Single-server queue with Markov dependent inter-arrival and service times

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Abstract

In this paper we study a single-server queue where the inter-arrival times and the service times depend on a common discrete time Markov Chain. This model generalizes the well-known $MAP/G/1$ queue by allowing dependencies between inter-arrival and service times. The waiting time process is directly analyzed by solving Lindley's equation by transform methods. The Laplace Stieltjes transforms (LST) of the steady-state waiting time and queue length distribution are both derived, and used to obtain recursive equations for the calculation of the moments. Numerical examples are included to demonstrate the effect of the auto-correlation of and the cross-correlation between the inter-arrival and service times.

1 Introduction

In the literature much attention has been devoted to single-server queues with Markovian Arrival Processes (MAP), see, e.g., [27] and the references therein. The $MAP/G/1$ queue provides a powerful framework to model dependences between successive inter-arrival times, but typically the service times are iid and independent of the arrival process. The present study concerns single-server queues where the inter-arrival times and the service times depend on a common discrete time Markov Chain; i.e., the so-called semi-Markov queues. As such the model under consideration is a generalization of the $MAP/G/1$ queue, by also allowing dependencies between successive service times and between inter-arrival times and service times.

The phenomenon of dependence among the inter-arrival times in the packet streams of voice and data traffic is well known; see, e.g., [19, 20, 33]. But [16] argue that in packet

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communication networks one should also expect two additional forms of dependence: between successive service times and among inter-arrival times and service times. These forms of dependence occur because of the presence of bursty arrivals and multiple sources with different mean service times (due to different packet lengths), and they may have a dominant effect on waiting times and queue lengths.

In [27] the MAP construction is generalized to the Batch Markovian Arrival Process (BMAP) to allow batch arrivals. The framework of the $BMAP/G/1$ queue can also be used to model dependence between inter-arrival times and service times. This is described in [11]. The important observation is that the arrival of a batch (the size of which depends on the state of the underlying Markov chain) can be viewed as the arrival of a super customer, whose service time is distributed as the sum of the service requests of the customers in the batch.

Models with dependencies between inter-arrival and service time have been studied by several authors. A review of the early literature can be found in [3]. The queueing model in [8] assumes dependence between service requests and the subsequent inter-arrival time; see also [7]. Models with a linear dependence between the service time and the preceding inter-arrival time have been studied in [9, 12, 13]; other papers [14, 17, 18, 29] analyze the $M/M/1$ queue where the service time and the preceding inter-arrival time have a bivariate exponential density with a positive correlation. The linear and bivariate exponential cases are both contained in the correlated $M/G/1$ queue studied by [4, 5, 6]. The correlation structure considered in [4, 5, 6] arises in the following framework: customers arrive according to a Poisson stream at some collection point, where after exponential time periods, they are collected in a batch and transported to a service system. In [11] it is shown that the collect system can also be modelled by using the BMAP framework.

A special case of Markov dependent inter-arrival and service times is the model with strictly periodic arrivals. These models arise, for example, in the modelling of inventory systems using periodic ordering policies; see [36, 37]. Queueing models with periodic arrival processes have been studied in, e.g., [30, 31, 23, 10].

The starting point of the analysis is Lindley's equation for the waiting times. We solve this equation using the transform techniques and derive the LST of the steady-state waiting time distribution. By exploiting a well-known relation between the waiting time of a customer and the number of customers left behind by a departing customer we find the LST of the queue length distribution at departure epochs and at arbitrary time points. The LST's are then used to derive simple recursive equations for the moments of the waiting time and queue length.

The paper is organized as follows. In Section 2 we present the queueing model under consideration. The waiting time process is analyzed in Section 3. We first derive the LST of the steady-state waiting time distribution in subsection 3.1. This transform is used in subsection 3.2 to obtain a system of recursive equations for the moments of the waiting time. In the subsequent section we study the queue length process. Some numerical examples are presented in Section 5 to demonstrate the effect of correlated inter-arrival and service times.

2 Queueing model

We consider a single-server queue, where customers are served in order of arrival. Let τ_n be the time of the nth arrival to the system ($n \geq 0$) with $\tau_0 = 0$. Define $A_n = \tau_n - \tau_{n-1}$, ($n \geq 1$). Thus A_n is the time between the nth and $(n-1)$th arrival. Let S_n ($n \geq 0$) be the service time of the nth arrival. We assume that the sequences $\{A_n, n \geq 1\}$ and $\{S_n, n \geq 0\}$ are auto-correlated as well as cross-correlated. The nature of this dependence is described below.

The inter-arrival and service times are regulated by an irreducible discrete-time Markov chain $\{Z_n, n \geq 0\}$ with state space $\{1, 2, \ldots, N\}$ and transition probability matrix P. More precisely, the tri-variate process $\{(A_{n+1}, S_n, Z_n), n \geq 0\}$ has the following probabilistic structure:

$$\begin{aligned} P(A_{n+1} \leq x, S_n \leq y, Z_{n+1} = j \mid Z_n = i, (A_{r+1}, S_r, Z_r), 0 \leq r \leq n-1) \\ = P(A_1 \leq x, S_0 \leq y, Z_1 = j \mid Z_0 = i) \\ = G_i(y) p_{i,j} (1 - e^{-\lambda_j x}), \end{aligned} \tag{1}$$

$x, y \geq 0; i, j = 1, 2, \ldots, N$. Thus A_{n+1}, S_n and Z_{n+1} are independent of the past, given Z_n. Further, A_{n+1} and S_n are conditionally independent, given Z_n and Z_{n+1}, where S_n has an arbitrary distribution and A_{n+1} has an exponential distribution (but the model can be easily extended to phase-type distributions; see Remark 2.3).

This model fits in the general class of semi-Markov queues that only assume the Markov property (1). Semi-Markov queues have been extensively studied in the literature; see, e.g., [35] and the references therein. However, explicit numerically tractable results have been obtained for special cases only. The model with arbitrary inter-arrival times and phase-type service times has been studied in [32], where it is shown that the waiting time distribution is matrix-exponential, with a phase-type representation.

Let γ_i be the mean and s_i^2 be the second moment of the service time distribution G_i, and $\pi = [\pi_1, \pi_2, \cdots, \pi_N]$ be the stationary distribution of $\{Z_n, n \geq 0\}$. Then the system is stable if (see [1, 24])

$$\sum_{i=1}^{N} \pi_i \gamma_i < \sum_{i=1}^{N} \pi_i \lambda_i^{-1},$$

since the left-hand side is the mean service time of a customer, and the right-hand side is the mean inter-arrival time between two consecutive customers in steady state. Let

$\Gamma = \text{diag}(\gamma_1, \ldots, \gamma_N)$, $\Lambda = \text{diag}(\lambda_1, \ldots, \lambda_N)$, and e be the column vector of ones. Then we can write the above stability condition in a matrix form as follows:

$$\pi(\Lambda^{-1} - \Gamma)e > 0. \tag{2}$$

Remark 2.1 In steady-state the auto-correlation between S_m and S_{m+n} is given by

$$\rho(S_m, S_{m+n}) = \rho(S_0, S_n) = \frac{\sum_{i=1}^{N} \sum_{j=1}^{N} \pi_i [p_{ij}^{(n)} - \pi_j]\gamma_i \gamma_j}{\sum_{i=1}^{N} \pi_i s_i^2 - (\sum_{i=1}^{N} \pi_i \gamma_i)^2}, \quad n \geq 1,$$

where

$$p_{ij}^{(n)} = P(Z_n = j \mid Z_0 = i), \quad 1 \leq i, j \leq N, \quad n \geq 0.$$

A similar expression holds for the auto-correlation between the inter-arrival times. Provided P is aperiodic, $p_{ij}^{(n)}$ converges to π_j geometrically as n tends to infinity. Hence, the auto-correlation function approaches zero geometrically as the lag goes to infinity.

Remark 2.2 In steady-state the cross-correlation between A_n and S_n is given by

$$\rho(A_n, S_n) = \rho(A_1, S_1) = \frac{\sum_{i=1}^{N} \pi_i (\lambda_i^{-1} - \lambda^{-1})(\gamma_i - \gamma)}{\left\{\sum_{i=1}^{N} \pi_i (\lambda_i^{-1} - \lambda^{-1})^2 \sum_{i=1}^{N} \pi_i (\gamma_i - \gamma)^2\right\}^{1/2}},$$

where $\lambda^{-1} = \sum_{i=1}^{N} \pi_i \lambda_i^{-1}$ and $\gamma = \sum_{i=1}^{N} \pi_i \gamma_i$.

Remark 2.3 The arrival process is a Markovian Arrival Process (MAP) (see [26]), where each transition of Z corresponds to an arrival. Transitions without arrivals can be easily included by allowing customers with zero service times. This device enables us to model general phase-type inter-arrival times between (real) customers with non-zero service times. clearly, the calculation of the auto-correlations and cross-correlations of inter-arrival times and service times of customers with non-zero service times becomes more complicated than explained in the previous remarks. An important property from a modelling point of view is that this class of arrival processes is dense in the class of marked point processes; see [2].

Remark 2.4 If the distribution of S_n does not depend on Z_n, i.e., $G_i(\cdot) = G(\cdot)$ for all i, then the model reduces to the $MAP/G/1$ queue, which is a special case of the $BMAP/G/1$ queue (see [27]).

Remark 2.5 Periodic arrivals can be derived as a special case of the present model by setting $p_{i,i+1} = p_{N,1} = 1$ for $i = 1, 2, \ldots, N-1$. The waiting time process for the periodic model has also been studied in [10], where functional equations for the stationary distributions of waiting times are derived. These equations formulate a Hilbert Boundary Value problem, which can be solved if the LST's of all the inter-arrival time distributions or the LST's of all the service time distributions are rational.

3 The waiting time process

In this section we study the LST of the limiting distribution of the waiting time, and its moments.

3.1 Steady state LST

Let W_n denote the waiting time of the nth customer. Let 1_A be the indicator random variable of the event A. Using the notation $E(X;A)$ to mean $E(X \cdot 1_A)$ for any event A, define

$$\phi_i^n(s) = E(e^{-sW_n}; Z_n = i), \qquad \mathrm{Re}(s) \geq 0, n \geq 0, i = 1, 2, \ldots, N.$$

and, assuming the limit exists, define

$$\phi_i(s) = \lim_{n\to\infty} \phi_i^n(s), \quad i = 1, 2, \ldots, N.$$

The next theorem gives the equations satisfied by the transforms

$$\phi(s) = [\phi_1(s),\ \phi_2(s),\ \cdots,\ \phi_N(s)].$$

First we need the following notation:

$$\begin{aligned} \tilde{G}_i(s) &= \int_0^\infty e^{-st} dG_i(t), \qquad 1 \leq i \leq N, \\ \tilde{G}(s) &= \mathrm{diag}(\tilde{G}_1(s),\ \tilde{G}_2(s),\ \cdots,\ \tilde{G}_N(s)), \\ H_{i,j}(s) &= \tilde{G}_i(s) p_{i,j} \lambda_j, \qquad 1 \leq i, j \leq N, \\ H(s) &= [H_{i,j}(s)] = \tilde{G}(s) P \Lambda. \end{aligned} \tag{3}$$

Theorem 3.1 *Provided condition (2) is satisfied, the transform vector $\phi(s)$ satisfies*

$$\phi(s)[H(s) + sI - \Lambda] = sv, \tag{4}$$

$$\phi(0)e = 1, \tag{5}$$

where $v = [v_1,\ v_2,\ \cdots,\ v_N]$ is given by

$$v_j = \sum_{i=1}^{N} \phi_i(\lambda_j) \tilde{G}_i(\lambda_j) p_{i,j}.$$

Proof: Let T_n denote the sojourn time of the nth customer, i.e., $T_n = W_n + S_n$, $n \geq 0$. The waiting times W_n satisfy Lindley's equation (see [22]),

$$W_{n+1} = (W_n + S_n - A_{n+1})^+ = (T_n - A_{n+1})^+, \qquad n \geq 0,$$

where $(x)^+ = \max\{x, 0\}$. From Lindley's equation we obtain the following equation for the transforms $\phi_j^{n+1}(s)$, $j = 1, \dots, N$,

$$\begin{aligned}
\phi_j^{n+1}(s) &= E(e^{-sW_{n+1}}; Z_{n+1} = j) \\
&= \sum_{i=1}^{N} P(Z_n = i) E(e^{-sW_{n+1}}; Z_{n+1} = j | Z_n = i) \\
&= \sum_{i=1}^{N} P(Z_n = i) p_{i,j} E(e^{-s(T_n - A_{n+1})^+} | Z_n = i, Z_{n+1} = j) \\
&= \sum_{i=1}^{N} P(Z_n = i) p_{i,j} \Bigg[E(\int_{x=0}^{T_n} e^{-s(T_n - x)} \lambda_j e^{-\lambda_j x} dx | Z_n = i) \\
&\qquad\qquad + E(\int_{x=T_n}^{\infty} e^{-s \cdot 0} \lambda_j e^{-\lambda_j x} dx | Z_n = i) \Bigg] \\
&= \sum_{i=1}^{N} P(Z_n = i) p_{i,j} E(e^{-sT_n} \frac{\lambda_j}{\lambda_j - s} (1 - e^{-(\lambda_j - s)T_n}) + e^{-\lambda_j T_n} | Z_n = i) \\
&= \sum_{i=1}^{N} p_{i,j} \left[\frac{\lambda_j}{\lambda_j - s} \phi_i^n(s) \tilde{G}_i(s) - \frac{s}{\lambda_j - s} \phi_i^n(\lambda_j) \tilde{G}_i(\lambda_j) \right], \qquad (6)
\end{aligned}$$

It is clear that $\phi_j^n(s)$ tends to $\phi_i(s)$ if the stability condition (2) is satisfied. Hence, letting $n \to \infty$ in (6), and rearranging, we get

$$\sum_{i=1}^{N} p_{i,j} \lambda_j \phi_i(s) \tilde{G}_i(s) - (\lambda_j - s) \phi_j(s) = s \sum_{i=1}^{N} p_{i,j} \phi_i(\lambda_j) \tilde{G}_i(\lambda_j), \qquad j = 1, \dots, N. \qquad (7)$$

Note that the sum on the right-hand side of the above equation is denoted by v_j. Then we can rewrite (7) in matrix form yielding (4). Equation (5) is just the normalization equation. This completes the proof. □

Clearly, we need to determine the unknown vector v in equation (4). For that we need to study the solutions of

$$\det(H(s) + sI - \Lambda) = 0. \qquad (8)$$

The following theorem gives the number and placement of the solutions to the above equation (cf. Theorem 5 in [25]).

Theorem 3.2 *Equation (8) has exactly N solutions s_i, $1 \le i \le N$, with $s_1 = 0$ and $Re(s_i) > 0$ for $2 \le i \le N$.*

Proof: See Appendix.

With the above result we now give a method of determining the vector v in the following theorem.

Theorem 3.3 *Suppose the condition of stability (2) is satisfied, and the $N-1$ solutions s_i, $2 \leq i \leq N$ with $Re(s_i) > 0$ to equation (4) are distinct. Let a_i be a non-zero column vector satisfying*

$$(H(s_i) + s_i I - \Lambda)a_i = 0, \quad 2 \leq i \leq N.$$

Then v is given by the unique solution to the following N linear equations:

$$va_i = 0, \quad 2 \leq i \leq N, \tag{9}$$

$$v\Lambda^{-1}e = \pi(\Lambda^{-1} - \Gamma)e. \tag{10}$$

Proof: Since s_i satisfies equation (8), it follows that there is a non-zero column vector a_i such that

$$(H(s_i) + s_i I - \Lambda)a_i = 0, \qquad i = 1, 2, \ldots, N.$$

In particular $a_1 = e$, the column vector of ones. Post-multiplying equation (4) with $s = s_i$ by a_i, we get

$$\phi(s_i)[H(s_i) + s_i I - \Lambda]a_i = s_i v a_i = 0, \qquad i = 1, 2, \ldots, N.$$

Since $s_i \neq 0$, for $2 \leq i \leq N$, v must satisfy equation (9). To derive the remaining equation, we take the derivative of equation (4) with respect to s, yielding

$$\phi(s)[H'(s) + I] + \phi'(s)(H(s) + sI - \Lambda) = v.$$

Setting $s = 0$ we get

$$\phi(0)(H'(0) + I) + \phi'(0)(P - I)\Lambda = v.$$

Post-multiplying by $\Lambda^{-1}e$ gives

$$\phi(0)(H'(0) + I)\Lambda^{-1}e + \phi'(0)(P - I)\Lambda\Lambda^{-1}e = b\Lambda^{-1}e.$$

Using $(P - I)e = 0$, $H'(0) = -\text{diag}(\gamma)P\lambda$ and $\phi(0) = \pi$ (where the latter follows from (4) with $s = 0$ and the normalization equation (5)), the above can be simplified to

$$\pi(\Lambda^{-1} - \Gamma)e = v\Lambda^{-1}e.$$

The uniqueness of the solution follows from the general theory of Markov chains that under the condition of stability, there is a unique stationary distribution and thus also a unique solution $\phi(s)$ to the equations (4) and (5). This completes the proof. □

3.2 Steady state moments

Once v is known, the entire transform vector $\phi(s)$ is known. We can use it to compute the moments of the waiting time in steady state. They are given in the following theorem. First some notation:

$$\begin{aligned}
m_{r,i} &= \lim_{n\to\infty} E(W_n^r; Z_n = i), \qquad r = 0, 1, 2, \ldots,\ 1 \le i \le N, \\
m_r &= [m_{r,1},\ m_{r,2},\ \cdots,\ m_{r,N}], \qquad r = 0, 1, 2, \ldots, \\
\gamma_{r,i} &= \int_0^\infty x^r dG_i(x), \qquad r = 0, 1, 2, \ldots,\ 1 \le i \le N, \\
\Gamma_r &= \mathrm{diag}[\gamma_{r,1},\ \gamma_{r,2},\ \cdots,\ \gamma_{r,N}], \qquad r = 0, 1, 2, \ldots
\end{aligned}$$

Note that $\gamma_{1,i} = \gamma_i$ and $\Gamma_1 = \Gamma$. We assume that the above moments exist.

Theorem 3.4 *The moment-vectors m_r satisfy the following recursive equations:*

$$m_0 = \pi, \tag{11}$$

$$m_1(I - P) = m_0(\Gamma_1 P - \Lambda^{-1}) + v\Lambda^{-1}, \tag{12}$$

$$m_1(\Lambda^{-1} - \Gamma_1)e = \frac{1}{2} m_0 \Gamma_2 e, \tag{13}$$

$$m_r(I - P) = r m_{r-1}(\Gamma_1 P - \Lambda^{-1}) + \sum_{k=2}^{r} \binom{r}{k} m_{r-k}\Gamma_k P, \tag{14}$$

$$m_r(\Lambda^{-1}e - \Gamma_1 e) = \frac{1}{r+1}\left[\sum_{k=2}^{r+1}\binom{r+1}{k} m_{r+1-k}\Gamma_k e\right]. \tag{15}$$

Proof: We have

$$\begin{aligned}
\phi(s) &= \sum_{r=0}^{\infty}(-1)^r m_r \frac{s^r}{r!}, \\
\tilde{G}(s) &= \sum_{r=0}^{\infty}(-1)^r \Gamma_r \frac{s^r}{r!}.
\end{aligned}$$

¿From equation (3) it follows that

$$H(s) = \sum_{r=0}^{\infty}(-1)^r \Gamma_r P \Lambda \frac{s^r}{r!}.$$

Substituting in equation (4) we get

$$\sum_{r=0}^{\infty}(-1)^r m_r \frac{s^r}{r!}\left(\sum_{r=0}^{\infty}(-1)^r \Gamma_r P\Lambda \frac{s^r}{r!} + sI - \Lambda\right) = sv.$$

Equating the coefficients of s^0 we get:

$$m_0(\Gamma_0 P\Lambda - \Lambda) = 0.$$

Since $\Gamma_0 = I$, and Λ is invertible this simplifies to

$$m_0(I - P) = 0.$$

We also know that $m_0 e = 1$. But these equations have a unique solution π, and hence we get equation (11). Next, equating the coefficients of s^1 and simplifying, we get (12). Multiplying this equation by e yields equation (10). Now $(I - P)$ is non-invertible with rank $N-1$, hence we need an additional equation. We get that by equating the coefficients of s^2, which, after simplification, yields

$$m_2(I - P) = m_0\Gamma_2 P + 2m_1(\Gamma_1 P - \Lambda^{-1}).$$

Multiplying by e, we get equation (13). This gives the required additional equation for m_1. Equations (12) and (13) uniquely determine m_1. Note that m_1 depends in Γ_2, the diagonal matrix of second moments of the service times, as expected. Proceeding in this fashion, equating coefficients of s^r, $r \geq 2$, we get equation (14). Using the same equation for $r + 1$ and multiplying it by e yields equation (15). Equations (14) and (15) uniquely determine m_r. This completes the proof. □

Remark 3.5 It is not essential to assume the existence of all the moments of all the service times. If k is the first integer for which at least one entry of Γ_k becomes infinite, the above equations can still be used to determine m_r for $0 \leq r \leq k - 1$.

4 Queue Length Distribution

Now we will analyze the queue length distribution in the queueing system described in Section 2.

4.1 Steady state LST

In this subsection we study the LST of the queue length distribution at departures and at arbitrary times in steady state. Toward this end, first define $N(t)$ to be the number of arrivals up to time t, with $N(0) = 0$, and the continuous time Markov chain $\{Z(t), t \geq 0\}$ to be

$$Z(t) = Z_{N(t)+1}, \qquad t \geq 0.$$

We are interested in

$$\psi_{i,j}(z, t) = E(z^{N(t)}; Z(t) = j \mid Z(0) = i), \qquad t \geq 0,\ 1 \leq i, j \leq N.$$

Let

$$\psi(z, t) = [\psi_{i,j}(z, t)]$$

be the matrix of the generating functions. The following theorem gives a method of computing $\psi(z, t)$. First we need some notation. Let $-\mu_i(z),\ 1 \leq i \leq N$, be the N eigenvalues of $\Lambda Pz - \Lambda$, assumed to be distinct, and let $y_i(z)$ and $x_i(z)$ be the orthonormal left and right eigenvectors corresponding to $-\mu_i(z)$. The matrices $A_i(z)$ are defined as

$$A_i(z) = x_i(z) y_i(z), \qquad 1 \le i \le N.$$

Then we have the following theorem.

Theorem 4.1 *The generating function matrix $\psi(z,t)$ is given by*

$$\psi(z,t) = e^{(\Lambda Pz - \Lambda)t} = \sum_{i=1}^{N} e^{-\mu_i(z)t} A_i(z).$$

Proof: It is easy to show that ψ satisfies the following differential equation (cf. Section 2.1 in [27]):

$$\frac{\partial}{\partial t}\psi(z,t) = (\Lambda Pz - \Lambda)\psi(z,t), \qquad t \ge 0, \tag{16}$$

with the initial condition $\psi(z,0) = I$. Hence the solution is given by

$$\psi(z,t) = e^{(\Lambda Pz - \Lambda)t}.$$

This proves the first equality. The second one holds, since we can write (see, e.g., [21])

$$(\Lambda Pz - \Lambda)^n = \sum_{i=1}^{N} (-\mu_i(z))^n A_i(z).$$

This completes the proof of the theorem. □

Using the above theorem we first derive the generating function of the queue length seen by departures in steady state. Let the random variable L^d denote the queue length and Z^d the state of $\{Z(t), t \ge 0\}$ seen by a departure in steady state. We use the following notation:

$$\begin{aligned} g_j(z) &= E(z^{L^d}; Z^d = j), \qquad 1 \le j \le N, \\ g(z) &= (g_1(z), \ldots, g_N(z)), \\ \theta_{i,j}(z) &= \phi_i(z)\tilde{G}_i(z)p_{i,j}, \qquad 1 \le i,j,k \le N, \\ \Theta(z) &= [\theta_{i,j}(z)]. \end{aligned}$$

The result is given in the next theorem:

Theorem 4.2 *The generating function $g(z)$ is given by*

$$g(z) = \sum_{i=1}^{N} e'\Theta(\mu_i(z))A_i(z). \tag{17}$$

Proof: Suppose that the process $\{(A_{n+1}, S_n, Z_n, W_n), n \geq 0\}$ is in steady state. Then the zeroeth customer finds the system in steady state, and finds $Z_0 = i$ with probability π_i. Given that he finds $Z_0 = i$, his service time S_0 has distribution $G_i(\cdot)$, and the LST of his sojourn time $T_0 = S_0 + W_0$ is given by $\phi_i(s)\tilde{G}_i(s)/\pi_i$. Upon arrival of the zeroeth customer the process $\{Z_n, n \geq 0\}$ jumps from $Z_0 = i$ to state $Z_1 = k$ with probability $p_{i,k}$; hence, $Z(0) = Z_1 = k$ with probability $p_{i,k}$. The new customers that arrive during the sojourn time of this customer are exactly the ones that are left behind by the zeroeth customer, so their number is equal to L^d (cf. [15]). Thus, conditional upon $Z_0 = i$ and $T_0 = t$, we have

$$E(z^{L^d}; Z^d = j \mid Z_0 = i; T_0 = t) = \sum_{k=1}^{N} p_{i,k}\psi_{k,j}(z,t)$$

Hence we obtain

$$\begin{aligned} g_j(z) &= \sum_{i=1}^{N} \int_0^\infty E(z^{L^d}; Z^d = j \mid Z_0 = i; T_0 = t) dP(T_0 \leq t; Z_0 = i) \\ &= \sum_{i=1}^{N} \int_0^\infty dP(T_0 \leq t; Z_0 = i) \sum_{k=1}^{N} p_{i,k}\psi_{k,j}(z,t) \\ &= \sum_{i=1}^{N} \int_0^\infty dP(T_0 \leq t; Z_0 = i) \sum_{k=1}^{N} p_{i,k} \sum_{l=1}^{N} e^{-\mu_l(z)}[A_l(z)]_{k,j} \\ &= \sum_{l=1}^{N}\sum_{i=1}^{N}\sum_{k=1}^{N} \phi_i(\mu_l(z))\tilde{G}_i(\mu_l(z))p_{i,k}[A_l(z)]_{k,j} \\ &= \sum_{l=1}^{N}\sum_{i=1}^{N}\sum_{k=1}^{N} \Theta_{i,k}(\mu_l(z))[A_l(z)]_{k,j} \\ &= \sum_{l=1}^{N}[e'\Theta(\mu_l(z))A_l(z)]_j, \end{aligned} \tag{18}$$

which completes the proof. □

We now connect the queue length as seen by a departure in steady state to the queue length at an arbitrary time. Let the random variable L denote the queue length at an arbitrary time, and Z the state of $\{Z(t), t \geq 0\}$ at an arbitrary time. Define

$$\begin{aligned} h_j(z) &= E(z^L; Z = j), \qquad 1 \leq j \leq N, \\ h(z) &= (h_1(z), \ldots, h_N(z)), \\ \lambda^{-1} &= \sum_{i=1}^{N} \pi_i \lambda_i^{-1}. \end{aligned}$$

Note that λ is the mean overall arrival rate. The connection between $h(z)$ and $g(z)$ is formulated in the following theorem (cf. Section 3.3 in [27]):

Theorem 4.3 *The generating function $h(z)$ satisfies*

$$h(z)(\Lambda P z - \Lambda) = \lambda(z-1)g(z). \tag{19}$$

Proof: Let $L(t)$ denote the queue length at time t; the process $\{(L(t), Z(t)), t \geq 0\}$ has state space $\{(n,i), n \geq 0, i = 1, \ldots, N\}$. In steady state the average number of transitions per unit time out of state (n,i) is equal to the number of transitions into state (n,i); hence

$$\begin{aligned} &P(L^d = n-1, Z^d = i)\lambda + P(L = n, Z = i)\lambda_i \\ &\quad = P(L^d = n, Z^d = i)\lambda + \sum_{j=1}^{N} P(L = n-1, Z = j)\lambda_j p_{j,i}. \end{aligned}$$

Taking the transforms gives equation (19). □

Remark 4.4 The classical approach for $M/G/1$-type models is to consider the embedded Markov chain at departure epochs first, and then to determine the waiting time distribution by using the connection between this distribution and the departure distribution (cf. [27]). For the present model, the type of customer to be served next does not only depend on the customer type of the departing customer, but also on the sojourn time of the departing customer. This feature essentially complicates an embedded Markov chain approach.

Remark 4.5 The analysis of the number of customers with nonzero service times in the system proceeds along the same lines; of course, in this case, $N(t)$ should only count arrivals of nonzero customers.

4.2 Steady state moments

The results of the previous subsection can be used to determine the factorial moments of the queue length distribution. Define

$$\begin{aligned} d_{r,i} &= E(L^d(L^d-1)\cdots(L^d-r+1); Z^d = i), \qquad r = 0,1,2,\ldots,\ 1 \leq i \leq N, \\ d_r &= [d_{r,1}, d_{r,2}, \ldots, d_{r,N}], \qquad r = 0,1,2,\ldots, \\ a_{r,i} &= E(L(L-1)\cdots(L-r+1); Z = i), \qquad r = 0,1,2,\ldots,\ 1 \leq i \leq N, \\ a_r &= [a_{r,1}, a_{r,2}, \ldots, a_{r,N}], \qquad r = 0,1,2,\ldots. \end{aligned}$$

The moment vectors d_r may be obtained by differentiating (17) and then using the relation $d_r = g^{(r)}(1)$. However, the derivatives of $\mu_i(z)$ and $A_i(z)$ at $z=1$ may be hard to determine. It is easier to proceed as follows. Let

$$M_r(t) = \frac{\partial^r}{\partial z^r}\psi(1,t), \qquad t \geq 0.$$

The moment matrices $M_r(t)$ satisfy the differential equations (see (16)

$$\frac{d}{dt}M_r(t) = r\Lambda P M_{r-1}(t) + (\Lambda P - \Lambda)M_r(t), \qquad t \geq 0, \tag{20}$$

with the initial condition $M_r(0) = I$ if $r = 0$, and $M_r(0) = 0$ if $r > 0$. Multiplying both sides of (20) by $e^{(\Lambda-\Lambda P)t}$, it is readily seen that these equations can be rewritten as

$$\frac{d}{dt}\left(e^{(\Lambda-\Lambda P)t}M_r(t)\right) = re^{(\Lambda-\Lambda P)t}\Lambda P M_{r-1}(t), \qquad t \geq 0.$$

Hence,

$$M_r(t) = re^{-(\Lambda-\Lambda P)t}\int_0^t e^{(\Lambda-\Lambda P)x}\Lambda P M_{r-1}(x)dx, \qquad t \geq 0, \tag{21}$$

from which it is obvious that $M_r(t)$ can be determined recursively, starting with $r = 0$. The following lemma gives the solutions for $r = 0$ and $r = 1$; the solutions for $r > 1$ can be obtained similarly. In the lemma we abbreviated $\mu_i(1)$ and $A_i(1)$ simply by μ_i and A_i.

Lemma 4.6 *The moment matrices $M_0(t)$ and $M_1(t)$ satisfy*

$$\begin{aligned} M_0(t) &= \sum_{i=1}^N e^{-\mu_i t}A_i; \\ M_1(t) &= \sum_{k=1}^N \sum_{\substack{l=1 \\ l\neq k}} \frac{e^{-\mu_k t} - e^{-\mu_l t}}{\mu_l - \mu_k} A_k \Lambda P A_l + \sum_{k=1}^N te^{-\mu_k t}A_k\Lambda P A_k. \end{aligned}$$

Proof: The expression for $M_0(t)$ immediately follows from Theorem 4.1. Substituting

$$e^{\pm(\Lambda-\Lambda P)t} = \sum_{k=1}^N e^{\pm\mu_k t}A_k$$

into (21) for $r = 1$ and using the fact that $A_k^2 = A_k$ and $A_kA_l = 0$ if $k \neq l$, yields

$$\begin{aligned} M_1(t) &= e^{-(\Lambda-\Lambda P)t}\left\{\sum_{k=1}^N \sum_{\substack{l=1 \\ l\neq k}}^N \frac{e^{(\mu_k-\mu_l)t} - 1}{\mu_k - \mu_l} A_k\Lambda P A_l + \sum_{k=1}^N tA_k\Lambda P A_k\right\} \\ &= \sum_{k=1}^N \sum_{\substack{l=1 \\ l\neq k}}^N \frac{e^{-\mu_l t} - e^{-\mu_k t}}{\mu_k - \mu_l} A_k\Lambda P A_l + \sum_{k=1}^N te^{\mu_k t}A_k\Lambda P A_k, \end{aligned}$$

which completes the proof of the lemma. □

Once $M_r(t)$ is known, the moment vectors d_r follow from (cf. (18))

$$d_{r,j} = \frac{d^r}{dz^r}g_j(1) = \sum_{i=1}^N \int_0^\infty dP(T_0 \leq t; Z_0 = i) \sum_{k=1}^N p_{i,k}[M_r(t)]_{k,j}.$$

The results for d_0 and d_1 are formulated in the following theorem.

Theorem 4.7 *The moment vectors d_0 and d_1 are given by*

$$\begin{aligned} d_0 &= \sum_{i=1}^{N} e'\Theta(\mu_i)A_i; \\ d_1 &= \sum_{i=1}^{N} e' \left\{ \sum_{k=1}^{N} \sum_{\substack{l=1 \\ l\neq k}}^{N} \frac{\Theta(\mu_k)-\Theta(\mu_l)}{\mu_l-\mu_k} A_k\Lambda P A_l - \sum_{k=1}^{N} \Theta'(\mu_k)A_k\Lambda P A_k \right\}. \end{aligned}$$

The moment vectors a_r can be found by exploiting the relationship formulated in Theorem 4.3.

Theorem 4.8 *The moment-vectors a_r satisfy the recursive equations:*

$$a_r\Lambda(P-I) = r(\lambda d_{r-1} - a_{r-1}\Lambda P), \tag{22}$$

$$a_r\Lambda e = \lambda d_r e, \tag{23}$$

where $a_{-1} = d_{-1} = 0$ by convention.

Proof: We have

$$\begin{aligned} g(z) &= \sum_{r=0}^{\infty} d_r \frac{(z-1)^r}{r!}, \\ h(z) &= \sum_{r=0}^{\infty} a_r \frac{(z-1)^r}{r!}. \end{aligned}$$

Substituting in equation (19) we get

$$\sum_{r=0}^{\infty} a_r \frac{(z-1)^r}{r!}\Lambda(P-I+(z-1)P) = \lambda(z-1)\sum_{r=0}^{\infty} d_r \frac{(z-1)^r}{r!},$$

which can be rearranged as

$$\sum_{r=0}^{\infty}\left(\frac{a_r}{r!}\Lambda(P-I) + \frac{a_{r-1}}{(r-1)!}\Lambda P\right)(z-1)^r = \sum_{r=0}^{\infty} \lambda d_{r-1}\frac{(z-1)^r}{(r-1)!}.$$

Equating the coefficients of $(z-1)^r$ gives

$$a_r\Lambda(P-I) = r(\lambda d_{r-1} - a_{r-1}\Lambda P). \tag{24}$$

Since $P-I$ is non-invertible with rank $N-1$ we need an extra equation. This one is obtained by post-multiplying (24) with e yielding

$$\lambda d_{r-1}e - a_{r-1}\Lambda e = 0.$$

This completes the proof of the theorem. □

Remark 4.9 For $r = 0$ the equations (22)-(23) reduce to (note that $d_0 e = 1$)

$$\begin{aligned} a_0 \Lambda (P - I) &= 0, \\ a_0 \Lambda e &= \lambda. \end{aligned}$$

Hence, we get

$$a_0 = \lambda \pi \Lambda^{-1}.$$

The mean queue length $a_1 e$ can also be obtained by application of Little's law, i.e.,

$$a_1 e = \lambda (m_1 + e' \Gamma_1) e.$$

5 Numerical Examples

In this section we present some examples to demonstrate the effects of auto-correlation and cross-correlation of the inter-arrival and service times. In each example we set $N = 4$ and we assume exponential service times. Further, we keep the mix of small and large inter-arrival and service times the same; the examples only differ in the dependence structure of inter-arrival and service times.

Example 1: Positively correlated inter-arrival and service times.
The arrival rates and mean service times are given by

$$[\lambda_1 \ \lambda_2 \ \lambda_3 \ \lambda_4] = [.1 \ 10 \ .1 \ 10], \quad [\gamma_1 \ \gamma_2 \ \gamma_3 \ \gamma_4] = u[2 \ 1 \ 2 \ 1],$$

where $u > 0$ is a parameter, and is used to explore the effect of increasing the mean service times (and hence the traffic intensity) on the expected waiting times. It is readily verified that the cross-correlation between the inter-arrival and service time is equal to +1 (see Remark 2.2). We study two cases. In case (a), the transition probability matrix is set to

$$P = \begin{pmatrix} 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \\ 1 & 0 & 0 & 0 \end{pmatrix}, \tag{25}$$

So the inter-arrival and service times are auto-correlated. For example, assuming a coefficient of variation of the service time to be one, we can compute (using Remark 2.1) the auto-correlation for the service times to be

$$\rho(S_0, S_n) = (-1)^n \frac{1}{11}, \quad n \geq 1. \tag{26}$$

The n-step auto-correlation for the inter-arrival times is given by

$$\rho(A_1, A_{n+1}) = (-1)^n .3289, \quad n \geq 1. \tag{27}$$

This situation will be compared with case (b), where P satisfies

$$P = \begin{pmatrix} .25 & .25 & .25 & .25 \\ .25 & .25 & .25 & .25 \\ .25 & .25 & .25 & .25 \\ .25 & .25 & .25 & .25 \end{pmatrix}. \tag{28}$$

In this case the successive inter-arrival times and service times are iid, i.e., there is no auto-correlation. In both cases, the traffic intensity is given by

$$\rho = \frac{\sum \pi_i \gamma_i}{\sum \pi_i \lambda_i^{-1}} = .297u.$$

The system is stable if $\rho < 1$. Thus u can be in allowed to vary in $(0, 3.37)$. In Figure 1 we present the mean waiting times as a function of the traffic intensity $\rho \in [.5, 1)$. The behavior in $\rho < .5$ is as expected and is not plotted. Figure 1 shows that the case with auto-correlation has slightly lower mean waiting times.

Figure 1: Cross-correlation +1: Mean waiting times as a function of ρ for case (a) with auto-correlation and case (b) without auto-correlation.

Example 2: Negatively correlated inter-arrival and service times.
The arrival rates and mean service times are given by

$$[\lambda_1\ \lambda_2\ \lambda_3\ \lambda_4] = [.1\ 10\ .1\ 10], \quad [\gamma_1\ \gamma_2\ \gamma_3\ \gamma_4] = u[1\ 2\ 1\ 2],$$

where $u \geq 0$. Now the cross-correlation between the inter-arrival and service time is equal to -1. We will again study two cases (c) and (d), with the transition probability matrices (25) and (27), respectively. The auto-correlation functions for the service times and the inter-arrival times in case (c) continue to be given by Equations (26) and (**??**), while they are uncorrelated in case (d). The results for the mean waiting times are presented in Figure 2, as a function of the traffic intensity $\rho = .297u$. Clearly, in this case, auto-correlation is able to exploit the big differences between the mean inter-arrival and service times. The mean waiting times for auto-correlated inter-arrival and service times are substantially less than the ones for no auto-correlation.

Figure 2: Cross-correlation -1: Mean waiting times as a function of ρ for case (c) with auto-correlation and case (d) without auto-correlation.

Example 3: Independent inter-arrival and service times.
Now the arrival rates and mean service times are given by

$$[\lambda_1\ \lambda_2\ \lambda_3\ \lambda_4] = [.1\ 10\ .1\ 10], \quad [\gamma_1\ \gamma_2\ \gamma_3\ \gamma_4] = u[2\ 2\ 1\ 1],$$

where $u \geq 0$. Again we consider two cases. In case (e), we use

$$P = \begin{pmatrix} 0 & .5 & 0 & .5 \\ .5 & 0 & .5 & 0 \\ 0 & .5 & 0 & .5 \\ .5 & 0 & .5 & 0 \end{pmatrix},$$

so that the inter-arrival times are auto-correlated with auto-correlation function given in Equation (**??**), but service times are iid. In case (f), P is given by

$$P = \begin{pmatrix} .25 & .25 & .25 & .25 \\ .25 & .25 & .25 & .25 \\ .25 & .25 & .25 & .25 \\ .25 & .25 & .25 & .25 \end{pmatrix}.$$

so that the inter-arrival times and service times are both independent (and hyper-exponentially distributed). However, there is no cross correlation between the inter-arrival times and the service times. The traffic intensity is given by $\rho = .297u$ as before. In Figure 3 the mean waiting times are shown for both cases; the results illustrate that the mean waiting times are substantially less for auto-correlated inter-arrival times.

Figure 3: Cross-correlation 0: Mean waiting times as a function of ρ for case (e) with auto-correlated and case (f) with independent inter-arrival times.

These examples clearly indicate the impact of auto-correlation and cross-correlation on the waiting times.

6 Appendix: Proof of Theorem 3.2

We first assume that for some $\epsilon > 0$ the transforms $\tilde{G}_i(s)$ are analytic for all s with $\mathrm{Re}(s) > -\epsilon$. This holds, e.g., for service time distributions with an exponential tail or for distributions with a finite support.

Let us consider the determinant

$$\det(\tilde{G}(s)P + s\Lambda^{-1} - I) = \det(H(s) + sI - \Lambda)/\det(\Lambda),$$

and let C_δ denote the circle with its center located at $\max_i \lambda_i$ and radius $\delta + \max_i \lambda_i$, with $0 < \delta < \epsilon$. We will prove that the determinant has exactly N zeros inside the circle C_δ for all δ sufficiently small. We follow the main idea in the proof of a similar theorem in Appendix 2 of [34]. We first prove the following lemma.

Lemma 6.1 *For $0 \leq u \leq 1$ and small $\delta > 0$,*

$$\det(u\tilde{G}(s)P + s\Lambda^{-1} - I) \neq 0, \qquad s \in C_\delta. \tag{29}$$

Proof: For $0 \le u \le 1$ and $s \in C_\delta$ with $\mathrm{Re}(s) \ge 0$, the matrix $u\tilde{G}(s)P + s\Lambda^{-1} - I$ is diagonally dominant, since

$$\begin{aligned} \mid u\tilde{G}_i(s)p_{i,i} + s/\lambda_i - 1 \mid \ &\ge\ \mid s/\lambda_i - 1 \mid - u\tilde{G}_i(0)p_{i,i} \ \ge\ 1 + \delta/\lambda_i - u\tilde{G}_i(0)p_{i,i} \\ &>\ \lambda\tilde{G}_i(0) - u\tilde{G}_i(0)p_{i,i} \ =\ \sum_{j \ne i} u\tilde{G}_i(0)p_{i,j} \ \ge\ \sum_{j \ne i} \mid u\tilde{G}_i(s)p_{i,j} \mid . \end{aligned} \tag{30}$$

Hence, its determinant is nonzero (see, e.g., pp. 146-147 in [28]). To prove this for $s \in C_\delta$ with $\mathrm{Re}(s) < 0$ we first note that the determinant is nonzero if and only if 0 is not an eigenvalue. So we proceed by studying the eigenvalues of $u\tilde{G}(s)P + s\Lambda^{-1} - I$ near s =0. If we write

$$u\tilde{G}(s)P + s\Lambda^{-1} - I = P - I + s\Lambda^{-1} + ((u-1)\tilde{G}(s) + \tilde{G}(s) - I)P, \tag{31}$$

we see that for (s, u) close to $(0, 1)$, the matrix above is a perturbation of $P - I$. Since P is irreducible, $P - I$ has a simple eigenvalue 0. Then in a neighborhood of $(0, 1)$, there exist differentiable $x(s, u)$ and $\mu(s, u)$ such that

$$(u\tilde{G}(s)P + s\Lambda^{-1} - I)x(s, u) = \mu(s, u)x(s, u), \qquad e'x(s, u) = 1,$$

and such that $\mu(0, 1) = 0$ and $x(0, 1) = e$. Differentiating this equation with respect to s and setting $s = 0$ and $u = 1$ in the result, we obtain

$$(P - I)\frac{\partial}{\partial s}x(0, 1) + (\Lambda^{-1} - \Gamma)e = \frac{\partial}{\partial s}\mu(0, 1)e.$$

Pre-multiplying both sides with π gives

$$\frac{\partial}{\partial s}\mu(0, 1) = \pi(\Lambda^{-1} - \Gamma)e. \tag{32}$$

Similarly, by differentiating with respect to u we get

$$\frac{\partial}{\partial u}\mu(0, 1) = 1.$$

Hence, for (s, u) close to $(0, 1)$, it holds that

$$\mu(s, u) \approx s\pi(\Lambda^{-1} - \Gamma)e + u - 1.$$

Since $\pi(\Lambda^{-1} - \Gamma)e > 0$ by virtue of (2), we can conclude that $\mu(s, u) \ne 0$ for $s \in C_\delta$ with $\mathrm{Re}(s) < 0$, for small $\delta > 0$ and u close to 1, say $1 - \hat{\delta} \le u \le 1$. Finally, for $0 \le u < 1 - \hat{\delta}$, it can be shown, similarly to (29), that $u\tilde{G}(s)P + s\Lambda^{-1} - I$ is diagonally dominant for $s \in C_\delta$ with $\mathrm{Re}(s) < 0$, provided δ is small enough such that

$$1 - \delta/\lambda_i > (1 - \hat{\delta})\tilde{G}_i(-\delta).$$

This completes the proof of the lemma. □

Let $f(u)$ be the number of zeros of $\det(u\tilde{G}(s)P + s\Lambda^{-1} - I)$ inside C_δ. Then we have

$$f(u) = \frac{1}{2\pi i} \int_{C_\delta} \frac{\frac{\partial}{\partial s} \det(u\tilde{G}(s)P + s\Lambda^{-1} - I)}{\det(u\tilde{G}(s)P + s\Lambda^{-1} - I)} ds.$$

So $f(u)$ is a continuous function on $[0, 1]$, integer-valued, and hence constant. Since $f(0) = N$, this implies that also $f(1) = N$. Letting δ tend to 0, we can conclude that $\det(\tilde{G}(s)P + s\Lambda^{-1} - I)$ has N zeros inside or on C_0.

Clearly, $s = 0$ satisfies

$$\det(\tilde{G}(s)P + s\Lambda^{-1} - I) = 0.$$

That $s = 0$ is a simple solution of the above equation is a consequence of P being irreducible and the stability condition (2); the arguments are presented below.

We evaluate the derivative of the determinant in $s = 0$. Let $D = diag(d_1, \ldots, d_N)$ be the diagonal matrix of eigenvalues of $P - I$, with $d_1 = 0$; since P is irreducible, 0 is a simple eigenvalue of $P - I$. Let Y and X denote the matrices of corresponding left and right eigenvectors, with $YX = I$. Hence,

$$P - I = YDX, \qquad YX = I.$$

Now in a neighborhood of the origin, there exist differentiable $D(s)$, $Y(s)$ and $X(s)$ such that

$$\tilde{G}(s)P + s\Lambda^{-1} - I = Y(s)D(s)X(s)$$

and such that $D(0) = D$, $Y(0) = Y$ and $X(0) = X$. Hence,

$$\det(\tilde{G}(s)P + s\Lambda^{-1} - I) = \det(Y(s)) \det(D(s)) \det(X(s)).$$

By differentiating this equation with respect to s, setting $s = 0$ in the result and using that $d_1 = 0$, we obtain that the derivative of $\det(\tilde{G}(s)P + s\Lambda^{-1} - I)$ in $s = 0$ is equal to $d_1'(0)d_2 \cdots d_N$. This is nonzero, since (see (31))

$$d_1'(0) = \pi(\Lambda^{-1} - \Gamma)e > 0.$$

To finally complete the proof of Theorem 3.2 we have to remove the initial assumption that for some $\epsilon > 0$ the transforms $\tilde{G}_i(s)$ are analytic for all s with $\mathrm{Re}(s) > -\epsilon$. To this end, first consider the 'truncated' service time distributions $G_i^K(x)$ defined as $G_i^K(x) = G_i(x)$ for $0 \leq x < K$ and $G_i^K(x) = 1$ for $x \geq K$. Then Theorem 3.2 holds for the distributions $G_i^K(x)$; by letting K tend to infinity, the result also follows for the original service time distributions. □

Remark 6.2 We have not only proved the existence of N solutions of Equation (8), but also that they are located inside or on the circle with its center at $\max_i \lambda_i$ and radius $\max_i \lambda_i$. This is useful in numerical procedures for finding these zeros.

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