regions and statistical tests can be extended to the composite case $H_0 : \theta \in \Theta_0$ *versus* $H_1 : \theta \notin \Theta_0$ *where* $\Theta_0 \subset \Theta$*.*

2.1 The case of the level-sets regions

In this section, we show that the so called "level-sets" confidence region (3) is a special case of the covering collection method. It is easier to consider here a decreasing covering collection (the corresponding version of Theorem 2.2 is immediate). Let us consider a random variable $X : (\Omega, \mathcal{A}) \to (E, \mathcal{B}_E)$ with law μ_{θ^*} where $\theta^* \in \Theta$. For every $u \geq 0$ and $\theta \in \Theta$, let us define

$$A(\theta, u) = \{x \in E \text{ such that } \mu_\theta(x) \geq u\}.$$

For every $\theta \in \Theta$, the collection $(A(\theta, u))_{u \geq 0}$ is decreasing with $A(\theta, 0) = E$ and there exists $u_{\max}$ that can be equal to $+\infty$ such that $A(\theta, u_{\max}) = \emptyset$. Also, $(A(\theta, u_{\max} - u))_{u \in [0, u_{\max}]}$ is a covering collection of E. Next, define

$$u(\theta, \alpha) = \sup \{u \in [0, u_{\max}] \text{ such that } \mu_\theta(A(\theta, u)) \geq 1 - \alpha\}$$

and

$$K(\theta, \alpha) = A(\theta, u(\theta, \alpha)).$$

We would like to construct a confidence region for θ^* form the observation of $X \sim \mu_{\theta^*}$. If

$$R_\alpha (X) = \{\theta \in \Theta \text{ such that } X \in K(\theta, \alpha)\} \tag{6}$$

then

$$\mathbb{P} (\theta^* \in R_\alpha (X)) = \mathbb{P} (X \in K(\theta^*, \alpha)) = \mu_{\theta^*}(K(\theta^*, \alpha)) \geq 1 - \alpha.$$

This shows that $R_\alpha(X)$ is a confidence region for θ^* with a coverage of at least $1 - \alpha$. Let us make precise the expression of the confidence region for the general multinomial case where $X \sim \mathcal{M}_d(n, p)$ with $p \in \Lambda_d$ and $d \geq 2$. Here the value of p used for the observed data X plays the role of θ^*. We have $\Theta = \Lambda_d$, $E = E_d$ as described by (1), $\mu_\theta = \mathcal{M}_d(n, \theta)$, and $u_{\max} = 1$. For every $\alpha \in (0, 1)$, the confidence region given by the level-sets method writes as in (3) given in the introduction.

Optimality

Let us focus on the case where E is a finite set. The confidence region constructed above is not optimal among all the $1 - \alpha$ conservative sets and thus could be improved by a more detailed analysis. Let us first notice that by its very construction, for all $\theta \in \Theta$, $K(\theta, \alpha)$ is minimal with respect to its cardinality that is, there does not exist a set $B(\theta, \alpha)$ so that $\mu_\theta(B(\theta, \alpha)) \geq 1 - \alpha$ and $\mathrm{card}(B(\theta, \alpha)) < \mathrm{card}(K(\theta, \alpha))$. However, in some circumstances, it may exist sets $L(\theta, \alpha)$ with the same cardinality as $K(\theta, \alpha)$ so that $\mu_\theta(K(\theta, \alpha)) \geq \mu_\theta(L(\theta, \alpha)) \geq 1 - \alpha$. The following theorem gives a condition that allows to build conservative sets but with a coverage closer to $1 - \alpha$ than the coverage of $R_\alpha(X)$. For all $\alpha \in [0, 1]$ and $\theta \in \Theta$, let us denote $\gamma(\theta, \alpha) = 1 - \mu_\theta(K(\theta, \alpha))$ and let us notice that $\gamma(\theta, \alpha) \leq \alpha$.

Theorem 2.7. *For each $\theta \in \Theta$, assume that it exist two subsets $V(\theta, \alpha) \subset K(\theta, \alpha)$ and $W(\theta, \alpha) \subset E \backslash K(\theta, \alpha)$ with the same cardinality so that*

$$\alpha - \gamma(\theta, \alpha) \geq \mu_\theta(V(\theta, \alpha)) - \mu_\theta(W(\theta, \alpha)) > 0.$$

Then, there exists a set $T_\alpha(X) \neq R_\alpha(X)$ so that

$$1 - \alpha \leq \mathbb{P}(\theta^* \in T_\alpha(X)) < \mathbb{P}(\theta^* \in R_\alpha(X)).$$

Proof. Let us consider the set $L(\theta, \alpha) = K(\theta, \alpha) \backslash V(\theta, \alpha) \bigcup W(\theta, \alpha)$ and notice that thanks to the conditions imposed the sets V and W we have for all $\theta \in \Theta$,

$$1 - \alpha \leq \mu_\theta(L(\theta, \alpha)) < \mu_\theta(K(\theta, \alpha)).$$

Now, set

$$T_\alpha(X) = \{\theta \in \Theta; X \in L(\theta, \alpha)\}.$$

But,

$$\begin{aligned}
\mathbb{P}(\theta^* \in T_\alpha(X)) &= \mathbb{P}(X \in L(\theta^*, \alpha)) \\
&= \mathbb{P}(X \in K(\theta^*, \alpha) \setminus V(\theta^*, \alpha) \textstyle\bigcup W(\theta^*, \alpha)) \\
&= 1 - \gamma(\theta^*, \alpha) - \mu_{\theta^*}(V(\theta^*, \alpha)) + \mu_{\theta^*}(W(\theta^*, \alpha)) \\
&\leq 1 - \gamma(\theta^*, \alpha).
\end{aligned}$$

On the other hand, we have already seen that for all $\theta \in \Theta$,

$$1 - \alpha \leq \mu_\theta \left(L(\theta, \alpha) \right).$$

This last inequality holds true when $\theta = \theta^*$ and thus

$$1 - \alpha \leq \mu_{\theta^*} \left(L(\theta^*, \alpha) \right) = \mathbb{P} \left(\theta^* \in T_\alpha \left(X \right) \right).$$

$\square$

This theorem can be used to build less conservative confidence sets than $R_\alpha(X)$. A convenient way to proceed is to take $V(\theta, \alpha) = \{y\}$ where y is such that

$$\mu_\theta(y) = \min_{z \in K(\theta,\alpha)} \mu_\theta(z)$$

and to iteratively try several sets W^k as follows. Set $W^0(\theta, \alpha) = \emptyset$, and at iteration $k \geq 1$, set $W^k(\theta, \alpha) = \{w_k\}$ and $L^k(\theta, \alpha) = K\left(\theta, \alpha\right) \setminus V\left(\theta, \alpha\right) \bigcup W^k\left(\theta, \alpha\right)$ where

$$w_k = \arg \max_{z \in L^{k-1}(\theta,\alpha)} \mu_\theta(z).$$

This process is iterated until the set $L^k(\theta, \alpha)$ is such that $\mu_\theta \left(L^k(\theta, \alpha) \right) - (1 - \alpha)$ is non-negative and minimum.

Since for $\theta \in \Theta$ there may exist $x \neq y$ with $\mu_\theta(x) = \mu_\theta(y)$, there also may exist several sets $(L^i(\theta, \alpha))_i$ which have the same mass $\mu_\theta(L^i(\theta, \alpha)) = 1 - \delta(\theta, \alpha)$. Several confidence sets with the same coverage can thus be derived using these sets. A simple way to choose between these concurrent confidence sets is to adopt the one that optimizes a criterion such as having a minimum volume (for the Lebesgue measure).

2.2 The case of the Clopper-Pearson regions

Consider the binomial case $d = 2$ for which $p = (p_1, 1-p_1)$. The well known Clopper-Pearson interval for p_1 relies on the exact distribution of X_1 in the binomial case [14, 20, 13]. It was considered for a long time as outstanding. This interval $[L, U]$ is given by

$$\begin{cases} L & = \inf\left\{\theta \in [0,1] \text{ such that } \sum_{i=x_1}^{n} \binom{n}{i}\theta^i(1-\theta)^{n-i} \geq \frac{1}{2}\alpha\right\} \\ U & = \sup\left\{\theta \in [0,1] \text{ such that} \sum_{i=0}^{x_1} \binom{n}{i}\theta^i(1-\theta)^{n-i} \geq \frac{1}{2}\alpha\right\}. \end{cases} \tag{7}$$

It has been shown that the Clopper-Pearson interval is often conservative. Also, some continuity corrections have been proposed, and give the so called "mid-p interval", see for instance [5] for a review. This trick reduces the staircase effect but the coverage probability can be less than $1 - \alpha$. The Beta-Binomial correspondence (see lemma 2.8 below) shows that the left and right limits L and R of the Clopper-Pearson confidence interval (7) are the $\frac{1}{2}\alpha$ and $(1 - \frac{1}{2}\alpha)$ quantiles of the Beta distribution $\text{Beta}\,(X_1; n - X_1 + 1)$.

Lemma 2.8 (Beta-Binomial correspondence)**.** *If* $X \sim \text{Binom}(n, p_1)$ *with* $p_1 \in [0, 1]$ *and* $0 \leq k \leq n$ *and* $B \sim \text{Beta}(k, n - k + 1)$ *then following identity holds true.*

$$\mathbb{P}(X \geq k) = \mathbb{P}(B \leq p_1). \tag{8}$$

Proof. We briefly recall here the classical proof (see [9, page 68]). Let $U_1, \ldots, U_n$ be iid uniform random variables on $[0, 1]$ and $U_{(1)} \leq \cdots \leq U_{(n)}$ be the reordered sequence. If we define $V_{p_1} = \sum_{i=1}^{n} \mathrm{I}_{\{U_i \leq p_1\}}$ then $V_{p_1} \sim \text{Bin}(n, p_1)$ and $U_{(k)} \sim \text{Beta}(k, n-k+1)$ and for every $1 \leq k \leq n$, $V_{p_1} \geq k$ if and only if $U_{(k)} \leq p_1$. □

The confidence interval obtained by the level-sets method does not coincide with the classical Clopper-Pearson confidence interval. Let us show why the Clopper-Pearson confidence interval can be considered as a special case of the method based on covering collections. Recall that we are in the case where $d = 2$ and $X_1 \sim \text{Binom}(n, p_1)$ for some unknown

$p_1 \in [0,1]$. Equivalently, we can write

$$(X_1, n - X_1) \sim \mathcal{M}_2(n, (p_1, 1 - p_1)).$$

The unidimensional nature of $E = \{0, \ldots, n\}$ suggests the following two covering collections $(A_k^1)_{k\in E}$ and $(A_k^2)_{k\in E}$ defined by $A_0^1 = \emptyset$ and $A_0^2 = \emptyset$, and for every $0 \le k \le n$,

$$A_{k+1}^1 = \{0, \ldots, k\} \quad \text{and} \quad A_{k+1}^2 = \{n-k, \ldots, n\}.$$

Here $\mathcal{K} = E$ for both the top-to-bottom and bottom-to-top sequences. The bottom-to-top sequence $(A_k^1)_{k\in E}$ leads to the $(1-\alpha)$ one-sided confidence interval for p_1 given by

$$R_\alpha^1(X_1) = \left\{ \theta \in [0,1] \text{ such that} \sum_{i=0}^{X_1} \binom{n}{i} \theta^i (1-\theta)^{n-i} \ge \alpha \right\} = [0, U_\alpha(X_1)] \tag{9}$$

where

$$U_\alpha(x) = \sup \left\{ \theta \in [0,1] \text{ such that} \sum_{i=0}^{x} \binom{n}{i} \theta^i (1-\theta)^{n-i} \ge \alpha \right\}.$$

On the other hand, the top-to-bottom covering collection $(A_k^2)_{k\in E}$ leads to an $(1-\alpha)$ confidence interval of p_1 given by

$$R_\alpha^2(X_1) = \left\{ \theta \in [0,1] \text{ such that} \sum_{i=X_1}^{n} \binom{n}{i} \theta^i (1-\theta)^{n-i} \ge \alpha \right\} = [L_\alpha(X_1); 1] \tag{10}$$

where

$$L_\alpha(x) = \sup \left\{ \theta \in [0,1] \text{ such that} \sum_{i=x}^{n} \binom{n}{i} \theta^i (1-\theta)^{n-i} \ge \alpha \right\}.$$

By virtue of Lemma 2.3, we can combine the one-sided confidence intervals (9) and (10) in order to obtain a two-sided $(1-\alpha)$ confidence interval of p_1, which is the two-sided interval

$$R_{\frac{1}{2}\alpha}^1(X_1) \bigcap R_{\frac{1}{2}\alpha}^2(X_1) = [L_{\frac{1}{2}\alpha}(X_1); U_{\frac{1}{2}\alpha}(X_1)].$$

We recognize the Clopper-Pearson interval (7). The discrete nature of E precludes the construction of a confidence interval of p_1 with coverage exactly equal to $1-\alpha$. Actually, the Clopper-Pearson interval is not exactly symmetric and there is no guaranty that

$$\mathbb{P}\left(p < L_{\frac{1}{2}\alpha}(X_1)\right) = \mathbb{P}\left(p > U_{\frac{1}{2}\alpha}(X_1)\right).$$

Our construction via a covering collection provides immediately an extension of the Clopper-Pearson interval in the general multinomial case where $X \sim \mathcal{M}_d(n,p)$ with $p \in \Lambda_d$ and $d > 2$. This construction consists in labeling the elements of E_d (note that $\mathrm{Card}(E_d) = \binom{n+d-1}{d-1}$) and constructing the covering collection $(A_k)_{k\in\mathcal{K}}$ which grows by adding the points one after the other. The choice of the total order on E_d is arbitrary when $d > 2$. Some additional constraints can help to reduce this choice. As advocated by Casella [12] for the binomial distribution, the proposed confidence region $R_\alpha(X)$ should be *equivariant*, that is not sensitive to the order chosen to label the d categories of the multinomial distribution.

Definition 2.9 (Equivariance)**.** *A confidence region $R_\alpha(X)$ is* equivariant *when*

$$\mathbb{P}\left(\sigma(\theta^*) \in R_\alpha\left(\sigma(X)\right)\right) = \mathbb{P}\left(\theta^* \in R_\alpha\left(X\right)\right) \tag{11}$$

for every permutation σ of $\{1,\ldots,d\}$. In other words, if and only if

$$\sigma\left(R_\alpha\left(X\right)\right) = R_\alpha\left(\sigma(X)\right).$$

The following lemma gives a criterion of equivariance for covering collections.

Theorem 2.10 (Equivariance criterion for covering collections)**.** *The confidence region $R_\alpha(X)$ constructed from a covering collection $(A_k)_{k\in\mathcal{K}}$ is equivariant if and only if A_k is invariant by permutation of coordinates for every $k \in \mathcal{K}$.*

Proof. Let σ be a permutation of $\{1,\ldots,d\}$, $i = (i_1,\ldots,i_d) \in E$, and for every $\theta \in \Theta$,

$$\sigma(\theta) = \left(\theta_{\sigma(1)},\ldots,\theta_{\sigma(d)}\right) \quad \text{and} \quad \sigma(i) = \left(i_{\sigma(1)},\ldots,i_{\sigma(d)}\right).$$

By invariance of A_k by permutation, we have $X \in A_k \Leftrightarrow X \in \sigma(A_k)$ and thus $k_X = k_{\sigma(X)}$. If $\theta \in \sigma\left(R_\alpha\left(X\right)\right)$ then $\mu_{\sigma^{-1}(\theta)}(A_{k_X}) \geq \alpha$. But, for every $i \in E$,

$$\mu_{\sigma^{-1}\theta)}(\{i\}) = \mu_\theta(\{\sigma(i)\}).$$

If A_k is invariant permutations, then for every $i \in A_k$, we have $\sigma(i) \in A_k$ and consequently

$$\mu_{\sigma^{-1}(\theta)}(A_k) = \mu_\theta(\sigma(A_k)) = \mu_\theta(A_k).$$

Thus, $\theta \in \sigma\left(R_\alpha\left(X\right)\right)$ if and only if $\mu_\theta(A_{k_X}) = \mu_\theta(A_{k_{\sigma(X)}}) \geq \alpha$, that is $\theta \in R_\alpha\left(\sigma(X)\right)$. □

Equivariance is a strong constraint on the covering collection. A large set A_{k_X} gives a large confidence region. Since confidence regions with small volume are desirable, it is interesting, when E is discrete, to consider a covering collection $(A_k)_{k\in\mathcal{K}}$ which grows by adding the points of E one after the other. Unfortunately, this method of construction is not compatible with equivariance: the A_k cannot be invariant by permutations of coordinates. A weaker condition consists in the existence of a subsequence $(A_{k_l})_l$ that is invariant by permutation of coordinates. An example of such a sequence for $d = 3$ is given by Figure 1.

Recall that when $d = 2$, the Beta-Binomial correspondence stated in Lemma 2.8 provides a clear link between the quantiles of the Beta distribution and the Clopper-Peason confidence interval. In fact, this can be seen as a special case of the Dirichlet-Multinomial correspondence valid for any $d \geq 3$ as stated in the following Lemma. This makes a link between Clopper-Pearson regions and Bayesian regions constructed with a Jeffrey prior (see for instance [24]). However, the notion of coverage that we use in the present article is purely frequentist and does not fit with the Bayesian paradigm without serious distortions.

Lemma 2.11 (Dirichlet-Multinomial correspondence)**.** *Let $p \in \Lambda_d$ and $k_0, k_1, \ldots, k_d$ such that $k_0 = 0 \leq k_1 \leq \cdots \leq k_{d-1} \leq n \leq k_d = n+1$. If*

$$X \sim \mathcal{M}_d(n,p) \quad \textit{and} \quad D \sim \mathrm{Dirichlet}_d(k_1 - k_0, k_2 - k_1, \ldots, k_d - k_{d-1})$$

then the following identity holds true.

$$\mathbb{P}(X_1 \geq k_1, X_1 + X_2 \geq k_2, \ldots, X_1 + \cdots + X_{d-1} \geq k_{d-1})$$
$$= \mathbb{P}(D_1 \leq p_1, D_1 + D_2 \leq p_2, \ldots, D_1 + \cdots + D_{d-1} \leq p_{d-1}). \quad (12)$$

Proof. The proof is a straightforward extension of the Beta-Binomial case given by Lemma 2.8. Let $I_1, \ldots, I_d$ be the sequence of adjacent sub-intervals of $[0,1]$ of respective lengths $p_1, \ldots, p_d$, $U_1, \ldots, U_n$ be iid uniform random variables on $[0,1]$ and $U_{(1)} \leq \cdots \leq U_{(n)}$ be the reordered sequence. For any $1 \leq r \leq d$, let us define

$$V_{p,r} = \sum_{i=1}^{n} \mathrm{I}_{\{U_i \in \mathrm{I}_r\}} = \mathrm{Card}\{1 \leq i \leq n \ \text{ such that } U_i \in \mathrm{I}_r\}.$$

We have $V_p = (V_{p,1}, \ldots, V_{p,r}) \sim \mathcal{M}_d(n,p)$. Now, for every $0 \leq k_1 \leq \cdots \leq k_{d-1} \leq n$,

$$V_{p,1} \geq k_1, \ldots, V_{p,1} + \cdots + V_{p,d-1} \geq k_{d-1} \quad \text{iff} \quad U_{(k_1)} \leq p_1, \ldots, U_{(k_{d-1})} \leq p_1 + \cdots + p_{d-1}.$$

But by using the notation $U_{(0)} = 0$ and $U_{(n+1)} = 1$, we have

$$(U_{(1)} - U_{(0)}, \ldots, U_{(n+1)} - U_{(n)}) \sim \mathrm{Dirichlet}_{n+1}(1, \ldots, 1).$$

and therefore, by the stability of Dirichlet laws by sum of blocs, with $k_0 = 0$ and $k_d = n+1$,

$$(U_{(k_1)} - U_{(k_0)}, \ldots, U_{(k_d)} - U_{(k_{d-1})}) \sim \mathrm{Dirichlet}_d(k_1, k_2 - k_1, \ldots, k_d - k_{d-1}).$$

□

3 Comparisons and examples

Recall that for every fixed $d \geq 2$, $n \geq 0$, and $p \in \Lambda_d$, a confidence region obtained from $X \sim \mathcal{M}(n,p)$ provides a single coverage probability and a distribution of volumes. In this section, we use coverage probabilities and mean volumes in order to compare the performance of our level-set method to other methods, in the case where $d \in \{2, 3\}$ and $n \in \{5, 10, 20, 30\}$. We also give two concrete examples, one for $d = 3$ and another one for $d = 4$ in relation with the χ^2-test. It turns out that the regions obtained by the Clopper-Pearson method and its multinomial extension have non-competitive volumes and we thus decided to ignore them in the comparisons.

3.1 Performances in the binomial case ($d = 2$)

In the binomial case $d = 2$, a confidence regions for $p = (p_1, 1 - p_1)$ is actually a confidence interval for p_1. It is well known that the Wald interval build from the Central Limit Theorem has poor coverage even when n is large but finite [10]. It is also widely accepted that the score interval or the Blyth-Still interval [6] should be preferred over the Wald interval. We thus compared the performances of the 95%-intervals provided by the level-sets method, the score method, and the Blyth-Still method. We computed the coverages and the mean widths of the intervals obtained with each method for $n \in \{5, 10, 20, 30\}$ and for all $p \in [0; 0.5]$. The results are respectively represented in figures 2 and 3. We can see that for some values of p, the coverage of the score method is smaller than the prescribed level 0.95, while the coverage of the Blyth-Still interval and the level-set interval are always greater than or equal to the nominal level 0.95. The coverages obtained with the level-set method are always closer to the prescribed level except for $n = 20$, $p \in [0.45, 0.48]$ and $n = 30$, $p \in [0.38, 0.42]$. The differences between the coverages of these three methods decrease with n.

Figures 2 and 3 show that the score method provides intervals with excellent mean width but fails to control the coverage. The level-set method gives intervals that have a mean width a bit narrower than the one obtained with the Blyth-Still method. This suggests that the level-set method is an excellent alternative to the Blyth-Still method. Moreover, and in contrast with the Blyth-Still method, the level-set method remains available when $d > 2$.

3.2 Performances in the trinomial case ($d = 3$)

To our knowledge, the Blyth-Still method has no counterpart for $d > 2$. Additionally, the regions obtained by the extended Clopper-Pearson method have non-competitive volumes. Also, we decided to compare the level-set method with the score method. We computed for $d = 3$ the coverage probabilities and the mean volumes of the 95%-regions obtained with both methods, for $n \in \{5, 10, 20\}$. Note that for the score method, only the trace over Λ_3 of the regions is taken into account for the computation of the volume. The graphics of figure 4 show the coverage of both methods as well as the difference of their mean volumes. Whatever the sample size is, the coverage of the level-set regions are very close to $1 - \alpha = 0.95$. On the contrary, the coverages of the score regions can be quite smaller than 0.95. Surprisingly and in contrast with the binomial case $d = 2$, the level-set method here provides confidence regions with comparable (for $n = 5$) or smaller mean volumes than their score's counterparts! We believe that this is due to the fact that we measure the performance by the mean volume. Anyway, the level-set method appears to be a reasonable way to build small confidence sets.

3.3 Concrete example in the trinomial case ($d = 3$)

The present example concerns antibiotics efficacy. A traditional way to evaluate whether or not one can use an antibiotic for a specific pathogen is to perform a susceptibility testing. In such an experiment, different isolates of a given pathogen are classified as Sensible, Intermediate or Resistant according to the antibiotics ability to stop their growth. Here, ten different isolates of Escherichia coli have been tested with ampicillin. The following results have been obtained : 8 isolates were Sensible, 2 Intermediate and 0 were Resistant. The count $x = (8, 2, 0)$ can be seen as the realization of $X \sim \mathcal{M}(10, p)$ where $p = (p_1, p_2, p_3)$ denotes the probabilities that an isolate belong to the different classes. We calculated a 95%-confidence region of p using the level-set method (figure 5). This region suggests that even if no resistant has been observed upon the 10 tested isolates, up to 30% of resistant and 20% of intermediate isolates are yet possible. This confidence region does not contain the situation where all the isolates are sensible and it is thus unlikely that this antibiotics works all the time when it meets this pathogen.

3.4 Concrete example in the quadrinomial case ($d = 4$)

The present example is simply a χ^2-test for independence. It deals with the difference of behavior of male and female veterinary students with respect to smoking habits. The following result has been observed in a group of 12 veterinary students in Toulouse:

	Smokers	Non-smokers
Female	3	8
Male	10	5

The χ^2-test rejects independence with P-value $P = 0.047$ and suggests that more males than females smoke. This P-value is close to the critical threshold 0.05 and has been obtained with a small sample size. Also, one can ask whether this result can be trusted. A possible solution is to build a confidence region. The table above can be seen as the realization $x = (3, 8, 10, 5)$ of a multinomial random variable $X \sim \mathcal{M}(26, p)$ with $p = (p_1, p_2, p_3, p_4)$. If the smoking habit and the gender are independent then p belongs to

$$H_0 = \left\{q \in \Lambda_4 \text{ such that } q = (uv, (1-u)v, u(1-v), (1-u)(1-v)) \text{ and } (u, v) \in [0,1]^2\right\}.$$

Since $p_4 = 1 - p_1 - p_2 - p_3 - p_4$, one can draw a graphic with only p_1, p_2, p_3. Figure 6 shows (in green) the 95% confidence region for p built with the level-set method. The surface corresponds to the null hypothesis H_0. The red area is the acceptance region of the χ^2-test. It turns out that $\hat{p} = (3/26, 8/26, 10/26)$ does not belong to the acceptance region of the χ^2-test. However, the 95%-region for p cuts H_0. Therefore, according to Remark 2.6 and in contrast with the result given by the χ^2-test, the independence hypothesis is not rejected.

4 Final discussion

The general concept of "covering collection" allows to construct confidence regions with controlled coverage, including the classical Clopper-Pearson interval for the binomial and its multinomial extensions. The covering collection construction involves an arbitrary growing collection of sets in the data space. Our “level-set” confidence regions are obtained by using a special collection based on level-sets of the data distribution. The level-set regions for the multinomial parameter can be easily computed for any d and n. It turns out that they have excellent coverage probabilities and mean volumes for $d \in \{2, 3\}$ and $n \leq 30$. They are competitive with the Blyth-Still intervals for $d = 2$. Also, we recommend the level-set regions for these ranges. Nevertheless, the level-set method can be computational expensive when d is very large. The behavior of these confidence regions when the ratio d/n tends to infinity is a very interesting open problem. In this extremal case, the observation X is sparse and belongs to the boundary of the observation simplex E_∞. Note that the critical n for which $X \sim \mathcal{M}(n, p)$ belongs to the interior of E_d corresponds to the classical “coupon collector problem” [15, 22, 19]. Another interesting open problem is the optimality of the level-set regions related to the control of $\mathbb{P}(p' \in R_\alpha(X))$ with $X \sim \mathcal{M}(n, p)$ and $p \neq p'$.

Acknowledgements

The present version of this article greatly benefited from the comments and criticism of an Associate Editor and three anonymous referees.

References

[1] A. Agresti. Dealing with discreteness: making ‘exact’ confidence intervals for proportions, differences of proportions, and odds ratios more exact. *Stat. Methods Med. Res.*, 12(1):3–21, 2003.

[2] A. Agresti and B. Caffo. Simple and effective confidence intervals for proportions and differences of proportions result from adding two successes and two failures. *Amer. Statist.*, 54(4):280–288, 2000.

[3] A. Agresti and B. A. Coull. Approximate is better than "exact" for interval estimation of binomial proportions. *Amer. Statist.*, 52(2):119–126, 1998.

[4] J. Albert. Pseudo-Bayes estimation of multinomial proportions. *Comm. Statist. A—Theory Methods*, 10(16):1587–1611, 1981.

[5] G. Berry and P. Armitage. Mid-p confidence intervals: a brief review. *The Statistician*, 44(4):417–423, 1995.

[6] C. R. Blyth. Approximate binomial confidence limits. *J. Amer. Statist. Assoc.*, 81(395):843–855, 1986.

[7] C. R. Blyth. Correction: "Approximate binomial confidence limits" [J. Amer. Statist. Assoc. **81** (1986), no. 395, 843–855]. *J. Amer. Statist. Assoc.*, 84(406):636, 1989.

[8] C. R. Blyth and H. A. Still. Binomial confidence intervals. *J. Amer. Statist. Assoc.*, 78(381):108–116, 1983.

[9] A. A. Borovkov. *Mathematical statistics*. Gordon and Breach Science Publishers, Amsterdam, 1998. Translated from the Russian by A. Moullagaliev and revised by the author.

[10] L. D. Brown, T. T. Cai, and A. DasGupta. Interval estimation for a binomial proportion. *Statist. Sci.*, 16(2):101–133, 2001. With comments and a rejoinder by the authors.

[11] L. D. Brown, T. T. Cai, and A. DasGupta. Confidence intervals for a binomial proportion and asymptotic expansions. *Ann. Statist.*, 30(1):160–201, 2002.

[12] G. Casella. Refining binomial confidence intervals. *Canad. J. Statist.*, 14(2):113–129, 1986.

[13] X. Chen, K. Zhou, and J. L. Aravena. On the binomial confidence interval and probabilistic robust control. *Automatica J. IFAC*, 40(10):1787–1789 (2005), 2004.

[14] C. J. Clopper and E. S. Pearson. The use of confidence or fiducial limits illustrated in the case of the binomial. *Biometrika*, 26:404–413, 1934.

[15] W. Feller. *An introduction to probability theory and its applications. Vol. I.* Third edition. John Wiley & Sons Inc., New York, 1968.

[16] M. García-Pérez. Exact finite-sample signifiance and confidence regions for goodness-of-fit statistics in one-way multinomials. *British Journal of Mathematical and Statistical Psychology*, 53:193–207, 2000.

[17] J. Glaz and C. P. Sison. Simultaneous confidence intervals for multinomial proportions. *J. Statist. Plann. Inference*, 82(1-2):251–262, 1999. Multiple comparisons (Tel Aviv, 1996).

[18] P. Hall. *The bootstrap and Edgeworth expansion.* Springer Series in Statistics. Springer-Verlag, New York, 1992.

[19] L. Holst. On birthday, collectors', occupancy and other classical urn problems. *Internat. Statist. Rev.*, 54(1):15–27, 1986.

[20] S. A. Julious. Two-sided confidence intervals for the single proportion: comparison of seven methods, letter to the editor. *Statist. Med.*, 24:3383–3384, 2005.

[21] D. Morales, L. Pardo, and L. Santamaría. Bootstrap confidence regions in multinomial sampling. *Appl. Math. Comput.*, 155(2):295–315, 2004.

[22] R. Motwani and P. Raghavan. *Randomized algorithms.* Cambridge University Press, Cambridge, 1995.

[23] R. G. Newcombe. Two-sided confidence intervals for the single proportion: comparison of seven methods. *Statist. Med.*, 17:857–872, 1998.

[24] C. P. Robert. *The Bayesian choice.* Springer Texts in Statistics. Springer-Verlag, New York, second edition, 2001. From decision-theoretic foundations to computational implementation, Translated and revised from the French original by the author.

[25] C. P. Sison and J. Glaz. Simultaneous confidence intervals and sample size determination for multinomial proportions. *J. Amer. Statist. Assoc.*, 90(429):366–369, 1995.

[26] A. Ullah, A. T. K. Wan, and A. Chaturvedi, editors. *Handbook of applied econometrics and statistical inference*, volume 165 of *Statistics: Textbooks and Monographs*. Marcel Dekker Inc., New York, 2002.

[27] E. B. Wilson. Probable inference, the law of succession, and statistical inference. *J. Amer. Statist.*, 22:209–212, 1927.

[28] X.-D. Zheng and W.-Y. Loh. Bootstrapping binomial confidence intervals. *J. Statist. Plann. Inference*, 43(3):355–380, 1995.

Djalil Chafaï, corresponding author, d.chafai[@]envt.fr

UMR181 INRA, ENVT, École Nationale Vétérinaire de Toulouse

23 Chemin des Capelles, F-31076 Cedex 3, Toulouse, France.

UMR 5219 CNRS, Institut de Mathématiques, Université de Toulouse

118 route de Narbonne, F-31062 Cedex 4, Toulouse, France.

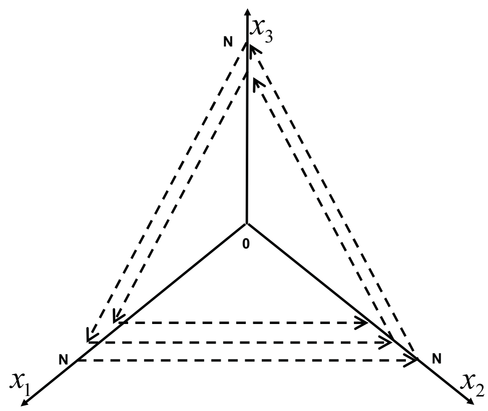

Figure 1: The construction of A_k when $d = 3$, with $A_0 = \emptyset$ and $A_1 = \{(n, 0, 0)\}$. The point in A_1 is at the beginning of the starting arrow represented in dotted line. Each time the arrow meets a point in the simplex, this point is added to A_k to give A_{k+1}. The set obtained with the three first arrows is invariant by permutation of coordinates.

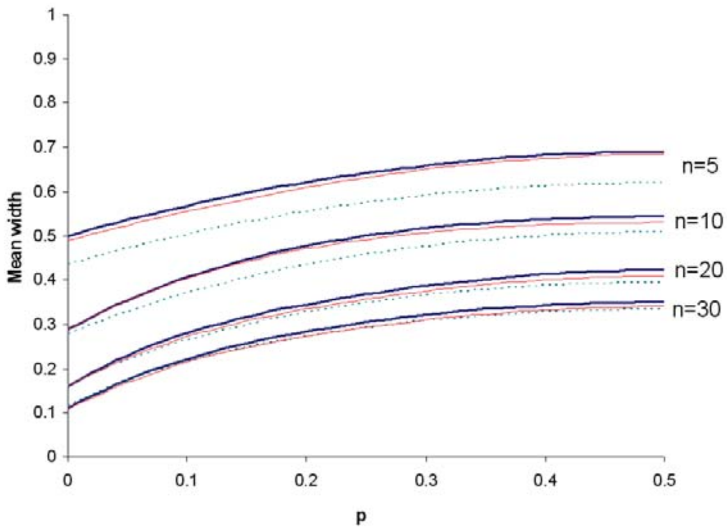

Figure 2: Binomial case $d = 2$. The curves are the mean width of the 95%-intervals obtained with the Blyth-Still method (thick line), the level-set method (thin line) and the score method (dotted line) for $p \in [0, 0.5]$. The Blyth-Still method gives intervals with higher mean width whatever p. The score method always gives intervals with smaller width. Note that the score method fails to control the coverage probability. When n increases the differences between the mean width of the respective intervals decrease.

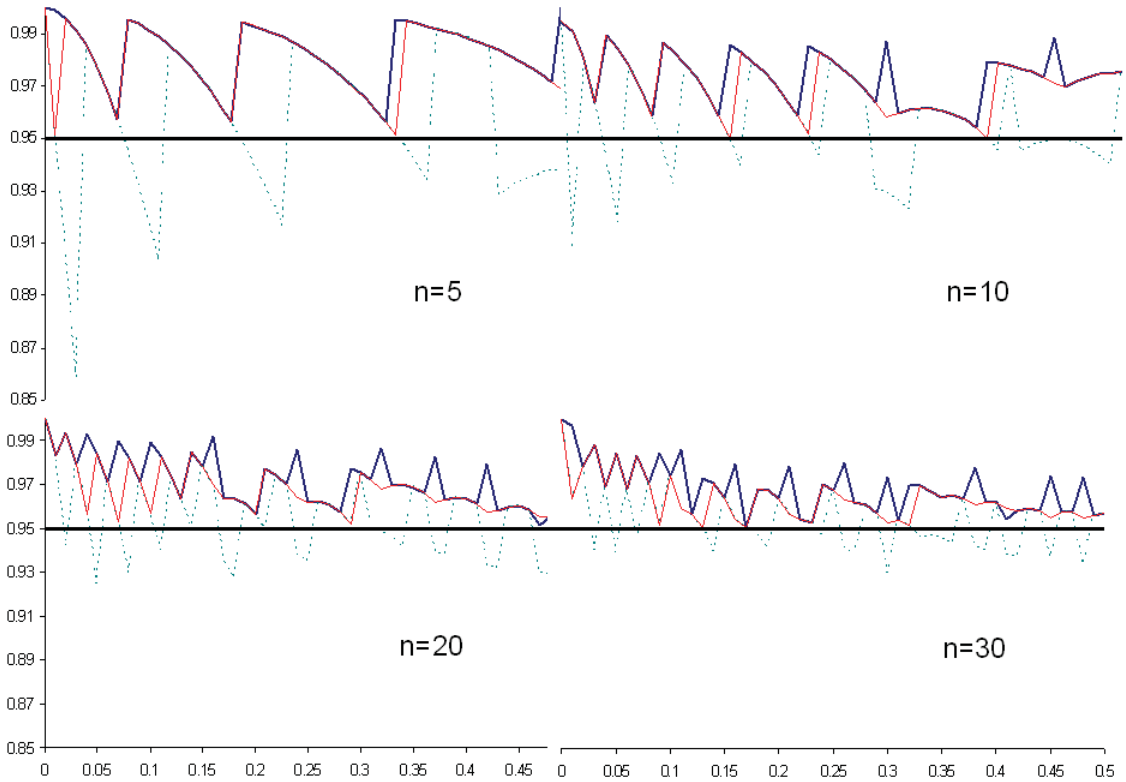

Figure 3: Binomial case $d = 2$. These curves are the coverage of the 95%-intervals obtained with the Blyth-Still method (thick line), the level-set method (thin line) and the score method (dotted line) for $p \in [0, 0.5]$. The score method fails to control the coverage. The level-set method seems (nearly) uniformly better than the Blyth-Still method: its coverages are closer to 0.95. When n increases the differences between these three methods decrease.

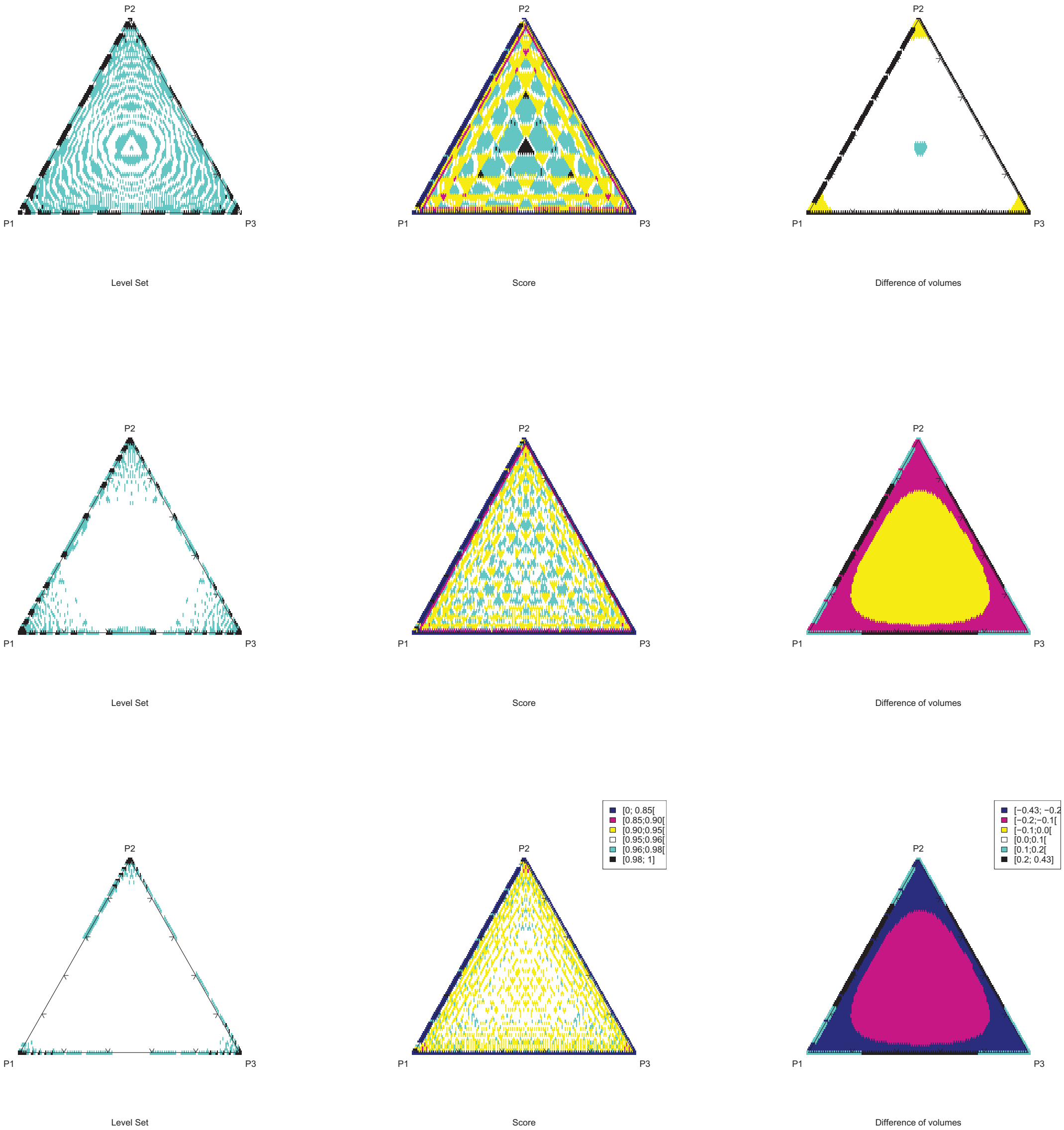

Figure 4: Trinomial case $d = 3$. The columns give the coverages of the level-set method, the coverages given by the score method and the difference of mean volumes. The three rows corresponds to $n \in \{5, 10, 20\}$. For the coverages graphs (first two columns), a clear color means that the coverage is close to 0.95 whereas a dark blue color means that the coverage is smaller than 0.85. For the volumes graphs (third column), a white color means that the difference of mean volumes is small whereas the blue, pink and yellow colors are used when the mean volume of the regions obtained with the level-set method are smaller than their counterpart obtained with the score method.

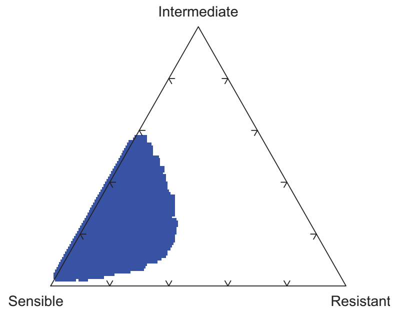

Figure 5: Trinomial case $d = 3$ (example 3.3). In barycentric coordinates, the 95%-region for p constructed from the observation $x = (0, 2, 8)$ of $\mathcal{M}_3(10, p)$. Note that the Wald method cannot be used here since the observation belongs to the boundary of the observation simplex E_3. In this example, the score and the level-set methods approximately give the same region.

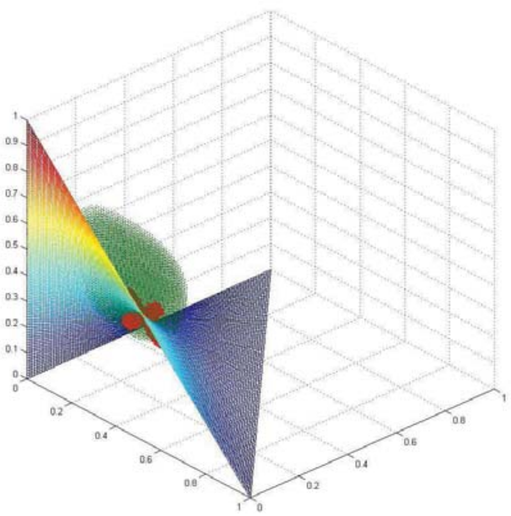

Figure 6: Quadrinomial case $d = 4$ (example 3.4). The axes correspond to p_1, p_2, and p_3. The null hypothesis H_0 of the χ^2-test is represented by the surface. The set in red is the acceptance region of the χ^2-test. The region in green is the 95%-region for p built with the level-set method. It turns out that $\hat{p} = (3/26, 8/26, 10/26)$does not belong to the acceptance region of the χ^2-test while it belongs to the 95%-region for p built with the level-set method. Additionally, since this confidence region cuts H_0, the corresponding test does not reject H_0, in contrast with the χ^2-test.