STAT253/317 Winter 2014 Lecture 20

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8.2.2 Steady-State Probabilities
8.5 The System $M/G/1$

8.2.2. Steady-State Probabilities

For a general queueing model, we are interested in three different limiting probabilities:

$$P_n = \lim_{t\to\infty} \mathrm{P}(X(t) = n),$$
where $X(t) = \#$ of customers in the system at time t

$a_n =$ proportion of customers arrive finding n in the system

$d_n =$ proportion of customers depart leaving n behind in the system

Here we assume they exist.
Though the three are defined differently, the latter two are identical in most of the queueing models.

Proposition 8.1 In any system in which customers arrive and depart one at a time

the rate at which arrivals find $n =$ the rate at which departures leave n

and

$$a_n = d_n$$

Proof of Proposition 8.1

Let

$N_{i,j}(t) =$ number of times the number of customers in the system goes from i to j by time t

$A(t) =$ number of customers arrived by time t

$D(t) =$ number of customers departed by time t

Note that an arrival will see n in the system whenever the number in the system goes from n to $n+1$; similarly, a departure will leave behind n whenever the number in the system goes from $n+1$ to n. Thus we know

$$\text{the rate at which arrivals find } n = \lim_{t\to\infty} \frac{N_{n,n+1}(t)}{t}$$

$$\text{the rate at which departures leave } n = \lim_{t\to\infty} \frac{N_{n+1,n}(t)}{t}$$

$$a_n = \lim_{t\to\infty} \frac{N_{n,n+1}(t)}{A(t)}, \quad d_n = \lim_{t\to\infty} \frac{N_{n+1,n}(t)}{D(t)}$$

Proof of Proposition 8.1 (Cont'd)

Since between any two transitions from n to $n+1$, there must be one from $n+1$ to n, and vice versa, we have

$$N_{n,n+1}(t) = N_{n+1,n}(t) \pm 1 \quad \text{for all } t.$$

Thus

$$\begin{aligned}\text{rate at which arrivals find } n &= \lim_{t\to\infty} \frac{N_{n,n+1}(t)}{t} \\ &= \lim_{t\to\infty} \frac{N_{n+1,n}(t) \pm 1}{t} \\ &= \text{rate at which departures leave } n\end{aligned}$$

For a_n and d_n, obviously $A(t) \geq D(t)$ and hence $\lim_{t\to\infty} A(t)/t \geq \lim_{t\to\infty} D(t)/t$. There are two possibilities:

- if $\lim_{t\to\infty} A(t)/t = \lim_{t\to\infty} D(t)/t$, then obviously $a_n = d_n$ for all n
- if $\lim_{t\to\infty} A(t)/t > \lim_{t\to\infty} D(t)/t$, then the queue size will go to infinity, implying that $a_n = d_n = 0$. The equality is still valid.

Example 8.1

Here is an example where $P_n \neq a_n$. Consider a queueing model in which

- service times = 1, always
- interarrival times are always > 1 [e.g., Uniform(1.5,2)].

Hence, as every arrival finds the system empty and every departure leaves it empty, we have

$$a_0 = d_0 = 1$$

However, $P_0 \neq 1$ as the system is not always empty of customers.

Proposition 8.2 (PASTA Principle)

Poisson Arrivals See Time Averages

If the arrival process is Poisson, then

$$P_n = a_n,$$

and hence $P_n = d_n$.

Why is PASTA True?

- By time T, the total amount of time there are n customers in the system is about $P_n T$
- Regardless of how many customers in the system, Poisson arrivals always arrive at rate λ. Thus by time T, the total number of arrivals that find n in the system is $\approx \lambda P_n T$.
- the overall number of customers arrived by time T is $\approx \lambda T$
- the proportion of arrivals that find the system in state n is

$$a_n = \frac{\lambda P_n T}{\lambda T} = P_n$$

$M/G/1$

The $M/G/1$ model assumes

- Poisson arrivals at rate λ;
- i.i.d service times with a general distribution G, $S_i \sim G$;
- a single server; and
- first come, first serve

A necessary condition for an $M/G/1$ to be stable is that the mean of service time $\mathbb{E}[S_n]$ must satisfies

$$\lambda \mathbb{E}[S_n] < 1.$$

This condition is necessary. Otherwise if

$$\begin{aligned} &\text{the average service time } \mathbb{E}[S_n] \\ &> \text{ the average interarrival time of customers } 1/\lambda, \end{aligned}$$

the queue will become longer and longer and the system will ultimately explode.

A Markov Chain embedded in $M/G/1$

Let $X(t) = \#$ of customers in the system at time t.
Unlike $M/M/k$ or $M/M/\infty$ systems, the process $\{X(t), t \geq 0\}$ in a $M/G/1$ system is NOT a continuous time Markov chain, because the service time is not memoryless. Future events will depend on the current service time.
However, there is a discrete-time Markov chain embedded in an $M/G/1$ system. Let

$$\begin{aligned} Y_0 &= 0 \\ Y_n &= \# \text{ of customers in the system} \\ &\qquad \text{leaving behind at the } n\text{th departure, } n \geq 1 \\ A_n &= \# \text{ of customers that enter the system} \\ &\qquad \text{during the service time of the } n\text{th customer, } n \geq 1 \end{aligned}$$

Observed that $\{Y_n, n \geq 0\}$ and $\{A_n, n \geq 1\}$ are related as follows

$$Y_{n+1} = A_{n+1} + (Y_n - 1)^+ = \begin{cases} Y_n - 1 + A_{n+1} & \text{if } Y_n > 0 \\ A_{n+1} & \text{if } Y_n = 0 \end{cases}$$

A Markov Chain embedded in $M/G/1$ (Cont'd)

Recall that S_n denotes the length of time to serve the nth customer.

Given S_n, A_n is Poisson with mean λS_n. From this we can conclude that $A_1, A_2, \ldots$ are i.i.d. since

- the service times $S_1, S_2, \ldots$ are i.i.d., and
- there is only 1 server, the service times of different customers are disjoint, and the number of events occurred in disjoint intervals are independent in a Poisson process.

That $\{A_n, n \geq 1\}$ are i.i.d. and Y_n is independent of A_{n+1} implies that

$$\boxed{\{Y_n, n \geq 0\} \text{ is a Markov chain.}}$$

Recall we have seen this Markov chain in Lecture 1 and in HW7.

A Markov Chain Embedded in $M/G/1$ (Cont'd)

Moreover, as A_n given S_n is Poisson with mean λS_n, we can find the distribution of A_n

$$\begin{aligned} \alpha_k = \mathrm{P}(A_n = k) &= \int_0^\infty \mathrm{P}(A_n = k | S_n = y) G(dy) \\ &= \int_0^\infty \frac{(\lambda y)^k}{k!} e^{-\lambda y} G(dy) \end{aligned}$$

from which we can find the transition probability P_{ij} for the Markov chain $\{Y_n, n \geq 0\}$:

$$\begin{aligned} P_{ij} &= \mathrm{P}(Y_{n+1} = j | Y_n = i) = \mathrm{P}(A_{n+1} = j - (i-1)^+) \\ &= \begin{cases} \alpha_j, & \text{if } i = 0 \\ \alpha_{j-i+1}, & \text{if } i \geq 1, j \geq i-1 \\ 0 & \text{if } i \geq 1, j < i-1 \end{cases} \end{aligned}$$

We can show that the Markov chain is irreducible and aperiodic and has a limiting distribution if and only if $\lambda \mathbb{E}[S_1] < 1$.

Idle Periods in $M/G/1$

Using the equation $Y_{n+1} = A_{n+1} + (Y_n - 1)^+$, we can find many properties of the Markov chain. First write the equation as

$$Y_{n+1} = A_{n+1} + Y_n - 1 + \mathbf{1}_{\{Y_n = 0\}}$$

Taking expectations we get

$$\begin{aligned} \mathbb{E}[Y_{n+1}] &= \mathbb{E}[A_{n+1}] + \mathbb{E}[Y_n] - 1 + \mathrm{P}(Y_n = 0) \\ &= \lambda \mathbb{E}[S] + \mathbb{E}[Y_n] - 1 + \mathrm{P}(Y_n = 0) \end{aligned}$$

where the last equality comes from that A_n given S_n is Poisson with mean λS_n, and hence $\mathbb{E}[A_n] = \lambda \mathbb{E}[S_n] = \lambda \mathbb{E}[S]$ since S_i's are i.i.d.

Let $n \to \infty$, since the MC has a limiting distribution, we have $\lim_{n\to\infty} \mathbb{E}[Y_{n+1}] = \lim_{n\to\infty} \mathbb{E}[Y_n]$ and from which we can get

$$\lim_{n\to\infty} \mathrm{P}(Y_n = 0) = 1 - \lambda \mathbb{E}[S]$$

By the PASTA principle, $\lim_{n\to\infty} \mathrm{P}(Y_n = 0) = d_0 = P_0$ is also the long-run proportion of time that the system is idle.

Length of Busy Periods in $M/G/1$

As in a birth & death queueing model, there is a alternating renewal process embedded in an $M/G/1$ system. We say a renewal occurs if the system become empty, then the system idles for a period of time until the next customer enters the system, and then a busy period begins until the system become empty again.
Using the alternating renewal theory, the long-run proportion of time that the system is empty is

$$\frac{\mathbb{E}[\text{Idle}]}{\mathbb{E}[\text{Idle}] + \mathbb{E}[\text{Busy}]},$$

and we just derived that it is $\lim_{t\to\infty} \mathrm{P}(X(t) = 0) = 1 - \lambda \mathbb{E}[S]$.
Since the length of an idle period $\sim Exp(\lambda)$, we have $\mathbb{E}[\text{Idle}] = 1/\lambda$. In summary, we have that

$$1 - \lambda \mathbb{E}[S] = \frac{1/\lambda}{(1/\lambda) + \mathbb{E}[\text{Busy}]} \quad \Rightarrow \quad \mathbb{E}[\text{Busy}] = \frac{\mathbb{E}[S]}{1 - \lambda \mathbb{E}[S]}$$

L of $M/G/1$ (Cont'd)

From the equation $Y_{n+1} = A_n - 1 + Y_n + \mathbf{1}_{\{Y_n=0\}}$, we have

$$\begin{aligned}
&\mathrm{Var}(Y_{n+1})\\
&= \mathrm{Var}(A_{n+1} - 1 + Y_n + \mathbf{1}_{\{Y_n=0\}})\\
&= \mathrm{Var}(A_{n+1}) + \mathrm{Var}(Y_n + \mathbf{1}_{\{Y_n=0\}}) \qquad (A_{n+1} \text{ and } Y_n \text{ are indep.})\\
&= \mathrm{Var}(A_{n+1}) + \mathrm{Var}(Y_n)\\
&\qquad + 2\mathrm{Cov}(Y_n, \mathbf{1}_{\{Y_n=0\}}) + \mathrm{Var}(\mathbf{1}_{\{Y_n=0\}}), \qquad (1)
\end{aligned}$$

in which

$$\mathrm{Var}(\mathbf{1}_{\{Y_n=0\}}) = \mathrm{P}(Y_n = 0)(1 - \mathrm{P}(Y_n = 0)) \qquad (2)$$

$$\begin{aligned}
\mathrm{Cov}(Y_n, \mathbf{1}_{\{Y_n=0\}}) &= \mathbb{E}[\underbrace{Y_n \mathbf{1}_{\{Y_n=0\}}}_{=0}] - \mathbb{E}[Y_n]\mathrm{P}(Y_n = 0)\\
&= -\mathbb{E}[Y_n]\mathrm{P}(Y_n = 0) \qquad (3)\\
\mathrm{Var}(A_n) &= \mathbb{E}[\mathrm{Var}(A_n|S_n)] + \mathrm{Var}(\mathbb{E}[A_n|S_n])\\
&= \mathbb{E}[\lambda S_n] + \mathrm{Var}(\lambda S_n)\\
&= \lambda\mathbb{E}[S] + \lambda^2\mathrm{Var}(S) \qquad (4)
\end{aligned}$$

L of $M/G/1$ (Cont'd)

Plugging in (2) (3) (4) into (1), letting $n \to \infty$, we have

$$\begin{aligned}
\lim_{n\to\infty} \mathrm{Var}(Y_{n+1}) &= \lambda\mathbb{E}[S] + \lambda^2\mathrm{Var}(S) + \lim_{n\to\infty} \mathrm{Var}(Y_n)\\
&\quad - 2\lim_{n\to\infty} \mathbb{E}[Y_n]\mathrm{P}(Y_n = 0)\\
&\quad + \lim_{n\to\infty} \mathrm{P}(Y_n = 0)(1 - \mathrm{P}(Y_n = 0))\\
&= \lambda\mathbb{E}[S] + \lambda^2\mathrm{Var}(S) + \lim_{n\to\infty} \mathrm{Var}(Y_n)\\
&\quad - 2\lim_{n\to\infty} \mathbb{E}[Y_n](1 - \lambda\mathbb{E}[S]) + (1 - \lambda\mathbb{E}[S])\lambda\mathbb{E}[S]
\end{aligned}$$

Again since the MC has a limiting distribution, we have $\lim_{n\to\infty} \mathrm{Var}[Y_{n+1}] = \lim_{n\to\infty} \mathrm{Var}[Y_n]$, and can get

$$\begin{aligned}
\lim_{n\to\infty} \mathbb{E}[Y_n] &= \frac{\lambda\mathbb{E}[S] + \lambda^2\mathrm{Var}(S)}{2(1 - \lambda\mathbb{E}[S])} + \frac{\lambda\mathbb{E}[S]}{2}\\
&= \frac{\lambda^2\mathbb{E}[S^2]}{2(1 - \lambda\mathbb{E}[S])} + \lambda\mathbb{E}[S] \quad (\text{since } \mathrm{Var}(S) = \mathbb{E}[S^2] - (\mathbb{E}[S])^2)
\end{aligned}$$

L of $M/G/1$ (Cont'd)

By the PASTA principle, we know $\lim_{n\to\infty} \mathbb{E}[Y_n] = \lim_{n\to\infty} \mathbb{E}[X(t)] = L$. From the cost identity $L = \lambda_a W$ and $L_Q = \lambda_a W_Q$, and that $\lambda_a = \lambda$, we have

$$L = \frac{\lambda^2\mathbb{E}[S^2]}{2(1 - \lambda\mathbb{E}[S])} + \lambda\mathbb{E}[S]$$

$$W = L/\lambda = \frac{\lambda\mathbb{E}[S^2]}{2(1 - \lambda\mathbb{E}[S])} + \mathbb{E}[S]$$

$$W_Q = W - \mathbb{E}[S] = \frac{\lambda\mathbb{E}[S^2]}{2(1 - \lambda\mathbb{E}[S])}$$

$$L_Q = \lambda W_Q = \frac{\lambda^2\mathbb{E}[S^2]}{2(1 - \lambda\mathbb{E}[S])}$$

Since $\mathbb{E}[S^2] = (\mathbb{E}[S])^2 + \mathrm{Var}(S)$, from the equations above we see for fixed mean service time $\mathbb{E}[S]$,

L, L_Q, W, and W_Q all increase as $\mathrm{Var}(S)$ increases.

Example

For an $M/M/1$ system, we have shown that if the service time is exponential with mean $1/\mu$ that the average waiting time is

$$W = \frac{1}{\mu - \lambda}$$

If the service time is exactly $1/\mu$, the average waiting time can be reduced to

$$W = \frac{\lambda\mathbb{E}[S^2]}{2(1 - \lambda\mathbb{E}[S])} + \mathbb{E}[S] = \frac{\lambda/\mu^2}{2(1 - \lambda/\mu)} + 1/\mu = \frac{1}{\mu - \lambda} - \frac{\lambda/\mu}{2(\mu - \lambda)}$$

For example, for $\lambda = 1/12$, $\mu = 1/8$

$$W = \begin{cases} 24 & \text{for } M/M/1\\ 16 & \text{if service time is exactly } 1/\mu = 8 \end{cases}$$

For $\lambda = 1/10$, $\mu = 1/8$

$$W = \begin{cases} 40 & \text{for } M/M/1\\ 24 & \text{if service time is exactly } 1/\mu = 8 \end{cases}$$