DAI Assignment-2

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September 8, 2024

Contents

1 Mathemagic

1.1 Task A

For a Bernoulli random variable $X \sim Ber(p)$, $P[X = 0] = 1 - p$, $P[X = 1] = p$ and $P[X = n] = 0 \forall n \geq 2$.

$$\begin{aligned}
G_{Ber}(z) &= \sum_{n=0}^{\infty} P[X = n] z^n \\
&= P[X = 0] + P[X = 1]z \\
G_{Ber} &= (1 - p) + pz
\end{aligned}$$

1.2 Task B

When $X \sim Bin(n, p)$, $P[X = k] = \binom{n}{k} p^k (1 - p)^{n-k}\ \forall k \in \mathbb{Z}, 0 \leq k \leq n$. $P[X = k] = 0\ \forall k > n$.

$$\begin{aligned}
G_{Bin}(z) &= \sum_{k=0}^{\infty} P[X = n] z^k \\
&= \sum_{k=0}^{n} \binom{n}{k} p^k (1 - p)^{n-k} z^k \\
&= \sum_{k=0}^{n} \binom{n}{k} (pz)^k (1 - p)^{n-k} \\
&= (pz + (1 - p))^n \\
G_{Bin}(z) &= G_{Ber}(z)^n
\end{aligned}$$

1.3 Task C

Let X_1 and X_2 be two random variables which take up non-negative integers and let $X = X_1 + X_2$.

$$
\begin{aligned}
G_X(z) &= \sum_{n=0}^{\infty} P[X=n]z^n \\
&= \sum_{n=0}^{\infty}\sum_{i=0}^{n}(P[X_1=i]P[X_2=n-i])z^n \\
&= \sum_{i=0}^{\infty}\sum_{n=i}^{\infty}(P[X_1=i]P[X_2=n-i])z^n \\
&= \sum_{i=0}^{\infty}\sum_{n=i}^{\infty}(P[X_1=i]z^i)(P[X_2=n-i]z^{n-i}) \\
&= \sum_{i=0}^{\infty}P[X_1=i]z^i\sum_{n=i}^{\infty}P[X_2=n-i]z^{n-i} \\
&= \sum_{i=0}^{\infty}P[X_1=i]z^i\sum_{n=0}^{\infty}P[X_2=n]z^n \\
G_X(z) &= G_{X_1}(z)G_{X_2}(z)
\end{aligned}
$$

For $k=1$, $G_\Sigma(z)=G(z)^k$.
If for $k=n-1$, $G_\Sigma(z)=G(z)^k$, then for $k=n$,

$$
\begin{aligned}
G_\Sigma(z) &= G_{X_1+X_2+\cdots+X_{n-1}}(z)G_{X_n}(x) \\
&= \prod_{i=1}^{n}G_{X_i}(z)G_\Sigma(z) && = G(z)^k
\end{aligned}
$$

From the *Principle of Mathematical Induction*, for all $k\in\mathbb{N}$, $G_\Sigma(z)=G(z)^k$

1.4 Task D

When $X\sim Geo(p)$, $P[X=n]=(1-p)^{n-1}p$ and $P[X=0]=0$.

$$
\begin{aligned}
G_{Geo}(z) &= \sum_{n=0}^{\infty}P[X=n]z^n \\
&= \sum_{n=1}^{\infty}(1-p)^{n-1}pz^n \\
&= pz\sum_{n=1}^{\infty}((1-p)z)^{n-1} \\
&= \frac{pz}{1-(1-p)z}
\end{aligned}
$$

1.5 Task E

$X \sim Bin(n,p)$ and $Y \sim NegBin(n,p)$. $Y = \sum_{k=1}^{n} Y_k$ where $Y_k \sim Geo(p)$.

$$
\begin{aligned}
G(Y) &= G(Y_1 + Y_2 + \cdots + Y_n) \\
&= G(Y_1)G(Y_2)\cdots G(Y_n) \\
&= \left(\frac{pz}{1-(1-p)z}\right)^n \\
G_Y^{(n,p)}(z) &= \left(\frac{pz}{1-(1-p)z}\right)^n \\
&= \left(\frac{1}{\frac{1}{pz} + (1-\frac{1}{p})}\right)^n \\
G_X^{(n,p^{-1})}(z^{-1}) &= \left(\frac{1}{pz} + (1-\frac{1}{p})\right)^n \\
\implies G_Y^{(n,p)}(z) &= \left(G_X^{(n,p^{-1})}(z^{-1})\right)^{-1}
\end{aligned}
$$

1.6 Task F

$P[Y=k] = \binom{k-1}{n-1} p^n (1-p)^{k-n}$.

$$
\begin{aligned}
G_Y^{(n,p)}(z) &= \sum_{k=0}^{\infty} P[Y=k] z^k \\
&= \sum_{k=n}^{\infty} \binom{k-1}{n-1} p^n (1-p)^{k-n} z^k \\
\left(\frac{1}{pz} + \left(1 - \frac{1}{p}\right)\right)^{-n} &= \sum_{k=n}^{\infty} \binom{k-1}{n-1} p^n (1-p)^{k-n} z^k
\end{aligned}
$$

Substitute $x = z$, $p = 2$

$$
\begin{aligned}
\left(\frac{1}{2x}+\frac{1}{2}\right)^{-n} &= \sum_{k=n}^{\infty}\binom{k-1}{n-1}2^n(-1)^{k-n}x^k \\
x^n(1+x)^{-n} &= \sum_{k=n}^{\infty}\binom{k-1}{n-1}(-1)^{k-n}x^k \\
(1+x)^{-n} &= \sum_{k=n}^{\infty}\binom{k-1}{n-1}(-1)^{k-n}x^{k-n} \\
(1+x)^{-n} &= \sum_{r=0}^{\infty}\binom{n+r-1}{n-1}(-1)^r x^r \\
(1+x)^{-n} &= \sum_{r=0}^{\infty}\binom{n+r-1}{r}(-1)^r x^r
\end{aligned}
$$

1.7 Task G

$$
\begin{aligned}
G(z) &= \sum_{n=0}^{\infty}P[X=n]z^n \\
G'(z) &= \sum_{n=0}^{\infty}nP[X=n]z^{n-1} \\
G'(1) &= \sum_{n=0}^{\infty}nP[X=n] \\
G'(1) &= E[X]
\end{aligned}
$$

For Bernoulli random variable $X \sim Ber(p)$,

$$
\begin{aligned}
G_{Ber}(z) &= (1-p)+pz \\
G'_{Ber}(z) &= p \\
G'_{Ber}(1) &= p \\
E[X] &= p
\end{aligned}
$$

For Binomial random variable $X \sim Bin(n,p)$,

$$
\begin{aligned}
G_{Bin}(z) &= (pz+(1-p))^n \\
G'_{Bin}(z) &= np(pz+(1-p))^{n-1} \\
G'_{Bin}(1) &= np \\
E[X] &= np
\end{aligned}
$$

For Geometric random variable $X \sim Geo(p)$,

$$\begin{aligned}
G_{Geo}(z) &= \frac{pz}{1-(1-p)z} \\
G'_{Geo}(z) &= \frac{p}{(1-(1-p)z)^2} \\
G'_{Geo}(1) &= \frac{p}{p^2} \\
E[X] &= \frac{1}{p}
\end{aligned}$$

For Negative Binomial random variable $X \sim NegBin(n, p)$,

$$\begin{aligned}
G_{NegBin}(z) &= \left(\frac{pz}{1-(1-p)z}\right)^n \\
G'_{NegBin}(z) &= n\left(\frac{pz}{1-(1-p)z}\right)^{n-1} \frac{p}{(1-(1-p)z)^2} \\
G'_{NegBin}(1) &= n\left(\frac{p}{p}\right)^{n-1} \frac{p}{p^2} \\
E[X] &= \frac{n}{p}
\end{aligned}$$

2 Normal Sampling

2.1 Task A

To prove that Y is uniformly distributed in [0,1], we need to prove that for an arbitrary interval [a,b] in [0, 1], Y lies in [a, b] with probability $b-a$.
The probability that $a \leq Y \leq b$ is the probability that the randomly chosen X is such that $a \leq F_X(x) \leq b$. The cumulative probability of x being in the interval $[a_0, b_0]$ such that $F_X(a_0) = a$ and $F_X(b_0) = b$ is $b-a$, since the cumulative probability that $x \leq b_0$ is b, and the cumulative probability of $x \leq a_0$ is a, hence probability that $a_0 \leq x \leq b_0$ is $b-a$. Our condition that $F_X()$ is invertible guarantees that the $\leq$ can be replaced with $<$, since it cannot be that $\lim_{x\to a_0^-} \neq \lim_{x\to a_0^+}$, otherwise between a_0^- and a_0^+, all elements would be mapped to by a_0.

2.2 Task B

Given a sample y, let us return $F_X^{-1}(y)$. To prove that this has the same CDF as X, we need to prove that if $F_X(a_0) = a$ and $F_X(b_0) = b$, the probability of getting $a_0 \leq F_X^{-1}(y) \leq b_0$ is $b-a$.
But this is fairly straightforward, since $a_0 \leq F_X^{-1}(y) \leq b_0 \iff a \leq y \leq b$, and the probability of getting $a \leq y \leq b$ is $b-a$.
Thus, this is the required function.

2.3 Task C

Code provided in 2c.ipynb.

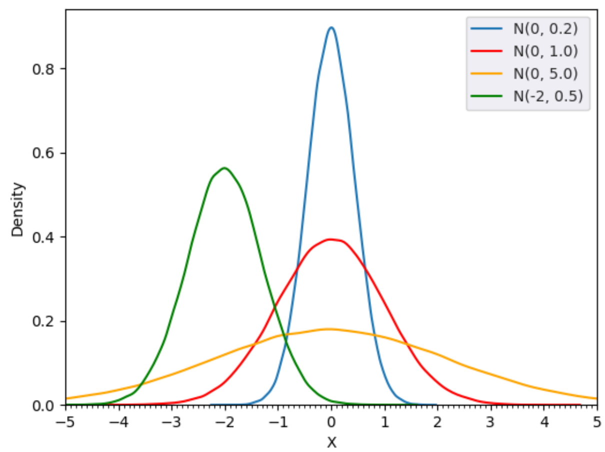

Figure 1: Histogram of the samples generated

2.4 Task D

Code provided in 2d.ipynb.

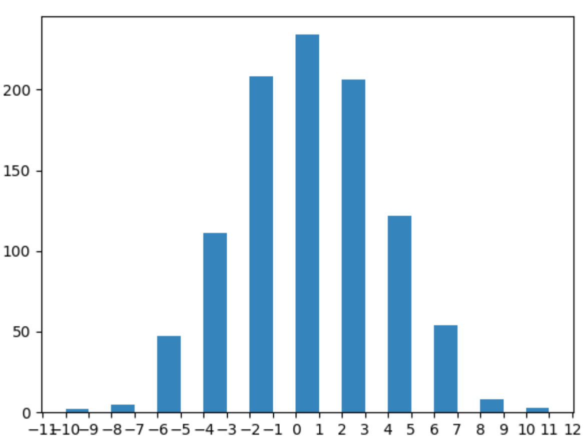

Figure 2: h=10

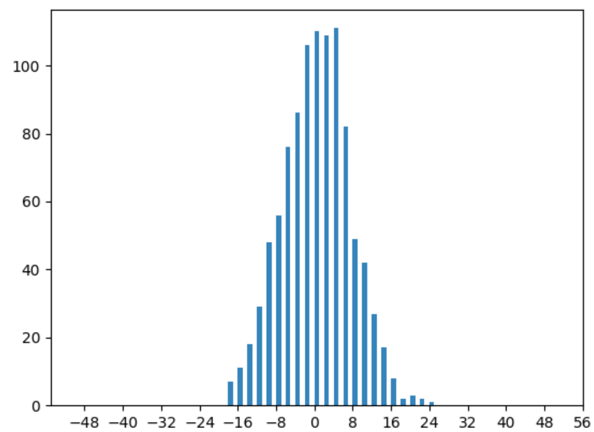

Figure 3: h=50

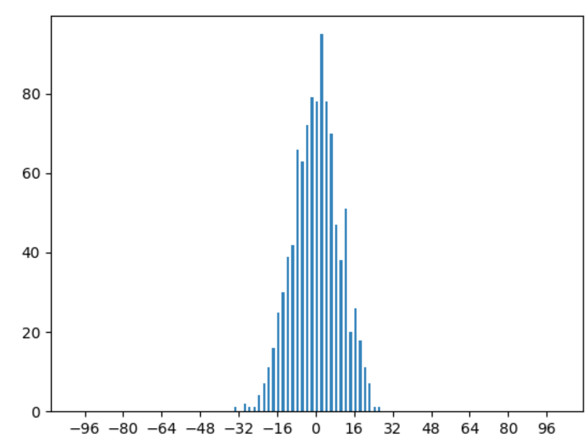

Figure 4: h=100

For h=100 alone, for better visibility, I have also written code with bins set to increment every 2 pockets. To observe the property of only even indexed bins being filled, simply replace range(-100, 102, 2) with range(-100, 102, 1).

2.5 Task E

$$P_h[X = 2i] = \frac{\binom{2k}{i+k}}{2^{2k}}$$

Since the likelihood of a ball arriving at pocket i is equal to the likelihood of having i+k collisions which move it to the right within the 2k possible collisions overall.
Applying Stirling's approximations,

$$\begin{aligned}
P_h[X = 2j] &\approx \frac{k^{2k}\sqrt{4\pi k}}{(k+j)^{k+j}\sqrt{2\pi(k+j)}(k-j)^{k-j}\sqrt{2\pi(k-j)}} \\
&\approx \frac{k^{2k}\sqrt{2k}(k-j)^j}{(k^2-j^2)^k\sqrt{2\pi(k^2-j^2)}(k+j)^j} \\
&\approx \frac{1}{\sqrt{\pi k}}\frac{k^{2k}(k-j)^j}{(k^2)^k(k+j)^j} && \text{(By approximating } k^2 - j^2 \approx j^2\text{)} \\
&= \frac{1}{\sqrt{\pi k}}\left(1 - \frac{2j}{k+j}\right)^j
\end{aligned}$$

Now, we independently approximate $\left(1 - \frac{2j}{k+j}\right)^j$ by taking its logarithm and using the property that for small x, $\ln(1-x) \approx -x$.

$$\begin{aligned}
\ln\left(\left(1 - \frac{2j}{k+j}\right)^j\right) &= j\ln\left(1 - \frac{2j}{k+j}\right) \\
&\approx -\frac{2j^2}{k+j} \\
\left(1 - \frac{2j}{k+j}\right)^j &\approx e^{-\frac{2j^2}{k}}
\end{aligned}$$

Setting $i = 2j$ as required, we get

$$\begin{aligned}
P_h[X = i] &\approx \frac{1}{\sqrt{\pi k}}e^{-\frac{i^2}{2k}} \\
&= \frac{1}{\sqrt{2\pi k}}e^{-\frac{i^2}{h}}
\end{aligned}$$

As required.

3 Quality in Inequalities

3.1 Task A

$$\begin{aligned}
\mathbb{E}[X] &= \int_0^\infty XP(X)\,dX \\
&= \int_0^a XP(X)\,dX + \int_a^\infty XP(X)\,dX \\
&\geq \int_0^a XP(X)\,dX + \int_a^\infty aP(X)\,dX \\
&\geq \int_0^a XP(X)\,dX + aP[X \geq a] \\
&\geq aP[X \geq a]
\end{aligned}$$

Here, we have used the fact that $\int_0^a XP(X)dX$ is ≥ 0 since $X \geq 0$ and $P(X) \geq 0$ everywhere in that interval.
From the above result, we get

$$\frac{\mathbb{E}[X]}{a} \geq P[X \geq a]$$

Intuitively, this corresponds to saying that if there is a high probability that $X \geq a$, then we can also expect that a random sample will give us a value greater than a.

3.2 Task B

First, we will prove a result which is useful later

$$\mathbb{E}[\left(X - \mathbb{E}[X] + \frac{\sigma^2}{\tau}\right)^2] = Var(X - \mathbb{E}[X] + \frac{\sigma^2}{\tau}) + \left(E[X - \mathbb{E}[X] + \frac{\sigma^2}{\tau}]\right)^2$$

Now, since the variance of a distribution does not change on adding a constant to the distribution, and since the expectation value of the sum of terms is the sum of expectations of the terms, we get that this is equal to

$$\mathbb{E}[\left(X - \mathbb{E}[X] + \frac{\sigma^2}{\tau}\right)^2] = \sigma^2 + \left(\frac{\sigma^2}{\tau}\right)^2$$

Now, we notice that whenever $X - \mu \geq \tau$, it also holds that $\left(X - \mu + \frac{\sigma^2}{\tau}\right)^2 \geq \left(\tau + \frac{\sigma^2}{\tau}\right)^2$. Since we are assuming that $x - \mu + \frac{\sigma^2}{\tau} \geq \tau + \frac{\sigma^2}{\tau} \geq 0$, this result is proven by the fact that x^2 is an increasing function over the non-negative reals.
Thus, $P[X - \mu \geq \tau] \leq P[\left(X - \mu + \frac{\sigma^2}{\tau}\right)^2 \geq \left(\tau + \frac{\sigma^2}{\tau}\right)^2]$.

Using Markov's inequality on the above result, we get

$$P[X-\mu \geq \tau] \leq P[\left(X-\mu+\frac{\sigma^2}{\tau}\right)^2 \leq \left(\tau+\frac{\sigma^2}{\tau}\right)^2] \leq \frac{\sigma^2+\left(\frac{\sigma^2}{\tau}\right)^2}{\left(\tau+\frac{\sigma^2}{\tau}\right)^2}$$
$$= \frac{\sigma^2}{\sigma^2+\tau^2}$$

Thus proving our desired result.

3.3 Task C

$$P[X \geq x] = \int_x^\infty P(X)\,dX$$

$$\begin{aligned}
e^{-tx}E[e^{tx}] &= e^{-tx}\int_{-\infty}^{\infty} e^{tX}P(X)\,dX \\
&= e^{-tx}\int_{-\infty}^{x} e^{tX}P(X)\,dX + e^{-tx}\int_{x}^{\infty} e^{tX}P(X)\,dX \\
&\geq e^{-tx}\int_{x}^{\infty} e^{tX}P(X)\,dX && \text{(since both P(X) and } e^{tX} \text{ are positive for all X)} \\
&\geq e^{-tx}e^{tx}\int_{x}^{\infty} P(X)\,dX && \text{(since } e^{tX} \text{ is increasing if t¿0)} \\
&= e^{-tx}e^{tx}P[X \geq x] \\
&= P[X \geq x]
\end{aligned}$$

This proves that

$$P[X \geq x] \leq e^{-tx}E[e^{tx}] \tag{1}$$

Similarly,

$$P[X \leq x] = \int_{-\infty}^x P(X)\,dX$$

$$\begin{aligned}
e^{-tx}E[e^{tx}] &= e^{-tx}\int_{-\infty}^{\infty} e^{tX}P(X)\,dX \\
&= e^{-tx}\int_{-\infty}^{x} e^{tX}P(X)\,dX + e^{-tx}\int_{x}^{\infty} e^{tX}P(X)\,dX \\
&\geq e^{-tx}\int_{-\infty}^{x} e^{tX}P(X)\,dX && \text{(since both P(X) and } e^{tX} \text{ are positive for all X)} \\
&\geq e^{-tx}\cdot e^{tx}\cdot\int_{-\infty}^{x} P(X)\,dX && \text{(since } e^{tx} \leq e^{tx'} \text{ where } x \geq x' \text{ when t is negative)} \\
&= \int_{-\infty}^{x} P(X)\,dX \\
&= P[X \leq x]
\end{aligned}$$

as required.

3.4 Task D

1. We know, by the linearity of expectation, that

$$\mathbb{E}[Y] = \sum_{i=1}^{n} \mathbb{E}[X_i] = \sum_{i=1}^{n} p_i$$

2. For this question, we make use of the first result proved in Task C above1.

$$\begin{aligned} P[Y \geq (1+\delta)\mu] &\leq \frac{\mathbb{E}[e^{tY}]}{e^{(1+\delta)\mu t}} \\ &= \frac{\mathbb{E}[e^{\left(t\sum_{i=1}^{n} X_i\right)}]}{e^{(1+\delta)\mu t}} \end{aligned}$$

We will now prove that for independent random variables, $\mathbb{E}[e^{\left(t\sum_{i=1}^{n} X_i\right)}] = \prod_{i=1}^{n} \mathbb{E}[e^{tX_i}]$. To prove that, we prove the simpler result that for independent variables, $\mathbb{E}[XY] = \mathbb{E}[X]\mathbb{E}[Y]$.

$$\begin{aligned} \mathbb{E}[XY] &= \int_{-\infty}^{\infty} (xy) p_{X,Y}(x,y)\, dy\, dx \\ &= \int_{-\infty}^{\infty} (xy) p_X(x) p_Y(y)\, dy\, dx && \text{(For independent RVs, p(xy)=p(x)p(y))} \\ &= \int_{-\infty}^{\infty} x p_X(x) \int_{-\infty}^{\infty} y p_Y(y)\, dy\, dx \\ &= \int_{-\infty}^{\infty} x p_X(x)\, dx \int_{-\infty}^{\infty} y p_Y(y)\, dy \\ &= \mathbb{E}[X]\mathbb{E}[Y] \end{aligned}$$

Since above, all X_i's are independent, so are e^{tX_i} for all i. Thus,

$$\begin{aligned} E[e^{\left(t\sum_{i=1}^{n} X_i\right)}] &= \mathbb{E}[e^{tX_1}] E[e^{\left(t\sum_{i=2}^{n} X_i\right)}] \\ &= \mathbb{E}[e^{tX_1}]\mathbb{E}[e^{tX_2}]\mathbb{E}[e^{\left(t\sum_{i=3}^{n} X_i\right)}] \\ &\dots \\ &= \prod_{i=1}^{n} \mathbb{E}[e^{tX_i}] \end{aligned}$$

For each Bernoulli random variable X_i, it is easily observed that $E[e^{tX_i}] = 1 + (e^t - 1)p_i$. Using this result,

$$E[e^{\left(t\sum_{i=1}^{n} X_i\right)}] = \prod_{i=1}^{n} (1 + (e^t - 1)p_i)$$

We will now prove $\prod_{i=1}^{n}(1+x_i) \leq e^{\sum_{i=1}^{n} x_i}$ for positive x_i's.

$$\begin{aligned}
\ln\left(\prod_{i=1}^{n}(1+x_i)\right) &= \sum_{i=1}^{n}\ln(1+x_i) \\
&\leq \sum_{i=1}^{n} x_i \qquad\qquad \text{(Well known result that } \ln(1+x) \leq x\text{)} \\
&= \ln\left(e^{\sum_{i=1}^{n} x_i}\right)
\end{aligned}$$

Since our initial assumption in using 1 required $t > 0$ and hence $e^t > 1$, all $(e^t - 1)p_i$ values are positive and we can use the above result. Hence, $\mathbb{E}[e^{tY}] \leq e^{\mu(e^t-1)}$.
Plugging this into **??**, we get

$$P[Y \geq (1+\delta)\mu] \leq \frac{e^{\mu(e^t-1)}}{e^{(1+\delta)t\mu}}$$

as required.

3. Since this holds for arbitrary positive t, we can find the t for which the value of $e^{\mu(e^t-1-(1+\delta)t)}$ is minimum. By differentiation, we observe the minima of the function is when $t = \ln(1+\delta)$. At the minima, we have

$$e^t - 1 - (1+\delta)t = \delta - (1+\delta)\ln(1+\delta)$$

Using the expansion $\ln(1+\delta) = \sum_{i=1}^{\infty} \frac{\delta^i(-1)^{i-1}}{i}$,

$$\begin{aligned}
\delta - (1+\delta)\ln(1+\delta) &= \delta + (1+\delta)(\sum_{i=1}^{\infty}\frac{\delta^i(-1)^i}{i}) \\
&= \delta + \sum_{i=1}^{\infty}\frac{\delta^i(-1)^i}{i} + \sum_{i=2}^{\infty}\frac{\delta^i(-1)^{i-1}}{i-1} \\
&= \sum_{i=2}^{\infty}\frac{\delta^i(-1)^{i-1}}{i(i-1)}
\end{aligned}$$

This is approximately $-\frac{\delta^2}{2} + \frac{\delta^3}{6}$ for small delta. Thus we get a good upper bound of $e^{\frac{-\mu\delta^2(3-\delta)}{6}}$ for $P[Y \geq (1+\delta)\mu]$.
If $\delta > 0$, this can be simplified further, in a way that is more useful for us in question 5. Observe that for $x > 0$

$$\ln(1+x) \geq \frac{2x}{2+x}$$

This can be seen by observing that at x=0, $\ln(1+x) - \frac{2x}{2+x}$ evaluates to 0 and that its derivative, $\frac{4}{(1+x)(2+x)^2}$ is always non-negative.

Thus,

$$\mu(\delta - (1+\delta)\ln(1+\delta)) \leq \mu(\delta - \frac{2\delta(1+\delta)}{2+\delta})$$
$$= \frac{-\mu\delta^2}{2+\delta}$$

Thus, we get,

$$P[Y \geq (1+\delta)\mu] \leq e^{\frac{-\mu\delta^2}{2+\delta}} \tag{2}$$

For the last few steps of simplification, https://math.mit.edu/ goemans/18310S15/chernoff-notes.pdf was used as a guide.

3.5 Task E

First, let us consider the corner cases where $\mu = 0$ or $\mu = 1$. In these cases, $A_n = 0$ or $A_n = 1$ always, and the result is trivial.
Now we can consider $0 < \mu < 1$.
Consider the random variable $B_n = nA_n$.
We need to prove $\lim_{n\to\infty} P[|B_n - n\mu| > n\epsilon] = 0$.
Note that we are done if we prove $\lim_{n\to\infty} P[B_n - n\mu >= n\epsilon] = 0$ and $\lim_{n\to\infty} P[n\mu - B_n >= n\epsilon] = 0$.
We will prove the first part first, assuming μ is not 0.

$$P[B_n \geq n\epsilon + n\mu] \leq e^{-\frac{n\epsilon^2}{2\mu+\epsilon}}$$

For a positive μ (since the distribution is the sum of Bernoulli RVs) and ϵ (given), $\frac{\epsilon^2}{2\mu+\epsilon}$ is a positive coefficient. e^{-n} tends to 0 as n tends to ∞, which we will use henceforth.
Note that the result that $P[X \geq (1+\delta)\mu] \leq e^{\frac{-\mu\delta^2}{2+\delta}}$ was not proven without referring to the internet, so I will prove the claim above using an alternative method which assumes only what has been proven by me.

$$P[X \geq (1+\delta)\mu] \leq e^{\mu(\delta-(1+\delta)\ln(1+\delta))}$$

We want to prove that $f(\delta) = \delta - (1+\delta)\ln(1+\delta)$ is strictly negative if δ is positive. For that, notice that $f(0) = 0$ and $f'(\delta) = -\ln(1+\delta)$, which is non-positive for all non-negative δ. Thus, $f(\delta)$ is strictly negative for all positive δ.
This again means that $P[B_n \geq n\epsilon + n\mu] \leq e^{-nk}$ for some positive coefficient k. Thus, $P[B_n \geq n\epsilon + n\mu]$ tends to 0 as n tends to ∞.
Similarly, $P[B_n \leq n\mu - n\epsilon]$ also tends to 0 as n tends to ∞.
We can prove this by considering the RV $1 - A_n$ instead of A_n and using the same logic as above.
$P[n - B_n - n - n\mu \geq n\delta] = P[n\mu - nB_n \geq n\delta] \leq e^{-nk}$ for some positive coefficient k.

Now, combining the two results,

$$
\begin{aligned}
P[|B_n - n\mu| > n\epsilon] &\leq P[B_n - n\mu \geq n\epsilon] + P[n\mu - B_n \geq n\epsilon] \\
&\leq e^{-nk} + e^{-nk'} && \text{(for some positive coefficients k and k')}
\end{aligned}
$$

If we get any positive value σ for the value of $P[|B_n - n\mu| > n\epsilon]$, simply set $n > \max(\frac{\ln(\frac{2}{\sigma})}{k}, \frac{\ln(\frac{2}{\sigma})}{k'})$ and we have

$$
\begin{aligned}
P[|B_n - n\mu| > n\epsilon] &\leq e^{-nk} + e^{-nk'} \\
&< \frac{\sigma}{2} + \frac{\sigma}{2} \\
&= \sigma
\end{aligned}
$$

which is a contradiction.
Thus, we have proven that $\lim_{n\to\infty} P[|B_n - n\mu| > n\epsilon] = \lim_{n\to\infty} P[|A_n - \mu| > \epsilon] = 0$.

4 A Pretty "Normal" Mixture

4.1 Task A

Let A_i be the event in which you choose i in the first step of the algorithm. Then the set of all the A_is is a valid partition, since they are completely disjoint and their union represents all the possibilities.
Hence we can use the *law of Total Probability* to get,

$$
\begin{aligned}
P[\mathcal{A} = x] &= \sum_{i=1}^{K} P[A_i] \cdot P[X_i = x] \\
&= \sum_{i=1}^{K} p_i \cdot P[X_i = x] \\
&= P[X = x] \\
\implies f_{\mathcal{A}}(x) &= f_X(x)
\end{aligned}
$$

4.2 Task B

4.2.1 Part 1

$$\begin{aligned}
\mathbb{E}[X] &= \sum_{x\in\mathbb{Z}}\left(x\cdot\sum_{i=1}^{K}p_i\cdot P[X_i=x]\right)\\
&= \sum_{i=1}^{K}\left(p_i\cdot\sum_{x\in\mathbb{Z}}x\cdot P[X_i=x]\right)\\
&= \sum_{i=1}^{K}\left(p_i\cdot\mathbb{E}[X_i]\right)=\sum_{i=1}^{K}\left(p_i\cdot\mu_i\right)
\end{aligned}$$

4.2.2 Part 2

- Calculating $\mathbb{E}[X^2]$

$$\begin{aligned}
\mathbb{E}[X^2] &= \sum_{x\in\mathbb{Z}}\left(x^2\cdot\sum_{i=1}^{K}p_i\cdot P[X_i=x]\right)\\
&= \sum_{i=1}^{K}\left(p_i\cdot\sum_{x\in\mathbb{Z}}x^2\cdot P[X_i=x]\right)\\
&= \sum_{i=1}^{K}\left(p_i\cdot\mathbb{E}[X_i^2]\right)=\sum_{i=1}^{K}\left(p_i\cdot\left(Var(X_i)+\left(\mathbb{E}[X_i]\right)^2\right)\right)\\
&= \sum_{i=1}^{K}\left(p_i\cdot\left(\sigma_i^2+\mu_i^2\right)\right)
\end{aligned}$$

- Calculating $\left(\mathbb{E}[X]\right)^2$

$$\begin{aligned}
\left(\mathbb{E}[X]\right)^2 &= \left(\sum_{i=1}^{K}p_i\cdot\mathbb{E}[X_i]\right)^2\\
&= \sum_{i=1}^{K}\left(p_i\cdot\mathbb{E}[X_i]\right)^2+2\sum_{1\le i<j\le K}\left(p_i\cdot\mathbb{E}[X_i]\cdot p_j\cdot\mathbb{E}[X_j]\right)\\
&= \sum_{i=1}^{K}\left(p_i\mu_i\right)^2+2\sum_{1\le i<j\le K}\left(p_i\mu_i p_j\mu_j\right)
\end{aligned}$$

- Calculating Var[X]

$$\begin{aligned}
Var[X] &= \left(\mathbb{E}[X]\right)^2-\mathbb{E}[X^2]\\
&= \left(\sum_{i=1}^{K}\left(p_i\mu_i\right)^2+2\sum_{1\le i<j\le K}\left(p_i\mu_i p_j\mu_j\right)\right)-\sum_{i=1}^{K}\left(p_i\cdot\left(\sigma_i^2+\mu_i^2\right)\right)
\end{aligned}$$

4.2.3 Part 3

$$
\begin{aligned}
M_X(t) = G(e^t) &= \int_{-\infty}^{\infty} e^{xt} P[X = x] dx = \int_{-\infty}^{\infty} e^{xt} \sum_{i=1}^{K} p_i P[X_i = x] dx \\
&= \int_{-\infty}^{\infty} \sum_{i=1}^{K} e^{xt} p_i P[X_i = x] dx = \sum_{i=1}^{K} p_i \int_{-\infty}^{\infty} e^{xt} P[X_i = x] dx \\
&= \sum_{i=1}^{K} p_i M_{X_i} = \sum_{i=1}^{K} p_i e^{\mu_i t + \frac{1}{2}\sigma_i^2 t^2}
\end{aligned}
$$

4.3 Task C

4.3.1 Part 1

$$
\mathbb{E}[Z] = \mathbb{E}\left[\sum_{i=1}^{K} p_i X_i\right] = \sum_{i=1}^{K} p_i \cdot \mathbb{E}\left[X_i\right] = \sum_{i=1}^{K} (p_i \mu_i)
$$

4.3.2 Part 2

$$
Var(Z) = Var\left(\sum_{i=1}^{K} p_i X_i\right) = \sum_{i=1}^{K} Var\left(p_i X_i\right) = \sum_{i=1}^{K} p_i^2 Var\left(X_i\right) = \sum_{i=1}^{K} p_i^2 \sigma_i^2
$$

4.3.3 Part 3

Since we already calculated μ_Z and $\sigma_Z = \sqrt{Var(Z)}$ We can just use the expression for Gaussian Random Variable (See Part 6 for the proof that Z is a Gaussian random variable)

$$
f_Z(u) = \frac{1}{\sigma_Z \sqrt{2\pi}} e^{\frac{-(u-\mu_Z)^2}{2\sigma_Z^2}}
$$

4.3.4 Part 4

Using the expression for Moment generating function of a Gaussian Random Variable (See Part 6 for the proof that Z is a gaussian random variable)

$$
M_Z(t) = e^{\mu_Z t + \frac{1}{2}\sigma_Z^2 t^2}
$$

4.3.5 Part 5

No, It can be seen that even though the expected value is the same in both the case, The values of both variance and MGF are different for both X and Z hence these are different random variables

4.3.6 Part 6

We can show that linear combination of independent random variables is also a random variable. For any random Gaussian variable X is given by $M_X(t) = \mathbb{E}[e^{tX}] = e^{\mu t + \frac{1}{2}\sigma^2 t^2}$
Let $Y = \sum_{i=1}^{K} a_i X_i$ where each X_i is an independent Gaussian random variable.Then

$$\begin{aligned}
M_Y(t) = \mathbb{E}[e^{tY}] = \mathbb{E}\left[e^{t\sum_{i=1}^{K} a_i X_i}\right] = \mathbb{E}\left[\prod_{i=1}^{K} e^{t a_i X_i}\right] \\
= \prod_{i=1}^{K} \mathbb{E}\left[e^{t a_i X_i}\right] \qquad \text{,Since all the } X_i \text{ are mutually independent} \\
= \prod_{i=1}^{K} M_{a_i X_i}(t) = \prod_{i=1}^{K}\left(e^{a_i \mu_i t + \frac{1}{2}(a_i\sigma_i)^2 t^2}\right) = exp\left(\sum_{i=1}^{K}(a_i\mu_i)t + \frac{1}{2}\sum_{i=1}^{K}((a_i\sigma_i)^2)t^2\right)
\end{aligned}$$

If we define $\mu_y = \sum_{i=1}^{K}(a_i\mu_i)$ and $\sigma_y^2 = \sum_{i=1}^{K}((a_i\sigma_i)^2)$ Then we get

$$M_y(t) = e^{\mu_y t + \frac{1}{2}\sigma_y^2 t^2} \tag{3}$$

This is just the Moment generating function of a Gaussian random variable, Hence Y is a Gaussian random variable. Hence Z is also a Gaussian random variable.

4.4 Task D

To prove that a PDF uniquely determines the MGF, we simply use the construction of the MGF. The MGF is defined as

$$M_X(t) = \mathbb{E}[e^{tX}] = \int_{-\infty}^{\infty} e^{tx} p_X(x)\, dx = \sum_{i=1}^{n} e^{tx_i} p(x_i)$$

where the support of the PDF is $\{x_1, \cdots x_n\}$.

To prove the other direction, let us assume that we can find two distinct solutions $\{x_1, \cdots x_n\}$ and $\{y_1, \cdots y_m\}$ from an arbitrary MGF which we know is of a finite discrete PDF. Then, by the above result, we know that

$$\sum_{i=1}^{n} e^{tx_i} p(x_i) = \sum_{j=1}^{m} e^{ty_j} p(y_j)$$

If any of the x_i's are not equal to any of the y_j's, and given that all $p(x_i)$'s and $p(y_j)$'s are non-zero, then we can rearrange the terms to get that for some a_i's

$$e^{a_0} t b_0 = \sum_{i=1}^{k} e^{a_i} t b_i$$

for some non-zero coefficients b_0 and for some k, and all a_i's being distinct. Now, note that the LHS has an annihilating polynomial $D - a_0$, where D is the derivative operator w.r.t t. The RHS has an annhitating polynomial $(D-a_1)\cdots(D-a_k)$. Since the LHS equals the RHS, if an operator makes the LHS 0, it also makes the RHS zero. Thus, the LHS $(e^{a_0 t}b_0)$ is also annhilated by the annhilator of the RHS.

$$
\begin{aligned}
&(D-a_1)\cdots(D-a_k)(e^{a_0 t}b_0) = 0 \\
\implies &(D-a_1)\cdots(D-a_{k-1})((a_0-a_k)e^{a_0 t}b_0) = 0 \\
\implies &(D-a_1)\cdots(D-a_{k-2})(a_0-a_{k-1})(a_0-a_k)e^{a_0 t}b_0 = 0 \\
&\cdots \\
\implies &(a_0-a_1)\cdots(a_0-a_k)e^{a_0 t}b_0 = 0
\end{aligned}
$$

Since $e^{a_0 t}$ is never 0, and we know $b_0 \neq 0$, $(a_0-a_1)\cdots(a_0-a_k) = 0$. But this is only possible if $a_0 = a_i$ for some $i \in \{1, \cdots, k\}$. But that contradicts our assumption that the a_i's are distinct.

Thus, it is not possible for two PDFs of finite discrete RVs to correspond to the same MGF.