Subgaussian random variables, Hoeffding's inequality, and Cramér's large deviation theorem

Jordan Bell

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1 Subgaussian random variables

For a random variable X, let $\Lambda_X(t) = \log E(e^{tX})$, the **cumulant generating function of** X. A b**-subgaussian random variable**, $b > 0$, is a random variable X such that

$$\Lambda_X(t) \leq \frac{b^2t^2}{2}, \qquad t \in \mathbb{R}.$$

We remark that for γ_{a,σ^2} a **Gaussian measure**, whose density with respect to Lebesgue measure on $\mathbb{R}$ is

$$p(x, a, \sigma^2) = \frac{1}{\sqrt{2\pi\sigma^2}} e^{-\frac{(x-a)^2}{2\sigma^2}},$$

we have

$$\int_{\mathbb{R}} e^{tx} d\gamma_{0,b^2}(x) = \int_{\mathbb{R}} e^{bty} \frac{1}{\sqrt{2\pi}} e^{-\frac{y^2}{2}} dy = \int_{\mathbb{R}} e^{\frac{b^2t^2}{2}} \frac{1}{\sqrt{2\pi}} e^{-\frac{(y-bt)^2}{2}} dy = e^{\frac{b^2t^2}{2}}.$$

We prove that a b-subgaussian random variable is centered and has variance $\leq b^2$.[1]

Theorem 1. If X is b-subgaussian then $E(X) = 0$ and $\mathrm{Var}(X) \leq b^2$.

Proof. For each $\omega \in \Omega$, $\sum_{k=0}^n \frac{t^k X(\omega)^k}{k!} \to e^{tX(\omega)}$, and by the dominated convergence theorem,

$$\sum_{k=0}^n \frac{t^k E(X)^k}{k!} \to E(e^{tX}) \leq e^{\frac{b^2t^2}{2}} = \sum_{k=0}^\infty \left(\frac{b^2t^2}{2}\right)^k \frac{1}{k!}.$$

Therefore

$$1 + tE(X) + t^2E(X^2) + O(t^3) \leq 1 + \frac{b^2t^2}{2} + O(t^4),$$

[1] Karl R. Stromberg, *Probability for Analysts*, p. 293, Proposition 9.8.

whence

$$tE(X) + t^2 E(X^2) \leq \frac{b^2 t^2}{2} + o(t^2),$$

and so, for $t > 0$,

$$E(X) + tE(X^2) \leq \frac{b^2 t}{2} + o(t).$$

First, this yields $E(X) = o(t)$, which means that $E(X) = 0$. Second, since $E(X) = 0$,

$$tE(X^2) \leq \frac{b^2 t}{2} + o(t),$$

and then

$$E(X^2) \leq \frac{b^2}{2} + o(1),$$

which measn that $E(X^2) \leq \frac{b^2}{2}$. □

Stromberg attributes the following theorem to Saeki; further, it is proved in Stromberg that if for some t the inequality in the theorem is an equality then the random variable has the Rademacher distribution.[2]

Theorem 2. If X is a random variable satisfying $E(X) = 0$ and $P(X \in [-1, 1]) = 1$, then

$$E(e^{tX}) \leq \cosh t, \qquad t \in \mathbb{R}.$$

Proof. Define $f : \mathbb{R} \to \mathbb{R}$ by

$$f(t) = e^t \left(\cosh t - E(e^{tX})\right) = \frac{e^{2t}}{2} + \frac{1}{2} - e^t E(e^{tX}).$$

Then

$$f'(t) = e^{2t} - e^t E(e^{tX}) - e^t E(X e^{tX});$$

the derivative of $E(e^{tX})$ with respect to t is obtained using the dominated convergence theorem. Let $Y = 1 + X$, with which

$$f'(t) = e^{2t} - E(e^{tY}) - E(X e^{tY}) = e^{2t} - E(e^{tY}) - E((Y-1)e^{tY}) = e^{2t} - E(Y e^{tY}).$$

$E(X) = 0$, so $E(Y) = 1$, hence

$$f'(t) = E(e^{2t} Y) - E(Y e^{tY}) = E(Y(e^{2t} - e^{tY})).$$

Because $P(Y \in [0, 2]) = 1$, for $t \geq 0$, we have almost surely $e^{2t} - e^{tY} \geq 0$, and therefore almost surely $Y(e^{2t} - e^{tY}) \geq 0$. Therefore, for $t \geq 0$,

$$f'(t) = E(Y(e^{2t} - e^{tY})) \geq 0,$$

[2] Karl R. Stromberg, *Probability for Analysts*, p. 293, Proposition 9.9; Omar Rivasplata, *Subgaussian random variables: An expository note*, `http://www.math.ualberta.ca/~orivasplata/publications/subgaussians.pdf`

which tells us that for $t \geq 0$,

$$f(0) \leq f(t).$$

As $f(0) = 0$, for $t \geq 0$,

$$0 \leq e^t \left(\cosh t - E(e^{tX})\right),$$

and so

$$E(e^{tX}) \leq \cosh t.$$

□

Corollary 3. If a random variable X satisfies $E(X) = 0$ and $P(|X| \leq b) = 1$, then X is b-subgaussian.

2 Hoeffding's inequality

We first prove **Hoeffding's lemma**.[3]

Lemma 4 (Hoeffding's lemma)**.** If a random variable X satisfies $E(X) = 0$ and $P(X \in [a, b]) = 1$, then X is $\frac{b-a}{2}$-subgaussian.

Proof. Because $P(X \in [a, b]) = 1$, it follows that

$$\mathrm{Var}(X) \leq \frac{(b-a)^2}{4},$$

not using that $P(X) = 0$. (Namely, Popoviciu's inequality.)

Write $\mu = X_* P$ and for $\lambda \in \mathbb{R}$ define

$$d\nu_\lambda(t) = \frac{e^{\lambda t}}{e^{\Lambda(\lambda)}} d\mu(t).$$

We check

$$\int_{\mathbb{R}} d\nu_\lambda(t) = \frac{1}{e^{\Lambda(\lambda)}} \int_{\mathbb{R}} e^{\lambda t} d(X_* P)(t) = \frac{1}{e^{\Lambda(\lambda)}} \int_{\Omega} e^{\lambda X} dP = 1.$$

There is a random variable $X_\lambda : (\Omega_\lambda, \mathscr{F}_\lambda, P_\lambda) \to \mathbb{R}$ for which $X_{\lambda *} P_\lambda = \nu_\lambda$. X_λ satisfies $P_\lambda(X_\lambda \in [a, b]) = 1$, and so

$$\mathrm{Var}(X_\lambda) \leq \frac{(b-a)^2}{4}.$$

We calculate

$$\Lambda_X'(t) = \frac{E(X e^{tX})}{E(e^{tX})}$$

[3]Stéphane Boucheron, Gábor Lugosi, and Pascal Massart, *Concentration Inequalities: A Nonasymptotic Theory of Independence*, p. 27, Lemma 2.2.

and

$$\Lambda_X''(t) = \frac{E(X^2 e^{tX})E(e^{tX}) - E(Xe^{tX})E(Xe^{tX})}{E(e^{tX})^2}.$$

But

$$E(X_\lambda) = \int_{\mathbb{R}} t d\nu_\lambda(t) = \int_{\mathbb{R}} t\frac{e^{\lambda t}}{e^{\Lambda(\lambda)}} d\mu(t) = \frac{1}{e^{\Lambda(\lambda)}} E(Xe^{\lambda X})$$

and

$$E(X_\lambda^2) = \int_{\mathbb{R}} t^2 d\nu_\lambda(t) = \frac{1}{e^{\Lambda(\lambda)}} E(X^2 e^{\lambda X}),$$

and so

$$\begin{aligned}\mathrm{Var}(X_\lambda) &= E(X_\lambda^2) - E(X_\lambda)^2 \\ &= \frac{E(X^2 e^{\lambda X})}{e^{\Lambda(\lambda)}} - \frac{E(Xe^{\lambda X})^2}{e^{2\Lambda(\lambda)}} \\ &= \Lambda_X''(\lambda).\end{aligned}$$

For $\lambda \in \mathbb{R}$, Taylor's theorem tells us that there is some θ between 0 and λ such that

$$\Lambda_X(\lambda) = \Lambda_X(0) + \lambda\Lambda_X'(0) + \frac{\lambda^2}{2}\Lambda_X''(\theta) = \frac{\lambda^2}{2}\Lambda_X''(\theta);$$

here we have used that $E(X) = 0$. But from what we have shown, $\mathrm{Var}(X_\theta) = \Lambda_X''(\theta)$ and $\mathrm{Var}(X_\theta) \leq \frac{(b-a)^2}{4}$, so

$$\Lambda_X(\lambda) = \frac{\lambda^2}{2}\mathrm{Var}(X_\theta) \leq \frac{\lambda^2}{2} \cdot \frac{(b-a)^2}{4},$$

which shows that X is $\frac{b-a}{2}$-subgaussian. □

We now prove **Hoeffding's inequality**.[4]

Theorem 5 (Hoeffding's inequality)**.** Let $X_1, \ldots, X_n$ be independent random variables such that for each $1 \leq k \leq n$, $P(X_k \in [a_k, b_k]) = 1$, and write $S_n = \sum_{k=1}^n X_k$. For any $a > 0$,

$$P(S_n - E(S_n) \geq a) \leq \exp\left(-\frac{2a^2}{\sum_{k=1}^n (b_k - a_k)^2}\right).$$

Proof. For $\lambda > 0$ and $\phi(t) = e^{\lambda t}$, because ϕ is nonnegative and nondecreasing, for X a random variable we have

$$1_{X \geq a}\phi(a) \leq \phi(X),$$

and so $E(1_{X\geq a}\phi(a)) \leq E(\phi(X))$, i.e.

$$P(X \geq a) \leq \frac{E(e^{\lambda X})}{e^{\lambda a}}.$$

[4] Stéphane Boucheron, Gábor Lugosi, and Pascal Massart, *Concentration Inequalities: A Nonasymptotic Theory of Independence*, p. 34, Theorem 2.8.

Using this with $X = S_n - E(S_n)$ and because the X_k are independent,

$$P(S_n - E(S_n) \geq a) \leq \frac{1}{e^{\lambda a}} E(e^{\lambda(S_n - E(S_n))}) = e^{-\lambda a} \prod_{k=1}^{n} E(e^{\lambda(X_k - E(X_k))}).$$

Because $P(X_k \in [a_k, b_k]) = 1$, we have $P(X_k - E(X_k) \in [a_k - E(X_k), b_k - E(X_k)]) = 1$, and as $(b_k - E(X_k)) - (a_k - E(X_k)) = b_k - a_k$, Hoeffding's lemma tells us

$$\log E(e^{\lambda(X_k - E(X_k))}) \leq \frac{(b_k - a_k)^2 \lambda^2}{8},$$

and thus

$$\begin{aligned} P(S_n - E(S_n) \geq a) &\leq e^{-\lambda a} \exp\left(\sum_{k=1}^{n} \frac{(b_k - a_k)^2 \lambda^2}{8} \right) \\ &= \exp\left(-\lambda a + \frac{\lambda^2}{8} \sum_{k=1}^{n} (b_k - a_k)^2 \right). \end{aligned}$$

We remark that λ does not appear in the left-hand side. Define

$$g(\lambda) = -\lambda a + \frac{\lambda^2}{8} \sum_{k=1}^{n} (b_k - a_k)^2,$$

for which

$$g'(\lambda) = -a + \frac{\lambda}{4} \sum_{k=1}^{n} (b_k - a_k)^2.$$

Then $g'(\lambda) = 0$ if and only if

$$\lambda = \frac{4a}{\sum_{k=1}^{n} (b_k - a_k)^2},$$

at which g assumes its infimum. Then

$$\begin{aligned} P(S_n - E(S_n) \geq a) &\leq \exp\left(-\frac{4a^2}{\sum_{k=1}^{n} (b_k - a_k)^2} + \frac{16a^2}{8} \frac{1}{\sum_{k=1}^{n} (b_k - a_k)^2} \right) \\ &= \exp\left(-\frac{2a}{\sum_{k=1}^{n} (b_k - a_k)^2} \right), \end{aligned}$$

proving the claim. □

3 Cramér's large deviation theorem

The following is **Cramér's large deviation theorem**.[5]

[5] Achim Klenke, *Probability Theory: A Comprehensive Course*, p. 508, Theorem 23.3.

Theorem 6 (Cramér's large deviation theorem)**.** Suppose that $X_n : (\Omega, \mathscr{F}, P) \to \mathbb{R}$, $n \geq 1$, are independent identically distributed random variables such that for all $t \in \mathbb{R}$,

$$\Lambda(t) = \log E(e^{tX_1}) < \infty.$$

For $x \in \mathbb{R}$ define

$$\Lambda^*(x) = \sup_{t\in\mathbb{R}}(tx - \Lambda(t)).$$

If $a > E(X_1)$, then

$$\lim_{n\to\infty} \frac{1}{n} \log P(S_n \geq an) = -\Lambda^*(a),$$

where $S_n = \sum_{k=1}^n X_k$.

Proof. For $a > E(X_1)$, let $Y_n = X_n - a$, let

$$L(t) = \log E(e^{tY_1}) = \log E(e^{tX_1}e^{-ta}) = -ta + \Lambda(t)$$

and let

$$L^*(x) = \sup_{t\in\mathbb{R}}(tx - L(t)) = \sup_{t\in\mathbb{R}}(t(x+a) - \Lambda(t)) = \Lambda^*(x+a).$$

Lastly, let $T_n = \sum_{k=1}^n Y_k = S_n - na$, with which

$$P(T_n \geq bn) = P(S_n \geq (b+a)n).$$

Thus, if we have

$$\lim_{n\to\infty} \frac{1}{n} \log P(T_n \geq 0) = -L^*(0),$$

then

$$\lim_{n\to\infty} \frac{1}{n} \log P(S_n \geq an) = -L^*(0) = -\Lambda^*(a).$$

Therefore it suffices to prove the theorem for when $E(X_1) < 0$ and $a = 0$.

Define

$$\phi(t) = e^{\Lambda(t)} = E(e^{tX_1}) = \int_\Omega e^{tX_1} dP = \int_\mathbb{R} e^{tx} d(X_{1*}P)(x), \qquad t \in \mathbb{R},$$

the moment generating function of X_1, and define

$$\rho = e^{-\Lambda^*(0)} = \exp\left(-\sup_{t\in\mathbb{R}}(-\Lambda(t))\right) = \exp\left(\inf_{t\in\mathbb{R}} \Lambda(t)\right) = \inf_{t\in\mathbb{R}} \phi(t),$$

using that $x \mapsto e^x$ is increasing.

Using the dominated convergence theorem, for $k \geq 0$ we obtain

$$\phi^{(k)}(t) = \int_\mathbb{R} x^k e^{tx} d(X_{1*}P)(x) = E(X_1^k e^{tX_1}).$$

In particular, $\phi'(t) = E(X_1 e^{tX_1})$, for which $\phi'(0) = E(X_1) < 0$, and $\phi''(t) = E(X_1^2 e^{tX_1}) > 0$ for all t (either the expectation is 0 or positive, and if it is 0 then $X_1^2 e^{tX_1}$ is 0 almost everywhere, which contradicts $E(X_1) < 0$).

Either $P(X_1 \leq 0) = 1$ or $P(X_1 \leq 0) < 1$. In the first case,

$$\phi'(t) = \int_\Omega X_1 e^{tX_1} dP = \int_{X_1 \leq 0} X_1 e^{tX_1} dP \leq 0,$$

so, using the dominated convergence theorem,

$$\rho = \inf_{t\in\mathbb{R}} \phi(t) = \lim_{t\to\infty} \phi(t) = \int_{X_1\leq 0} \left(\lim_{t\to\infty} e^{tX_1} \right) dP = \int_{X_1=0} dP = P(X_1 = 0).$$

Then

$$P(S_n \geq 0) = P(X_1 = 0, \dots, X_n = 0) = P(X_1 = 0) \cdots P(X_n = 0) = \rho^n.$$

That is, as $a = 0$,

$$P(S_n \geq a) = e^{-n\Lambda^*(a)},$$

and the claim is immediate in this case.

In the second case, $P(X_1 \leq 0) < 1$. As $\phi''(t) > 0$ for all t, there is some $\tau \in \mathbb{R}$ at which $\phi(\tau) < \phi(t)$ for all $t \neq \tau$ (namely, a unique global minimum). Thus,

$$\phi(\tau) = \rho, \qquad \phi'(\tau) = 0.$$

And $\phi'(0) = E(X_1) < 0$, which with the above yields $\tau > 0$. Because $\tau > 0$, $S_n(\omega) \geq 0$ if and only if $\tau S_n(\omega) \geq 0$ if and only if $e^{\tau S_n(\omega)} \geq 1$. Applying Chebyshev's inequality, and because X_n are independent,

$$P(S_n \geq 0) = P(e^{\tau S_n} \geq 1) \leq E(e^{\tau S_n}) = E(e^{\tau X_1}) \cdots E(e^{\tau X_n}) = \phi(\tau)^n = \rho^n,$$

thus $\log P(S_n \geq 0) \leq n \log \rho$ and then

$$\limsup_{n\to\infty} \frac{1}{n} \log P(S_n \geq 0) \leq \limsup_{n\to\infty} \log \rho = \log \rho = \log e^{-\Lambda^*(0)} = -\Lambda^*(0).$$

To prove the claim, it now suffices to prove that, in the case $P(X_1 \leq 0) < 1$,

$$\liminf_{n\to\infty} \frac{1}{n} \log P(S_n \geq 0) \geq \log \rho. \tag{1}$$

Let $\mu = X_{1*}P$, and let

$$d\nu(x) = \frac{e^{\tau x}}{\rho} d\mu(x).$$

ν is a Borel probability measure: it is apparent that it is a Borel measure, and

$$\nu(\mathbb{R}) = \int_\mathbb{R} d\nu(x) = \int_\mathbb{R} \frac{e^{\tau x}}{\rho} d\mu(x) = \frac{1}{\rho} \int_\mathbb{R} e^{\tau x} d\mu(x) = \frac{\phi(\tau)}{\rho} = 1.$$

There are independent identically distributed random variables Y_n, $n \geq 1$, each with $Y_{n*}Q = \nu$. [6] Define

$$\psi(t) = E(e^{tY_1}) = \int_{\mathbb{R}} e^{tx} d\nu(x) = \int_{\mathbb{R}} e^{tx} \frac{e^{\tau x}}{\rho} d\mu(x) = \frac{1}{\rho} \int_{\mathbb{R}} e^{(t+\tau)x} d\mu(x) = \frac{\phi(t+\tau)}{\rho},$$

the moment generating function of Y_1. As $\phi'(\tau) = 0$,

$$E(Y_1) = \psi'(0) = \frac{\phi'(\tau)}{\rho} = 0.$$

As $\rho > 0$ and $\phi''(t) > 0$ for all t,

$$\mathrm{Var}(Y_1) = E(Y_1^2) = \psi''(0) = \frac{\phi''(\tau)}{\rho} \in (0, \infty).$$

For $T_n = \sum_{k=1}^n Y_k$, using that the X_n are independent and that the Y_n are independent,

$$\begin{aligned} P(S_n \geq 0) &= \int_{x_1+\cdots+x_n \geq 0} d(S_{n*}P)(x) \\ &= \int_{x_1+\cdots+x_n \geq 0} d\mu(x_1) \cdots d\mu(x_n) \\ &= \int_{x_1+\cdots+x_n \geq 0} \left(\frac{\rho}{e^{\tau x_1}} d\nu(x_1) \right) \cdots \left(\frac{\rho}{e^{\tau x_n}} d\nu(x_n) \right) \\ &= \rho^n \int_{x_1+\cdots+x_n \geq 0} e^{-\tau(x_1+\cdots+x_n)} d(T_{n*}Q). \end{aligned}$$

But

$$\begin{aligned} \int_{x_1+\cdots+x_n \geq 0} e^{-\tau(x_1+\cdots+x_n)} d(T_{n*}Q) &= \int_{T_n \geq 0} e^{-\tau T_n} dQ \\ &= E(1_{\{T_n \geq 0\}} \cdot e^{-\tau T_n}), \end{aligned}$$

hence

$$P(S_n \geq 0) = \rho^n E(1_{\{T_n \geq 0\}} \cdot e^{-\tau T_n}).$$

Thus, (1) is equivalent to

$$\liminf_{n \to \infty} \frac{1}{n} \log \left(\rho^n E(1_{\{T_n \geq 0\}} \cdot e^{-\tau T_n}) \right) \geq \log \rho,$$

so, to prove the claim it suffices to prove that

$$\liminf_{n \to \infty} \frac{1}{n} \log \left(E(1_{\{T_n \geq 0\}} \cdot e^{-\tau T_n}) \right) \geq 0.$$

[6] Gerald B. Folland, *Real Analysis: Modern Techniques and Their Applications*, p. 329, Corollary 10.19.

For any $c > 0$,

$$\begin{aligned}\log\left(E(1_{\{T_n\geq 0\}}\cdot e^{-\tau T_n})\right) &\geq \log E\left(1_{\{0\leq T_n\leq c\sqrt{n}\}}\cdot e^{-\tau T_n}\right)\\ &\geq \log\left(e^{-\tau c\sqrt{n}}\cdot Q\left(0\leq T_n\leq c\sqrt{n}\right)\right)\\ &= -\tau c\sqrt{n}+\log Q\left(\frac{T_n}{\sqrt{n}}\in[0,c]\right).\end{aligned}$$

Because the Y_n are independent identically distributed L^2 random variables with mean 0 and variance $\sigma^2 = \mathrm{Var}(Y_1) = \frac{\phi''(\tau)}{\rho}$, the central limit theorem tells us that as $n \to \infty$,

$$Q\left(\frac{T_n}{\sqrt{n}}\in[0,c]\right)\to\gamma_{0,\sigma^2}([0,c]),$$

where γ_{a,σ^2} is the Gaussian measure, whose density with respect to Lebesgue measure on $\mathbb{R}$ is

$$p(t,a,\sigma^2)=\frac{1}{\sigma\sqrt{2\pi}}e^{-\frac{(t-a)^2}{2\sigma^2}}.$$

Thus, because for $c > 0$ we have $\gamma_{0,\sigma^2}([0,c]) > 0$,

$$\begin{aligned}\liminf_{n\to\infty}\frac{1}{n}\log\left(E(1_{\{T_n\geq 0\}}\cdot e^{-\tau T_n})\right) &\geq \liminf_{n\to\infty}\left(\frac{-\tau c}{\sqrt{n}}+\frac{1}{n}\log Q\left(\frac{T_n}{\sqrt{n}}\in[0,c]\right)\right)\\ &= \lim_{n\to\infty}-\frac{\tau c}{\sqrt{n}}+\lim_{n\to\infty}\frac{1}{n}\log Q\left(\frac{T_n}{\sqrt{n}}\in[0,c]\right)\\ &= \lim_{n\to\infty}\frac{1}{n}\log\gamma_{0,\sigma^2}([0,c])\\ &= 0,\end{aligned}$$

which completes the proof. □

For example, say that X_n are independent identically distributed random variables with $X_{1*}P = \gamma_{0,1}$. We calculate that the cumulant generating function $\Lambda(t) = \log E(e^{tX_1})$ is

$$\begin{aligned}\Lambda(t) &= \log\left(\int_{\mathbb{R}}e^{tx}d\gamma_{0,1}(x)\right)\\ &= \log\left(\int_{\mathbb{R}}e^{tx}\frac{e^{-\frac{x^2}{2}}}{\sqrt{2\pi}}dx\right)\\ &= \log\left(\int_{\mathbb{R}}\frac{e^{-\frac{1}{2}(x-t)^2}}{\sqrt{2\pi}}e^{\frac{t^2}{2}}dx\right)\\ &= \log e^{t^2}2\\ &= \frac{t^2}{2},\end{aligned}$$

thus $\Lambda(t) < \infty$ for all t. Then

$$\Lambda^*(x) = \sup_{t\in\mathbb{R}}(tx - \Lambda(t)) = \sup_{t\in\mathbb{R}}\left(tx - \frac{t^2}{2}\right) = \frac{x^2}{2}.$$

Now applying Cramér's theorem we get that for $a > E(X_1) = 0$, for $S_n = \sum_{k=1}^n X_k$ we have

$$\lim_{n\to\infty}\frac{1}{n}\log P(S_n \geq an) = -\frac{a^2}{2}.$$

Another example: If X_n are independent identically distributed random variables with the **Rademacher distribution**:

$$X_{n*}P = \frac{1}{2}\delta_{-1} + \frac{1}{2}\delta_1.$$

Then

$$E(e^{tX_1}) = \int_{\mathbb{R}} e^{tx} d\left(\frac{1}{2}\delta_{-1} + \frac{1}{2}\delta_1\right)(x) = \frac{1}{2}e^{-t} + \frac{1}{2}e^t = \cosh t,$$

so the cumulant generating function of X_1 is

$$\Lambda(t) = \log\cosh t,$$

and indeed $\Lambda(t) < \infty$ for all $t \in \mathbb{R}$. Then, as $\frac{d}{dt}(tx - \log\cosh t) = x - \tanh t$,

$$\Lambda^*(x) = \sup_{t\in\mathbb{R}}(tx - \log\cosh t) = \operatorname{arctanh} x \cdot x - \log\cosh\operatorname{arctanh} x.$$

For $x \in (-1, 1)$,

$$\operatorname{arctanh} x = \frac{1}{2}\log\frac{1+x}{1-x}.$$

Then

$$\cosh\operatorname{arctanh} x = \frac{1}{2}\left(e^{\operatorname{arctanh} x} + e^{-\operatorname{arctanh} x}\right) = \frac{1}{2}\sqrt{\frac{1+x}{1-x}} + \frac{1}{2}\sqrt{\frac{1-x}{1+x}} = \frac{1}{\sqrt{1-x^2}}.$$

With these identities,

$$\begin{aligned}
\Lambda^*(t) &= \frac{x}{2}\log\frac{1+x}{1-x} + \frac{1}{2}\log(1-x^2) \\
&= \frac{x}{2}\log(1+x) - \frac{x}{2}\log(1-x) + \frac{1}{2}\log(1+x) + \frac{1}{2}\log(1-x) \\
&= \frac{1+x}{2}\log(1+x) + \frac{1-x}{2}\log(1-x).
\end{aligned}$$

With $S_n = \sum_{k=1}^n X_k$, applying Cramér's theorem, we get that for any $a > E(X_1) = 0$,

$$\lim_{n\to\infty}\frac{1}{n}\log P(S_n \geq an) = -\frac{1+x}{2}\log(1+x) - \frac{1-x}{2}\log(1-x).$$

For a Borel probability measure μ on $\mathbb{R}$, we define its **Laplace transform** $\check{\mu} : \mathbb{R} \to (0, \infty]$ by

$$\check{\mu}(t) = \int_{\mathbb{R}} e^{ty} d\mu(y).$$

Suppose that $\int_{\mathbb{R}} |y| d\mu(y) < \infty$ and let $M_1 = \int_{\mathbb{R}} y d\mu(y)$, the first moment of μ. For any t the function $x \mapsto e^{tx}$ is convex, so by Jensen's inequality,

$$e^{tM_1} \leq \int_{\mathbb{R}} e^{ty} d\mu(y) = \check{\mu}(t).$$

Thus for all $t \in \mathbb{R}$,

$$tM_1 - \log \check{\mu}(t) \leq 0.$$

For a Borel probability measure μ with finite first moment, we define its **Cramér transform** $I_\mu : \mathbb{R} \to [0, \infty]$ by[7]

$$I_\mu(x) = \sup_{t \in \mathbb{R}} (tx - \log \check{\mu}(t)).$$

For $t = 0$, $tx - \log \check{\mu}(t) = -\log \check{\mu}(0) = -\log(1) = 0$, which shows that indeed $0 \leq I_\mu(x) \leq \infty$ for all $x \in \mathbb{R}$. But $tM_1 - \log \check{\mu}(t) \leq 0$ for all t yields

$$I_\mu(M_1) = 0.$$

[7] Heinz Bauer, *Probability Theory*, pp. 89–90, §12.