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Periodic Little’s Law

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Motivated by our recent study of patient-flow data from an Israeli emergency department (ED), we establish a sample-path periodic Little’s law (PLL), which extends the sample-path Little’s law (LL) of Stidham (1974). The ED data analysis led us to propose a periodic stochastic process to represent the aggregate ED occupancy level, with the length of a periodic cycle being one week. Because we conducted the ED data analysis over successive hours, we construct our PLL in discrete time. The PLL helps explain the remarkable similarities between the simulation estimates of the average hourly ED occupancy level over a week, using our proposed stochastic model fit to the data, to direct estimates of the ED occupancy level from the data.

We also establish a steady-state stochastic PLL, similar to the time-varying LL of Bertsimas and Mourtzinou (1997) and Fralix and Riano (2010).

Key words: Little’s law, $L = \lambda W$, periodic queues, service systems, data analysis, emergency departments

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1. Introduction

Many service systems with customer response times extending over hours or days can be modeled as periodic queues with the length of a periodic cycle being one week. Examples are hospitals wards, order-fulfillment systems and loan-processing systems. In this paper we establish a periodic version of Little’s law, which can provide insight into the performance of these periodic systems.

We formulate our periodic Little’s law (PLL) in discrete time, assuming that there are $d$ discrete time periods within each periodic cycle. In discrete time, the PLL states that, under appropriate
conditions,

\[ L_k = \sum_{j=0}^{\infty} \lambda_{k-j} F_{k-j,j}, \quad k = 0, 1, \ldots, d - 1, \quad (1) \]

where \( d \) is the number of time periods within each periodic cycle, \( L_k \) is the long-run average number in system in time period \( k \), \( \lambda_k \) is the long-run average number of arrivals in time period \( k \), and \( F_{k,j}, j \geq 0 \), is the long-run proportion of arrivals in time period \( k \) that remain in the system for at least \( j \) time periods, which can be viewed as the complementary cumulative distribution function (ccdf) of the length of stay of an arbitrary arrival. The long-run averages are over all indices of the form \( k + md, m \geq 0 \). These quantities \( \lambda_k, F_{k,j} \), and \( L_k \) are periodic functions of the time index \( k \), exploiting the extension of these periodic functions to all integers, negative as well as positive.

We were motivated to develop the PLL because of a remarkable similarity between two curves we observed in our recent study of patient-flow data from an Israeli emergency department (ED) in Whitt and Zhang (2017b). As part of that study, we developed an aggregate stochastic model of an emergency department (ED) based on a statistical analysis of patient arrival and departure data from the ED of an Israeli hospital, using 25 weeks of data from the data repository associated with the study by Armony et al. (2015). In §6 of Whitt and Zhang (2017b), we conducted simulation experiments to validate the aggregate model of ED patient flow. One of these comparisons compared direct estimates of the average ED occupancy from data to estimations from simulations of the stochastic model, where the distributions of the daily number of arrivals, the arrival-rate function and the LoS distribution are estimated from the data. Figure 1 below shows that the two curves are barely distinguishable. The PLL provides an explanation.

The main contribution of this paper is the sample-path PLL in discrete time, Theorem 1, extending the sample-path Little’s law (LL, \( L = \lambda W \)) established by Stidham (1974); also see El-Taha and Stidham (1999), Fiems and Bruneel (2002), Little (1961, 2011), Whitt (1991, 1992) and Wolff and Yao (2014). This sample-path PLL is different in detail from all previous sample-path LL results (known to us). For example, in addition to the usual limits of averages of the arrival rates and LoS (waiting times), we need to assume a limit for the entire LoS distribution. The necessity of this condition is shown by Example 1 in Section 2.4.
We also establish steady-state stochastic versions of the PLL, which relate more directly to the time-varying LL in Bertsimas and Mourtzinou (1997) and Fralix and Riano (2010). This involves the usual two forms of stationarity associated with arrival times and arbitrary times, that emerges from the Palm theory of stochastic point processes; e.g., see Baccelli and Bremaud (1994) and Sigman (1995), but now both are in discrete time, as in Section 1.7.4 of Baccelli and Bremaud (1994) and Miyazawa and Takahashi (1992). Our steady-state stochastic versions of the PLL extend (and are consistent with) an early PLL for the $M_t/GI/1$ queue in Proposition 2 of Rolski (1989).

We also conduct additional analysis of the ED data to provide additional support for the stochastic model proposed in Whitt and Zhang (2017b), despite the negligible support provided by Figure 1.

The rest of the paper is organized as follows: In §2 we state and discuss the sample-path PLL. In §3, we establish the steady-state stochastic versions of the PLL. In §3.4 and the e-companion we elaborate on the ED application, reviewing the model we built in Whitt and Zhang (2017b), illustrating how it relates to the PLL and providing evidence that the conditions in the theorems are satisfied in our application. In §4 we provide the proof of Theorems in §2. Finally in §5 we draw conclusions.
2. Sample-Path Version of the Periodic Little’s Law

In this section we develop the sample-path PLL. This version is general in that (i) we do not directly make any stochastic assumptions and (ii) we do not directly impose any periodic structure. Instead, we assume that natural limits exist, which we take to be with probability 1 (w.p.1). It turns out that the periodicity of the limit emerges automatically from the assumed existence of the limits.

2.1. Notation and Definitions

We consider discrete time periods indexed by integers \( i, i \geq 0 \). Since multiple events can happen at these times, we need to carefully specify the order of events, just as in the large literature on discrete-time queues, e.g., Bruneel and Kim (1993). We assume that all arrivals in a time period occur before any departures. Moreover, we count the number of customers (patients in the ED in our intended application) in the system in a time period after the arrivals but before the departures. Thus, each arrival can spend \( j \) time periods in the system for any \( j \geq 0 \).

With these conventions, we focus on a single sequence, \( X \equiv \{ X_{i,j} : i \geq 0; j \geq 0 \} \), with \( X_{i,j} \) denoting the number of arrivals in period \( i \) that have length of stay (LoS) \( j \) periods. We also could have customers at the beginning, but without lost of generality, we can view them as a part of the arrivals in time period 0. We define other quantities of interest in terms of \( X \). In particular, with \( \equiv \) denoting equality by definition, the key quantities are:

\[
Y_{i,j} \equiv \sum_{s=j}^{\infty} X_{i,s}: \text{the number of arrivals in time period } i \text{ with LoS greater or equal to } j, j \geq 0,
\]

\[
A_i \equiv Y_{i,0} = \sum_{s=0}^{\infty} X_{i,s}: \text{the total number of total arrivals in time period } i,
\]

\[
Q_i \equiv \sum_{j=0}^{i} Y_{i-j,j} = \sum_{j=0}^{i} A_{i-j} Y_{i-j,j}/A_{i-j}: \text{the number in system in time period } i,
\]

all for \( i \geq 0 \). In the last line, and throughout the paper, we understand \( 0/0 \equiv 0 \), so that we properly treat time periods with 0 arrivals.
2.2. Periodic Averages

We do not directly make any periodic assumptions, but with the periodicity in mind, we consider the following averages over $n$ periods:

\[
\tilde{\lambda}_k(n) \equiv \frac{1}{n} \sum_{m=1}^{n} A_{k+(m-1)d},
\]

\[
\tilde{Q}_k(n) \equiv \frac{1}{n} \sum_{m=1}^{n} Q_{k+(m-1)d} = \frac{1}{n} \sum_{m=1}^{n} \left( \sum_{j=0}^{(k+(m-1)d-j,j)} Y_{k+(m-1)d-j,j} \right),
\]

\[
\tilde{Y}_{k,j}(n) \equiv \frac{1}{n} \sum_{m=1}^{n} Y_{k+(m-1)d,j}, \quad j \geq 0,
\]

\[
\tilde{F}_{k,j}^c(n) \equiv \frac{\tilde{Y}_{k,j}(n)}{\tilde{\lambda}_k(n)} = \frac{1}{\sum_{m=1}^{n} A_{k+(m-1)d}} \sum_{m=1}^{n} Y_{k+(m-1)d,j}, \quad j \geq 0, \text{ and}
\]

\[
\tilde{W}_k(n) \equiv \sum_{j=0}^{\infty} \tilde{F}_{k,j}^c(n), \quad 0 \leq k \leq d - 1,
\]

where $d$ is a positive integer.

Clearly, $\tilde{\lambda}_k(n)$ is the long-run average arrival rate in time period $k$; we think of it applying to all the time periods $(m-1)d + k$ for $0 \leq k \leq d - 1$ and $m \geq 1$. Similarly, $\tilde{Q}_k(n)$ is the long-run average number of customers in the system in time period $k$, while $\tilde{Y}_{k,j}(n)$ is the long-run average number of customers that arrive in time period $k$ that have a LoS greater or equal to $j$ time periods. Thus, $\tilde{F}_{k,j}^c(n)$ is the empirical ccdf, which is the natural estimator of the LoS ccdf of an arrival in time period $k$. Finally, $\tilde{W}_k(n)$ is the sample mean LoS of customers that arrive in time period $k$. We write $n$ as a parameter to indicate that the estimator is computed by averaging over $n$ periodic cycles. We will let $n \to \infty$.

2.3. The Limit Theorem

With the framework introduced above, we can state our main theorem, the sample-path version of the PLL. We first introduce our assumptions, which are just as in the sample-path LL, with one exception. In particular, we assume that

\[(A1) \quad \tilde{\lambda}_k(n) \to \lambda_k, \quad \text{w.p.1 as} \quad n \to \infty, \quad 0 \leq k \leq d - 1,\]

\[(A2) \quad \tilde{F}_{k,j}^c(n) \to F_{k,j}^c, \quad \text{w.p.1 as} \quad n \to \infty, \quad 0 \leq k \leq d - 1, \quad j \geq 0, \text{ and}\]

\[(A3) \quad \tilde{W}_k(n) \to W_k \equiv \sum_{j=0}^{\infty} F_{k,j}^c, \quad \text{w.p.1 as} \quad n \to \infty, \quad 0 \leq k \leq d - 1, \quad (3)\]
where the limits are deterministic and finite. For the sample-path LL, \( d = 1 \) and we do not need \((A2)\).

The assumptions above only assume the existence of limits within the first period, but the limits immediately extend to all \( k \geq 0 \), showing that the limit functions must be periodic functions. We then extend these periodic functions to the entire real line, including the negative time indices. We give a proof of the following in §4.1.

**Lemma 1.** *(periodicity of the limits)* If the three assumptions in (3) hold, then the limits hold for all \( k \geq 0 \), with the limit functions being periodic with period \( d \).

We are now ready to state our main theorem; we give the proof in §4.2.

**Theorem 1.** *(sample-path PLL)* If the three assumptions \((A1), (A2)\) and \((A3)\) in (3) hold, then \( Q_k(n) \) defined in (2) converges as \( n \to \infty \) to a limit that we call \( L_k \). Moreover,

\[
L_k = \sum_{j=0}^{\infty} \lambda_{k-j} F^c_{k-j,j} < \infty \quad \text{as} \quad n \to \infty \quad \text{w.p.1} \tag{4}
\]

for \( 0 \leq k \leq d - 1 \), where \( \lambda_k \) and \( F^c_{k,j} \) are the periodic limits in \((A1)\) and \((A2)\) extended to all integers, negative as well as positive.

### 2.4. Indirect Estimation of \( L_k \) via the PLL

The PLL in Theorem 1 provides an indirect way to estimate the long-run average occupancy level \( L_k \) through the right hand side of (4), as discussed in Glynn and Whitt (1989b) for the ordinary LL. Here we show that the indirect estimator for \( L_k \) is consistent with the direct estimator.

Since we only have data going forward in time from time period 0, we start by rewriting (1) as

\[
\sum_{j=0}^{\infty} \lambda_{k-j} F^c_{k-j,j} = \sum_{i=0}^{k} \lambda_i \sum_{l=0}^{\infty} F^c_{i,k-i+l,d} + \sum_{i=k+1}^{d-1} \lambda_i \sum_{l=1}^{\infty} F^c_{i,k-i+l,d}, \quad 0 \leq k \leq d - 1. \tag{5}
\]

Guided by (5), we let our indirect estimator for \( L_k \) be

\[
\bar{L}_k(n) \equiv \sum_{i=0}^{k} \bar{\lambda}_i(n) \sum_{l=0}^{\infty} \bar{F}^c_{i,k-i+l,d}(n) + \sum_{i=k+1}^{d-1} \bar{\lambda}_i(n) \sum_{l=1}^{\infty} \bar{F}^c_{i,k-i+l,d}(n), \quad 0 \leq k \leq d - 1, \tag{6}
\]
where $\bar{\lambda}_i(n)$ and $\bar{F}_{c,i,j}(n)$ are defined in (2). With data, it is likely that the infinite sums in (6) would be truncated to finite sums, but at a level growing with $n$; we do not address that truncation modification, which we regard as minor.

We now show that the estimator $\bar{L}_k(n)$ in (6) is asymptotically equivalent to the direct estimator $\bar{Q}_k(n)$ in (2); we will prove this result together with Theorem 1 in §4.2.

**Theorem 2. (indirect estimation through the PLL)** Under the conditions of Theorem 1,

$$\lim_{n \to \infty} \bar{L}_k(n) = L_k \text{ w.p.1 for } 0 \leq k \leq d-1,$$

where $\bar{L}_k(n)$ is defined in (6) and $L_k$ is as in Theorem 1.

In applications, the LoS often can be considered to be bounded, i.e., for some $m > 1$, $X_{i,j} = 0$ when $j \geq md$. In that case, condition (A3) is directly implied by condition (A2) and it is possible to bound the error between the direct and indirect estimators for $L_k$, defined as

$$\bar{E}_k(n) \equiv |\bar{L}_k(n) - \bar{Q}_k(n)|,$$

for $\bar{Q}_k(n)$ in (2) and $\bar{L}_k(n)$ in (6), as we show now.

**Corollary 1. (the bounded case)** If, in addition to conditions (A1) and (A2) in Theorem 1, $X_{i,j} = 0$ for $i \geq 0, j \geq dm_u$, then assumption (A3) is necessarily satisfied. If, in addition, $A_i \leq \lambda_u$ for $i \geq 0$, then

$$\bar{R}_n \equiv \max_{0 \leq k \leq d-1} \{|\bar{E}_k(n)|\} \leq \frac{\lambda_u d(m_u + 2)^2}{2n}, \quad n \geq m_u,$$

for $\bar{E}_k(n)$ in (8).

**Proof:** Here we show the proof of the first part of the corollary, i.e. if the LoS is bounded, then assumption (A3) is implied from (A2), and we postpone the second half of proof in §4.3, since it depends on part of the proof of Theorems 1 and 2.

If $X_{i,j} = 0$ for $i \geq 0, j > dm_u$, then $\bar{F}_{k,j}(n) = 0$ for $0 \leq k \leq d-1$ and $j \geq dm_u$. So

$$\bar{W}_k(n) = \sum_{j=1}^{dm_u} \bar{F}_{c,k,j}(n), \quad 0 \leq k \leq d-1,$$
is a finite summation and \( F_{c,k,j}^c = 0 \) for \( 0 \leq k \leq d - 1 \) and \( j > dm_u \). Then

\[
\lim_{n \to \infty} \bar{W}_k(n) = \lim_{n \to \infty} \sum_{j=0}^{dm_u} F_{c,k,j}^c(n) = \sum_{j=0}^{dm_u} \lim_{n \to \infty} F_{c,k,j}^c(n) = \sum_{j=0}^{dm_u} F_{c,k,j}^c = W_k,
\]

which is assumption (A3).

**Remark 1.** (from assumption (A2) to (A3)) In addition to the boundedness condition presented in Corollary 1, there are other mathematical conditions under which (A2) implies (A3), i.e., under which we can interchange the order of the limits. Uniform integrability is a standard condition for this purpose; see p. 185 of Billingsley (1995) and Section 2.6 of El-Taha and Stidham (1999). We prefer (A3) plus (A2) because that makes our conditions easier to compare to the conditions in the ordinary LL.

**Remark 2.** (insightful figure) Let

\[
\bar{E}(n) \equiv \sum_{k=0}^{d-1} \bar{E}_k(n)
\]

for \( \bar{E}_k(n) \) defined in (8). Observe that \( n\bar{E}(n) \) is the total time spent in the system after time period \( nd \) by customers that arrived between time period 0 to time period \( nd - 1 \). Thus, Figure 2 illustrates the relationship between \( \bar{Q}_k(n) \), \( \bar{L}_k(n) \) and \( \bar{E}(n) \). In Figure 2, we plot the time intervals that each of the first 35 arrivals spends in the system as horizontal bars, each with height 1 placed in order of the arrival times. The left end point is the arrival time, while the right end point is the departure time, which need not be in order of arrival. The area of region \( A \) depicts \( n \sum_{k=0}^{d-1} \bar{Q}_k(n) \), while the area of region \( A \cup B \) depicts \( n \sum_{k=0}^{d-1} \bar{L}_k(n) \), so that the area of region \( B \) depicts

\[
n\bar{E}(n) = n \sum_{k=0}^{d-1} \bar{E}_k(n) = n(\sum_{k=0}^{d-1} \bar{L}_k(n) - \sum_{k=0}^{d-1} \bar{Q}_k(n)).
\]

Figure 2 is a variant of Figure 1 in Whitt (1991) and Figures 2 and 3 in Kim and Whitt (2013), as well as similar figures in earlier papers. The figures in Kim and Whitt (2013) are different, because of the initial edge effect, which we avoid by treating arrivals before time 0 in the system as arrivals at time 0.
Figure 2  An example of a periodic queueing system with \(d = 5\) and \(n = 4\). Area A is where we take average to compute \(\bar{Q}_k(4)\); Area \(A \cup B\) is where we take average to compute \(\bar{F}_{k,j}(n)\), and so is included in calculating \(\bar{L}_k(4)\).

2.5. The Assumptions in Theorem 1

When \(d = 1\), the PLL reduces to the non-time-varying ordinary LL. In that case, \(k = 0\) represents all time indices since it is non-time-varying. In the Theorem 1, \(L_0 \equiv \lim_{n \to \infty} \bar{Q}_0(n)\) is the limiting time-average number of customers in the system while \(\lim_{n \to \infty} \bar{\lambda}_0(n) = \lambda_0\) is the limiting time-average arrival rate and the right hand side of (4) becomes

\[
\sum_{j=0}^{\infty} \lambda_{k-j} F_{k-j,j} = \lambda_0 \sum_{j=0}^{\infty} F_{0,j} = \lambda_0 W_0.
\]  

(11)

And Theorem 1 claims that

\[ L_0 = \lambda_0 W_0, \]

which is exactly the ordinary Little’s law. Of course, the ordinary LL can be applied to the time-varying case as well, but we will lose the time structure and get overall averages.

However, there is a difference between our assumptions in (3) and the assumptions in the LL. For the LL, we let \(L\) be the limiting time-average number in the system, \(\lambda\) be the limiting average
arrival rate of customers and $W$ be the limiting customer-average waiting time (time spent in the system or length of stay). Then, if both $\lambda$ and $W$ exist and are finite, then $L$ exists and is finite, and $L = \lambda W$.

Our limit for $\bar{\lambda}_k(n)$ in (A1) is the natural extension; the only difference is that now we require that it holds for each $k$, $0 \leq k \leq d - 1$. The third limit for $\bar{W}_k(n)$ in (A3) parallels the limit for the average waiting time, but again we require that it holds for each $k$, $0 \leq k \leq d - 1$. However, those two limits are not sufficient to determine the number of customers for the periodic case. Now we need to require that the LoS distribution converges for each $k$, $0 \leq k \leq d - 1$, as stated in (A2). We show that this extra condition is needed in the following example.

**Example 1.** (need for convergence of the complementary cdf’s) We now show that we need to assume the limit $\bar{F}_{k,j}(n) \to F_{k,j}$ in (3). For simplicity, let $d = 2$, so that we have 2 time periods in each periodic cycle. Suppose we have 2 systems. In the first one we deterministically have 2 arrivals in the first time period of each periodic cycle, (i.e. 2 arrivals at even indexed time period), with one of them having LoS 0 and LoS 2. In the other one, we also deterministically have 2 arrivals in the first time period of each periodic cycle, but with both of them having LoS 1. Suppose there is no arrival at the odd indexed time period for both of the two systems. Now the two systems have the same $\lambda_k$ and $W_k$. ($\lambda_0 = 2$, $\lambda_1 = 0$, $W_0 = 2$ and $W_1 = 0$.) However, if we count the number of customers in the system, we have $\lim_{n \to \infty} \bar{Q}_0(n) = 3$ for the first system and $\lim_{n \to \infty} \bar{Q}_0(n) = 2$ for the second one.

**3. Steady-State Stochastic Versions of the Periodic Little’s Law**

We now discuss stochastic analogs of Theorem 1. We start by deriving a steady-state stochastic PLL by applying Theorem 1. Afterwards, we consider other variants. For this initial version, we aim for simplicity. Thus, we assume that the basic stochastic process is both stationary and ergodic, so that steady-state means coincide with long-run averages.

The main idea is that we now interpret the key quantities $L_k$ and $\lambda_k$ appearing in (1) as appropriate expected values of random variables associated with the system in periodic steady state. As we see it, there are two main issues:
1. What is meant by periodic steady state?

2. What is \( F_{k,j}^c \) or, equivalently, what is the probability distribution of the LoS of an arbitrary arrival in period \( k \)?

### 3.1. Periodic Steady State

In order to construct periodic steady-state, we assume that the basic stochastic process \( Y \equiv \{ Y_n : n \in \mathbb{Z} \} \) with

\[
Y_n \equiv \{ Y_{nd+k,j} : 0 \leq k \leq d-1; j \geq 0 \}
\]  

introduced in §2.1 is a stationary sequence of nonnegative random elements indexed by the integer \( n \). For each integer \( n \), the random element \( Y_n \) takes values in the space \((Z^d)^\infty \equiv \mathbb{Z}^d \times \mathbb{Z}^d \times \cdots \). See Chapter 6 of Breiman (1968) and Baccelli and Bremaud (1994), Sigman (1995) for background on stationary processes and their application to queues. Just as for the time-varying LL, as discussed in Fralix and Riano (2010), it is important to apply the Palm transformation, but we avoid that issue by exploiting the established limits for the averages.

Without loss of generality, we now regard our stochastic processes as stationary processes on the integers \( \mathbb{Z} \), negative as well as positive; see Proposition 6.5 of Breiman (1968). As usual, we mean strictly stationary; i.e., the finite-dimensional distributions are independent of time shifts, which in turn means that, for each \( k \) and each \( k \)-tuple \((n_1, \ldots, n_k)\) of integers in \( \mathbb{Z} \),

\[
(Y_{n_1}, \ldots, Y_{n_k}) \overset{d}{=} (Y_{n_1+m}, \ldots, Y_{n_k+m}) \quad \text{for all} \quad m \in \mathbb{Z},
\]

with \( \overset{d}{=} \) denoting equality in distribution.

As a consequence of the stationarity assumed for \( \{ Y_n : n \in \mathbb{Z} \} \), we also have stationarity for the associated stochastic process \( \{ (A_{nd+k}, Q_{nd+k}) : n \in \mathbb{Z} \} \), where \( A_{nd+k} \) is the number of arrivals in time period \( nd+k \) and \( Q_{nd+k} \) is the number of customers in the system in period \( nd+k \), both of which are defined in §2.1, but now that we have stationarity on all integers, negative as well as positive. Thus, we have

\[
A_{nd+k} \equiv Y_{nd+k,0} \quad \text{and} \quad Q_{nd+k} \equiv \sum_{j=0}^{\infty} Y_{nd+k-j,j}.
\]
Hence, with some abuse of notation, we let \( \{Y_{k,j} : j \geq 0\}, A_k, Q_k \) be a stationary random element.

In this stochastic setting, we have

\[
\lambda_k \equiv E[A_k] = E[Y_{k,0}] \quad \text{and} \quad L_k \equiv E[Q_k] = \sum_{j=0}^{\infty} E[Y_{k-j,j}].
\]  

(13)

3.2. The Stochastic PLL

We now come to the second issue: In the stochastic setting it remains to define \( F_{c,k,j} \equiv P(W_k > j) \), where \( W_k \) is the time in system for an arbitrary arrival in period \( k \). It is natural to define \( F_{c,k,j} \) by requiring that it agree with the limit of the averages \( \bar{F}_{c,k,j}(n) \) in (2). That limit is well defined if we assume that the basic sequence is ergodic as well as stationary, with \( 0 < E[Y_{k,0}] < \infty \) for all \( k \).

With the stationary framework on all the integers, positive and negative, the stochastic PLL becomes very elementary, because there are no edge effects.

**Theorem 3.** (stochastic PLL) Suppose that \( \{Y_n : n \in \mathbb{Z}\} \) in (12) is stationary and ergodic with \( 0 < \lambda_k \equiv E[Y_{k,0}] < \infty, 0 \leq k \leq d - 1 \). Then, for each \( k \), \( 0 \leq k \leq d - 1 \), and \( j \geq 0 \),

\[
\bar{F}_{c,k,j}(n) \to F_{c,k,j} \equiv \frac{E[Y_{k,j}]}{E[Y_{k,0}]} \quad \text{w.p.1 as } n \to \infty
\]  

and

\[
L_k \equiv E[Q_k] = \sum_{j=0}^{\infty} \lambda_k F_{c,k-j,j}.
\]  

(15)

**Proof.** First, the stationary and ergodic condition with the specified moment assumptions implies that

\[
\bar{Y}_{k,j}(n) \to Y_{k,j} \quad \text{as } n \to \infty \quad \text{w.p.1}
\]

for all \( k \) and \( j \). Then (14) follows immediately by continuity. If we multiply and divide by \( \lambda_k \) within the representation for \( E[Q_k] \) in (13), then we see that

\[
E[Q_k] = \sum_{j=0}^{\infty} E[Y_{k-j,j}] = \sum_{j=0}^{\infty} \lambda_k \frac{E[Y_{k-j,j}]}{\lambda_k} = \sum_{j=0}^{\infty} \lambda_k F_{c,k-j,j},
\]

\[\blacksquare\]
3.3. The Discrete-Time Periodic $G_t/GI_t/\infty$ Model

The candidate model for the ED proposed in Whitt and Zhang (2017b) was a special case of the periodic $G_t/GI_t/\infty$ infinite-server model, in which the LoS variables are mutually independent and independent of the arrival process, with a cpdf $F_{k,j} \equiv P(W_k \geq j)$ for a steady-state LoS in time period $k$ that depends only on the time period $k$ within the periodic cycle (a week). This strong local condition provides a sufficient condition for the steady-state stochastic PLL. That can be seen from the following proposition.

**Proposition 1.** *(the $G_t/GI_t/\infty$ special case)* For the stationary $G_t/GI_t/\infty$ infinite-server model specified above, where $Y \equiv \{Y_n : n \in \mathbb{Z}\}$ in (12) is strictly stationary with finite mean values, then

$$E[Y_{k,j}] = F_{k,j}^c E[Y_{k,0}],$$

consistent with formula (14).

**Proof.** Let \{\(W_{k,i} : i \geq 1\)\} be a sequence of i.i.d. LoS variables associated with arrivals in time period $k$, which is also independent of the arrival process and the other LoS variables. Under those conditions,

$$E[Y_{k,j}] = \sum_{m=1}^{\infty} \sum_{i=1}^{m} P(W_{k,i} > j)P(A_k = m)$$

$$= \sum_{m=1}^{\infty} m F_{k,j}^c P(A_k = m) = F_{k,j}^c E[A_k] = F_{k,j}^c E[Y_{k,0}].$$

3.4. Further Statistical Tests of the Infinite-Server Model

We now report results directly testing the $GI_t$ assumption in the stochastic model proposed in Whitt and Zhang (2017b). We briefly review the data analysis in the e-companion here.

Now we first test whether the LoS distribution in period $k$ can be regarded as being independent of the number of arrivals in period $k$. To be specific, let $A_k^{(m)}$ be the number of arrivals in hour $k$ of week $m$, where in our ED case $1 \leq k \leq 7 \times 24 = 168$ and $m$ is from 1 to 25. And let $W_k^{(i)}$ be the average LoS of arrivals in hour $k$ of week $m$. For each $k$, we compute the estimated (sample) Pearson
correlation coefficients of $A_k^{(i)}$ and $W_k^{(m)}$, using samples where $A_k^{(m)} > 0$; See p.169 of Casella and Berger (2002) for background. The plot in the top of Figure 3 shows the correlation coefficients ($r_k$) of all 168 hours in a week, where (if $A_k^{(m)} > 0$ for all $m$)

$$r_k = \frac{\sum_{m=1}^{25} (A_k^{(m)} - \bar{A}_k)(W_k^{(m)} - \bar{W}_k)}{\sqrt{\sum_{m=1}^{25} (A_k^{(m)} - \bar{A}_k)^2} \sqrt{\sum_{m=1}^{25} (W_k^{(m)} - \bar{W}_k)^2}},$$

where $\bar{A}_k \equiv (1/25) \sum_{i=1}^{25} A_k^{(m)}$ and $\bar{W}_k \equiv (1/25) \sum_{i=1}^{25} W_k^{(m)}$. For any $m$ such that $A_k^{(m)} = 0$, we remove those terms in the summation and the average correspondingly.

The middle left plot compares the quantile of the coefficients with the normal distribution, indicates that the distribution of the 168 correlation coefficients are approximately normal, with sample mean 0.056 and sample standard error 0.21. It shows that there is no significant evidence indicating that the LoS distribution is related to the number of arrivals within each hour. The middle right plot is a scatter plot of $(A_{36}^{(m)}, W_{36}^{(m)})$, i.e. at noon on Monday, excluding samples with no arrivals. and the solid line is the mean of $W_{36}^{(m)}$'s with same number of arrivals. Finally, the bottom plot shows the number of hours with no arrival for each hour in a week, i.e., the number of $i$'s such that $A_k^{(m)} = 0$ for $k = 1, 2, \ldots, 7 \times 24 = 168$.

In summary, the statistical results in this section provide additional support for the stochastic model of the ED proposed in Whitt and Zhang (2017b). In the e-companion to this paper we review the data analysis in Whitt and Zhang (2017b) and show the results of our studies that give evidence that the ED data are consistent with the conditions of Theorem 1 and Proposition 1.
Figure 3  
Top: the estimated linear correlation between the number of arrivals and the mean LoS for each hour of a week; 
Middle left: the $Q-Q$ (quantile) plot of the correlation coefficients compared to a Gaussian distribution; 
Middle right: an example of the relationship between the number of arrivals and the mean LoS at noon on Monday, while the solid line is the average for each column of points; 
Bottom: the number of hours with no arrival for each hour in a week.
3.5. Exploiting the Palm Theory in Continuous Time

Finally, we observe that the Palm theory in Rolski (1989) and Fralix and Riano (2010) also provides a general steady-state stochastic PLL in the conventional continuous-time setting. For this steady-state stochastic result, we now assume that the arrival process is a simple point process (arrivals occur on a time) on the entire real line with a well defined arrival rate $\lambda(t)$ at time $t$. This was satisfied in the ED because the arrival times actually have very detailed time stamps.

Thus, letting $N(t) - N(s)$ be the number of arrivals in the interval $[s, t]$, we assume that

$$E[N(t)] - E[N(s)] = \int_s^t \lambda(s) \; ds,$$

where the arrival-rate function $\lambda(t)$ is a periodic function with periodic cycle length $c$, which is also right continuous with left limits.

As in §2 of Fralix and Riano (2010), we let $W(t)$ be the waiting time of the last arrival before time $t$. That convention yields a well-defined waiting-time process $\{W(t) : t \in \mathbb{R}\}$.

As a continuous-time analog of the periodic stationarity assumed in Section 3.1, we assume that the queue-length (number in system) process $Q \equiv \{Q(t) : t \in \mathbb{R}\}$ has a distribution that is invariant under time shifts by $c$. The arrival process is included by the upward jumps on $Q$. As a consequence of this $c$-stationarity, the set of Palm measures $\{N_s : s \in \mathbb{R}\}$ associated with the arrival process $N$ is periodic with period $c$. The mean queue length is expressed in terms of the tail probabilities

$$F^c_{t,s} \equiv P_t(W(t) > x)$$

under the Palm measures $P_t$, which are periodic with period $c$.

Then, paralleling the remark after Theorem 3.1 of Fralix and Riano (2010), Theorem 3.1 of Fralix and Riano (2010) implies the following continuous-analog of (1), which was already given for the $M_t/GI/1$ special case in Rolski (1989).

**Theorem 4.** (*continuous-time PLL following from Rolski (1989) and Fralix and Riano (2010)*)

Under the conditions above,

$$E[Q(t)] = \int_0^\infty F^c_{t-s,s} \lambda(t - s) \; ds$$

where $F^c_{t-s,s}$ is defined in (17).
4. Proofs

We now provide the postponed proofs of Lemma 1, Theorem 1, Theorem 2 and Corollary 1. In the proof, all the limits are in the sense of almost sure convergence.

4.1. Proof of Lemma 1

We will show that under the three assumptions in (3), we have \( \lim_{n \to \infty} \bar{\lambda}_{k+ld}(n) = \bar{\lambda}_{k+ld} = \lambda_k \), \( \lim_{n \to \infty} \bar{F}_{k+ld,j}^c(n) = \bar{F}_{k+ld,j}^c = F_{k,j}^c \), and \( \lim_{n \to \infty} \bar{W}_{k+ld}(n) = \bar{W}_{k+ld} = W_k \),

\[
\begin{align*}
\lim_{n \to \infty} \bar{\lambda}_{k+ld}(n) &\equiv \lambda_{k+ld} = \lambda_k, \\
\lim_{n \to \infty} \bar{F}_{k+ld,j}^c(n) &\equiv F_{k+ld,j}^c = F_{k,j}^c, \quad \text{and} \\
\lim_{n \to \infty} \bar{W}_{k+ld}(n) &\equiv W_{k+ld} = W_k,
\end{align*}
\]

where \( \lambda_k \), \( F_{k,j}^c \) and \( W_k \) are the same constants as in (3).

**Proof:** By the definition of \( \lambda_k \),

\[
\lambda_k = \lim_{n \to \infty} \bar{\lambda}_k(n) = \lim_{n \to \infty} \frac{1}{n} \sum_{m=1}^{n} A_{k+(m-1)d}
= \lim_{n \to \infty} \frac{1}{n} (A_k + \sum_{m=1}^{n-1} A_{(k+d)+(m-1)d}) = \lim_{n \to \infty} \frac{n-1}{n} \frac{1}{n-1} \sum_{m=1}^{n-1} A_{(k+d)+(m-1)d}
= \lambda_{k+d}.
\]

Next, by (2), we have \( \bar{Y}_{k,j}(n) = \bar{\lambda}_k(n) \bar{F}_{k,j}^c(n) \). By assumption (3), we know that \( \lim_{n \to \infty} \bar{Y}_{k,j}(n) = \lambda_k F_{k,j}^c \) exists for all \( 0 \leq k \leq d-1 \) and \( j \geq 0 \). Using the same argument as for \( \lambda_k \), we know that \( \lim_{n \to \infty} \bar{Y}_{k+d,j}(n) = \lim_{n \to \infty} \bar{\lambda}_{k+d}(n) \) for all \( 0 \leq k \leq d-1 \), \( j \geq 0 \). Then,

\[
F_{k+d,j}^c = \lim_{n \to \infty} \bar{F}_{k+d,j}^c(n) = \frac{\lim_{n \to \infty} \bar{Y}_{k+d,j}(n)}{\lim_{n \to \infty} \lambda_{k+d}(n)} = \frac{\lambda_k F_{k,j}^c}{\lambda_k} = F_{k,j}^c, \quad \text{for all} \quad 0 \leq k \leq d-1, \quad j \geq 0.
\]

Similarly, for \( W_{k+d} \) we have

\[
W_{k+d} = \sum_{j=0}^{\infty} F_{k+d,j}^c = \sum_{j=0}^{\infty} F_{k,j}^c = W_k \quad \text{for all} \quad 0 \leq k \leq d-1.
\]

By induction, we proved (18). \( \square \)
4.2. Proof of Theorem 1 and 2

Proof: The proof is done in two steps. In step 1, we prove Theorem 2, i.e. show that \( \lim_{n \to \infty} \bar{L}_k(n) = \sum_{j=0}^{\infty} \lambda_{k-j} F_{k-j,j}^{c} \) for \( k = 0, 1, \cdots, d-1 \), and then in step 2, we show that \( \lim_{n \to \infty} \bar{Q}_k(n) = \lim_{n \to \infty} \bar{L}_k(n) \), thus complete the proof of Theorem 1.

Step 1: For any given \( k = 0, 1, \cdots, d-1 \) and any \( \epsilon > 0 \) fixed, by assumption (A1), there exists \( N_1 \), such that for any \( n > N_1 \), we have \( \sup_{0 \leq k \leq d-1} |\bar{\lambda}_k(n) - \lambda_k| < \epsilon \). Given assumption (A2), by the series form of Scheffé’s lemma (p.215 of Billingsley (1995)), we know that assumption (A3) is equivalent to

\[
\lim_{n \to \infty} \sum_{j=0}^{\infty} |\bar{F}_{k,j}^{c}(n) - F_{k,j}^{c}| = 0. \tag{22}
\]

So there exists \( N_2 \), such that for any \( n > N_2 \), we have \( \sup_{0 \leq k \leq d-1} \sum_{j=0}^{\infty} |\bar{F}_{k,j}^{c}(n) - F_{k,j}^{c}| < \epsilon \). Let \( N_3 = \max\{N_1, N_2\} \), then when \( n > N_3 \),

\[
|\bar{L}_k(n) - \sum_{j=0}^{\infty} \lambda_{k-j} F_{k-j,j}^{c}| = |\sum_{i=0}^{k} \bar{\lambda}_i(n) \sum_{l=0}^{\infty} F_{i,k-i+ld}^{c}(n) + \sum_{i=k+1}^{d-1} \bar{\lambda}_i(n) \sum_{l=1}^{\infty} F_{i,k-i+ld}^{c}(n)
- \sum_{i=0}^{k} \lambda_i \sum_{l=0}^{\infty} F_{i,k-i+ld}^{c} - \sum_{i=k+1}^{d-1} \lambda_i \sum_{l=1}^{\infty} F_{i,k-i+ld}^{c}|
= |\sum_{j=0}^{k} \lambda_{k-j} \bar{F}_{k-j,j}^{c}(n) + \sum_{m=1}^{d} \sum_{j=1}^{\infty} \lambda_{d-j} \bar{F}_{d-j,(m-1)d+j+k}^{c}(n)
- \sum_{j=0}^{k} \lambda_{k-j} \bar{F}_{k-j,j}^{c} - \sum_{m=1}^{d} \lambda_{d-j} \bar{F}_{d-j,(m-1)d+j+k}^{c}|
\leq |\sum_{j=0}^{k} \lambda_{k-j} \bar{F}_{k-j,j}^{c}(n) + \sum_{m=1}^{d} \sum_{j=1}^{\infty} \lambda_{d-j} \bar{F}_{d-j,(m-1)d+j+k}^{c}(n)
- \sum_{j=0}^{k} \lambda_{k-j} \bar{F}_{k-j,j}^{c} - \sum_{m=1}^{d} \lambda_{d-j} \bar{F}_{d-j,(m-1)d+j+k}^{c}|
+ \epsilon \left( \sum_{j=0}^{k} \bar{F}_{k-j,j}^{c}(n) + \sum_{m=1}^{d} \sum_{j=1}^{\infty} \bar{F}_{d-j,(m-1)d+j+k}^{c}(n) \right)
= \left( \max_{0 \leq k \leq d-1} \lambda_k \right) \left( \sum_{j=0}^{k} |\bar{F}_{k-j,j}^{c}(n) - \bar{F}_{k-j,j}^{c}| + \sum_{m=1}^{d} \sum_{j=1}^{\infty} |\bar{F}_{d-j,(m-1)d+j+k}^{c}(n) - F_{d-j,(m-1)d+j+k}^{c}| \right)
+ \epsilon \left( \sum_{j=0}^{k} \bar{F}_{k-j,j}^{c}(n) + \sum_{m=1}^{d} \sum_{j=1}^{\infty} \bar{F}_{d-j,(m-1)d+j+k}^{c}(n) \right)
\]
\[ \leq \left( \max_{0 \leq k \leq d-1} \lambda_k \right) e + ed \left( \max_{0 \leq k \leq d-1} W_k + \varepsilon \right). \] 

\[ \text{Hence } \lim_{n \to \infty} (\bar{L}_k(n) - \sum_{j=0}^{\infty} \lambda_{k-j} F_{k-j,1}) = 0 \text{ for all } 0 \leq k \leq d - 1 \text{ and we have proved Theorem 2.} \]

\textbf{Step 2:} Next, we show that \( L_k \equiv \lim_{n \to \infty} \bar{Q}_k(n) = \lim_{n \to \infty} \bar{L}_k(n) \). To do so, actually we will prove that

\[ \bar{E}(n) \equiv \sum_{k=0}^{d-1} \bar{E}_k(n) \equiv \sum_{k=0}^{d-1} |\bar{L}_k(n) - \bar{Q}_k(n)| \to 0 \text{ as } n \to \infty. \] 

We further divide this step into 2 substeps. In the first substep, we compute the expression of \( \bar{E}(n) \), and then in the second substep we show it goes to 0 as \( n \) goes to infinity.

\textbf{Step 2.1:} By some transformation, we know that

\[ \bar{Q}_k(n) = \frac{1}{n} \sum_{m=1}^{n} Q_{k+(m-1)d} = \frac{1}{n} \sum_{m=1}^{n} \left( \sum_{j=0}^{k+(m-1)d} Y_{k+(m-1)d-j,j} \right) \]

\[ = \frac{1}{n} \sum_{m=1}^{n} \sum_{j=0}^{k} Y_{k-j+(m-1)d,j} + \frac{1}{n} \sum_{m=2}^{n} \sum_{j=k+1}^{k+(m-1)d} Y_{k-j+(m-1)d,j} \]

\[ = \frac{1}{n} \sum_{m=1}^{n} \sum_{j=0}^{k} Y_{k-j+(m-1)d,j} + \frac{1}{n} \sum_{m=1}^{n-1} \sum_{j=1}^{md} Y_{d-j+(m-1)d,j+k} \]

\[ = \frac{1}{n} \sum_{m=1}^{n} \sum_{j=0}^{k} Y_{k-j+(m-1)d,j} + \frac{1}{n} \sum_{m=1}^{n-1} \sum_{j=1}^{d} Y_{d-j+(m-1)d,j+k}(s-1)d, \]

and

\[ \bar{L}_k(n) = \sum_{j=0}^{k} \bar{Y}_{k-j,j}(n) + \sum_{s=1}^{\infty} \sum_{j=1}^{d} \bar{Y}_{d-j,j+k+(s-1)d}(n) \]

\[ = \frac{1}{n} \sum_{m=1}^{n} \sum_{j=0}^{k} Y_{k-j+(m-1)d,j} + \frac{1}{n} \sum_{s=1}^{\infty} \sum_{m=1}^{n} \sum_{j=1}^{d} Y_{d-j+(m-1)d,j+k+(s-1)d}. \]

We may now study the absolute difference between \( \bar{L}_k(n) \) and \( \bar{Q}_k(n) \). Here

\[ \bar{E}_k(n) = |\bar{L}_k(n) - \bar{Q}_k(n)| \]

\[ = \frac{1}{n} \sum_{s=1}^{n-1} \sum_{m=n-s+1}^{n} \sum_{j=1}^{d} Y_{d-j+(m-1)d,j+k+(s-1)d} + \frac{1}{n} \sum_{s=n}^{n} \sum_{m=1}^{n} \sum_{j=1}^{d} Y_{d-j+(m-1)d,j+k+(s-1)d} \]

\[ = \frac{1}{n} \sum_{m=1}^{n} \sum_{j=1}^{d} \sum_{s=n-m+1}^{\infty} Y_{d-j+(m-1)d,j+k+(s-1)d}. \]
and summing over \( k = 0, 1, \ldots, d - 1 \) further gives

\[
\tilde{E}(n) \equiv \sum_{k=0}^{d-1} \tilde{E}_k(n) = \sum_{k=0}^{d-1} \frac{1}{n} \sum_{m=1}^{n} \sum_{d-j+(m-1)d+k+(s-1)d} Y_{d-j+(m-1)d+s} = \frac{1}{n} \sum_{m=1}^{n} \sum_{j=1}^{d} \sum_{s=(n-m)d}^{\infty} Y_{d-j+(m-1)d+s}. \tag{27}
\]

**Step 2.2:** Now it suffices to show that \( \tilde{E}(n) \to 0 \) as \( n \to \infty \). For that purpose, let \( N_1, N_2 \) and \( N_3 \) be the same as in the beginning of the proof, depending on given \( \epsilon \). Then, when \( n > N_3 \), we have

\[
\left| \sum_{j=0}^{\infty} \tilde{Y}_{k,j}(n) - \sum_{j=0}^{\infty} \lambda_k F_{k,j}^c \right| = \left| \sum_{j=0}^{\infty} \tilde{\lambda}_k(n) F_{k,j}^c(n) - \sum_{j=0}^{\infty} \lambda_k F_{k,j}^c \right| \\
\leq \left| \sum_{j=0}^{\infty} \lambda_k F_{k,j}^c(n) - \sum_{j=0}^{\infty} \lambda_k F_{k,j}^c \right| + \epsilon \sum_{j=0}^{\infty} F_{k,j}^c(n) \\
\leq \lambda_k \epsilon + \epsilon (W_k + \epsilon). \tag{28}
\]

Assumptions \((A1)\) and \((A2)\) indicate that \( \lim_{n \to \infty} \tilde{Y}_{k,j}(n) = \lambda_k F_{k,j}^c \). Again, by Scheffé’s lemma, \( (28) \) is equivalent to

\[
\lim_{n \to \infty} \sum_{j=0}^{\infty} \left| \tilde{Y}_{k,j}(n) - \lambda_k F_{k,j}^c \right| = 0. \tag{29}
\]

For any \( \epsilon > 0 \), since \( \sum_{j=0}^{\infty} \lambda_k F_{k,j}^c \) converges for each \( k = 0, 1, \ldots, d - 1 \), we know that there exists \( J \), such that

\[
\sum_{k=0}^{d-1} \sum_{j=0}^{\infty} \lambda_k F_{k,j}^c < \epsilon.
\]

Let \( N_4 \equiv \lfloor J/d \rfloor \) where \( \lfloor x \rfloor \) means the smallest integer that greater than \( x \), then when \( n \geq N_4 \), we have \( nd > J \). By \( (29) \), there exists \( N_5 \) such that when \( n > N_5 \),

\[
\sum_{k=0}^{d-1} \sum_{j=0}^{\infty} \left| \tilde{Y}_{k,j}(n) - \lambda_k F_{k,j}^c \right| < \epsilon.
\]

And let \( N_6 \equiv \max\{N_4, N_5\} \), then when \( n \geq N_6 \),

\[
\sum_{k=0}^{d-1} \sum_{j=N_6d}^{\infty} \tilde{Y}_{k,j}(n) \leq 2\epsilon. \tag{30}
\]
And let $N_7 \equiv \lceil N_6 / \epsilon \rceil$, then when $n > N_7$, we have $(n - N_6)/n > 1 - \epsilon$. Finally, when $n > \max\{N_7, 2N_6\}$,

$$
\bar{E}(n) = \frac{1}{n} \sum_{m=1}^{n} \sum_{j=1}^{d} \sum_{s=(n-m)d}^{\infty} Y_{d-j+(m-1)d,j+s} \\
= \sum_{j=1}^{d} \frac{1}{n} \sum_{m=1}^{n-N_6} \sum_{s=(n-m)d}^{\infty} Y_{d-j+(m-1)d,j+s} + \sum_{j=1}^{d} \frac{1}{n} \sum_{m=n-N_6+1}^{n} \sum_{s=(n-m)d}^{\infty} Y_{d-j+(m-1)d,j+s} \\
\leq \sum_{j=1}^{d} \frac{1}{n} \sum_{m=1}^{n-N_6} \sum_{s=N_6d}^{\infty} Y_{d-j+(m-1)d,j+s} + \sum_{j=1}^{d} \frac{1}{n} \sum_{m=n-N_6+1}^{n} \sum_{s=0}^{\infty} Y_{d-j+(m-1)d,s} \\
= \sum_{j=1}^{d} \sum_{s=N_6d}^{\infty} \bar{Y}_{d-j,j+s}(n) + \sum_{j=1}^{d} \sum_{s=0}^{\infty} (\bar{Y}_{d-j,s}(n) - \frac{n-N_6}{n} \bar{Y}_{d-j,s}(n-N_6)) \\
\leq 2\epsilon + 2\epsilon + \epsilon \left( \sum_{k=0}^{d-1} \sum_{j=0}^{\infty} \lambda_k F_{k,j}^c + \epsilon \right). \tag{31}
$$

The inequality in the third line is just relaxing the index $s$, the next inequality is relaxing the index $m$ for the first term, and the last inequality is given by the definition of $N_6$ and $N_7$, where the first term is bounded by $2\epsilon$ by using (30), and the second term is bounded by

$$
\sum_{j=1}^{d} \sum_{s=0}^{\infty} (\bar{Y}_{d-j,s}(n) - \frac{n-N_6}{n} \bar{Y}_{d-j,s}(n-N_6)) \\
\leq \sum_{j=1}^{d} \sum_{s=0}^{\infty} (\bar{Y}_{d-j,s}(n) - (1-\epsilon) \bar{Y}_{d-j,s}(n-N_6)) \\
= \sum_{j=1}^{d} \sum_{s=0}^{\infty} (\bar{Y}_{d-j,s}(n) - \bar{Y}_{d-j,s}(n-N_6)) + \epsilon \bar{Y}_{d-j,s}(n-N_6) \\
\leq 2\epsilon + \epsilon \left( \sum_{k=0}^{d-1} \sum_{j=0}^{\infty} \lambda_k F_{k,j}^c + \epsilon \right). \tag{32}
$$

So we have proved that $L_k \equiv \lim_{n \to \infty} \bar{Q}_k(n) = \lim_{n \to \infty} \bar{L}_k(n)$, which completes the proof of Theorem 1.

4.3. Proof of Second Half of Corollary 1

Proof: Here we give the second half of the proof of Corollary 1, i.e., the explicit bound in (9).
From equation (26), we know that

\[ \bar{E}_k(n) = \frac{1}{n} \sum_{m=1}^{n} \sum_{j=1}^{d} \sum_{s=n-m+1}^{\infty} Y_{d-j+(m-1)d, j+k+(s-1)d}. \]

By the condition, we know that \( Y_{i,j} = 0 \) for \( i \geq 0 \) and \( j \geq md_u \). Hence, when \( n \geq m_u \),

\[ \bar{E}_k(n) \leq \frac{1}{n} \sum_{j=1}^{d} \sum_{m=n-m_u}^{n} \sum_{s=n-m+1}^{m_u+1} A_{d-j+(m-1)d} \leq \frac{1}{n} \sum_{j=1}^{d} \sum_{m=n-m_u}^{n} \sum_{s=n-m+1}^{m_u+1} \lambda_u \]

\[ = \frac{1}{n} d (m_u + 1)(m_u + 2) \leq \lambda_u d(m_u + 2)^2. \quad (33) \]

So we have proved Corollary 1.

5. Conclusions

In §2 and §3 we have established sample-path and stationary versions of a periodic Little’s law (PLL), which we think can add insight into the performance of periodic stochastic models, which are natural for many service systems. In particular, these new theorems explain the extraordinary model fit we found in our data anlysis of an emergency department in Whitt and Zhang (2017b), as shown in Figure 1. Nevertheless, in §3.4 we present additional evidence supporting the infinite-server model proposed in Whitt and Zhang (2017b).

There are many directions for future research. We ourselves have already established a central limit theorem (CLT) version of the PLL in Whitt and Zhang (2017a), which parallels the CLT versions of Little’s law in Glynn and Whitt (1986, 1987, 1988) and Whitt (2012); these have important statistical applications as in Glynn and Whitt (1989b), Kim and Whitt (2013).

There may well be more related good theory to develop associated with periodic Palm measures and their application to queues, especially in discrete time, supplementing Section 1.7.4 of Baccelli and Bremaud (1994), Miyazawa and Takahashi (1992) and Whitt (1983).

As noted in El-Taha and Stidham (1999), Glynn and Whitt (1989a) and Whitt (1991), there are many important generalizations of Little’s law, such as the relation \( H = \lambda G \). It remains to establish such results in a periodic setting.
Acknowledgments

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References


(See Correction Note on $L = \lambda W$, Queueing Systems, 12 (4), 1992, 431-432. The results are correct; minor but important change needed in proofs.).


Whitt, W. 1992. Correction note on $L = \lambda W$. *Queueing Systems* **12** 431–432. The results in the previous papers are correct, but minor important changes are needed in some proofs.


Wolff, R. W., Y.-C. Yao. 2014. Little's law when the average waiting time is infinite. *Queueing Systems* 76, 267–281.
Review of the Emergency Department Data Analysis

In this e-companion we review the data analysis in Whitt and Zhang (2017b), which led to Figure 1, and present results of further tests, all of which provide background for §3.4 in the main paper.

EC.1. General Background

Our work exploited the fact that Armony et al. (2015) made their hospital data publicly available. The data is from the large Rambam Hospital located in Haifa, Israel. The hospital has approximately 1000 beds, with about 40 belonging to the ED. In our analysis, we focus on the Emergency Internal Medical Unit, which is the largest division of the ED, accounting for 60% of the patients visiting the ED. Also the unit has independent resources, so it is reasonable to focus on it alone. For convenience, we refer to the Emergency Internal Medical Unit as the ED.

The data contains the arrival time and departure time of each patient that visited the ED from Jan. 2004 to Oct. 2007, but we only used data from the 25 weeks between December 2004 to May 2005. It is important to understand the definition of the departure time. The departure time is the time when an admission decision is made in the ED for the patient, i.e. whether to release the patient from the ED or to admit the patient to an Internal Ward (another department of the hospital) from the ED. The length of stay (LoS) of each patient is the length of time between arrival and departure, as defined above. Thus, the LoS does not include the important boarding time (the time between an admission decision to admit and receiving a bed within the Internal Ward).

During the 25-week study period, 23,409 patients visited the ED, which is about 134 patients per day. The mean LoS was about 4 hours, while the longest LoS was less than a week. Most of the patients only stayed a few hours in the ED. Hence, the boundedness assumption in Corollary 1 is appropriate in the present setting.

After carefully analyzing the data, we proposed an $M_t^T/GI_1/\infty$ queueing system as an aggregate model for the patient flow of the ED. We also conducted simulations to validate our model. We will briefly review the modeling and analysis in the following subsections. The details and supporting materials is presented in Whitt and Zhang (2017b).
EC.2. The Model Components

EC.2.1. The Arrival Process

We confirmed the observation in Armony et al. (2015) that the arrival rate is time-varying, but with significant day-to-day variation. We concluded that the arrival rate function can be regarded as periodic over a week. We observed over-dispersion in the arrival process compared to the Poisson process; i.e., the dispersion (ratio of the estimated variance $\hat{\sigma}^2$ to the estimated mean $\hat{\mu}$) of the daily number of arrivals is significantly larger than 1. Hence, we proposed a two-time-scale model for the arrival process, in which we model the daily totals first and then describe the process within a day.

We used a linear model to model the successive daily totals. We assumed that the daily totals are determined by

$$T(d_w) = A + C \times d_w + \epsilon,$$

where $T(d_w)$ represents the daily total, $A, C$ are constant coefficients, $d_w$ is a factor (qualitative) variable indicating the day-of-week and $\epsilon \sim N(0, \sigma^2)$ is a Gaussian distributed random error. We fit the model to the $25 \times 7 = 175$ daily totals. The statistical analysis indicated a good fit, with the Gaussian residual assumption being supported. The regression coefficients are given in Table EC.1. The estimated variance and dispersion were $\hat{\sigma}^2 = 202$ and $\hat{\sigma}^2 / \hat{\mu} = 202 / 134 \approx 1.50$.

We applied statistical tests to test the assumption of a nonhomogeneous Poisson process within days, using the tests described in Kim and Whitt (2014b,a). We concluded that the within-day arrivals (given the daily totals) fit a nonhomogeneous Poisson process well.

EC.2.2. The Admission Decision

Let $p(t)$ be the probability that a patient arriving at time $t$ is admitted to the Internal Ward from the ED. Our data analysis indicated that this probability also is time-varying, thus we include the arrival time $t$ as a parameter. In our model we assume that the successive admission decisions are mutually independent Bernoulli random variables, with a probability that is determined by the
arrival time. We concluded that the function $p(t)$ can be regarded as periodic with a period length of one day. The admission probability is higher in the daytime than it is in the night. We found that a truncated quadratic function provides a good fit.

**EC.2.3. The Length of Stay**

We found that the patient LoS given their arrival time should also be modelled as time-varying. Indeed, the corresponding $M^T_t/GI/\infty$ model with i.i.d. service times failed to predict the number of patients in the system. Because it is difficult to estimate continuously changing distributions, we assumed that the LoS distribution is fixed within each hour but can vary from hour to hour. Similar to the arrival process, we let the LoS distributions be periodic with a length of one week. We considered the LoS for admitted and non-admitted patients separately. The model assumes that each new LoS is independent of the system state upon arrival, having a distribution that depends only on the hour of arrival (within the week) and the admission decision.

**EC.3. Model Fitting and Simulations**

In summary, we fitted our model as follows:

1. Fit the Gaussian model (EC.1) using daily total arrivals data of 25 weeks. The fitting result is shown in Table EC.1.

<table>
<thead>
<tr>
<th>Coefficients</th>
<th>Estimate</th>
<th>SE</th>
</tr>
</thead>
<tbody>
<tr>
<td>A</td>
<td>133.766</td>
<td>2.842</td>
</tr>
<tr>
<td>C.Sun</td>
<td>28.234</td>
<td>4.019</td>
</tr>
<tr>
<td>C.Mon</td>
<td>14.634</td>
<td>4.019</td>
</tr>
<tr>
<td>C.Tue</td>
<td>11.274</td>
<td>4.019</td>
</tr>
<tr>
<td>C.Wed</td>
<td>-1.366</td>
<td>4.019</td>
</tr>
<tr>
<td>C.Thu</td>
<td>-0.526</td>
<td>4.019</td>
</tr>
<tr>
<td>C.Fri</td>
<td>-23.886</td>
<td>4.019</td>
</tr>
<tr>
<td>C.Sat</td>
<td>-28.366</td>
<td>4.019</td>
</tr>
</tbody>
</table>

*Table EC.1* Estimated regression coefficients for the single-factor model in (EC.1).*
2. Get the empirical hourly arrival rate function in a week view by combining data of 25 weeks.

3. Fit a quadratic function ($\hat{p}(t)$) to the empirical admission rate (hourly proportion of admitted patients) $p(t)$, which is computed by combining all the days, and make it periodic with period 1 day. $\hat{p}(t) = -0.001082(x - 13.5)^2 + 0.451996$, where $x = ((t - 1.5) \mod 24) + 1.5$ and $t \in [0, 24]$.

4. Get the hourly empirical LoS distributions for a week by combining 25 weeks data for each of the 2 groups.

Then we conducted simulations to test our model as follows:

1. Generate i.i.d. samples of the daily numbers of arrivals for a week and 5 days before the week from the fitted Gaussian model (EC.1).

2. Simulate the arrival times of patients within each day given the daily total arrivals (say $N$) by generating $N$ i.i.d. samples of random variables with a density proportional to the empirical arrival rate function.

3. Independently determine the admission decision for each patient upon arrival (say $t$ being the arrival epoch) using independent Bernoulli random variables with mean $\hat{p}(t)$.

4. Independently draw a sample of LoS from the corresponding empirical LoS distribution which depends on the arrival time and admission decision for each patient.

5. Compute the average occupancy level for each hour of the week.

6. Repeat the above procedure 1000 times. (Use 1000 i.i.d. replications.)

As we can see in Figure 1, the estimated mean number of patients from the model simulation coincides almost perfectly with the empirical directly from the data. We think that the sample-path PLL provides a good explanation.

**EC.4. Testing the Assumptions of the Sample-Path PLL**

As we can see in Figure 1, the estimated mean number of patients from the model simulation coincides almost perfectly with the empirical directly from the data. We think that the PLL provides a good explanation. However, to support that conclusion, we should investigate to what extent the data satisfies the assumptions in the PLL theorems.
First, we investigate Assumptions (A1), (A2) and (A3) in the sample-path PLL in Theorem 1. Toward that end, Figure EC.1 shows the convergence of the arrival rate, the empirical cumulative distribution functions (empirical cdfs) and the mean LoS as the sample size increases. In particular, the top figure shows the estimated hourly arrival rate as a function of the sample size (the number of weeks), increasing from 1 to 25, with the shading becoming darker as the sample size increases. Similarly, the two figures in the middle show the empirical cdfs at 10 a.m. and 4 p.m. as a function of the sample size, while the bottom two figures show the average LoS at 10 a.m. and 4 p.m. as as a function of the sample size, increasing from 1 week to 25. All five plots provide empirical support for the assumptions.

Our model actually used these estimators. The simulation used arrivals from a nonhomogeneous Poisson process given daily total arrivals according to the empirical arrival rate. The daily total arrivals came from the Gaussian model and so are also unbiased. So our model replicated the limiting arrival rate. Then we sampled the LoS of each patient based on the arrival time from the corresponding empirical distribution. By law of large numbers, the average occupancy level from our simulation converges to \( \bar{L}(25) \), which by Theorem 1 is close to \( \bar{Q}(25) \).

We test the independence assumptions in Proposition 1 in §3.4.
Figure EC.1  Top: the estimated hourly arrival rate as a function of the sample size (number of weeks), ranging from 1 to 25, with the shading becoming darker as the sample size increases; Middle: the estimated LoS empirical cdfs at 10 a.m. and 4 p.m. as a function of the sample size; Bottom: the average LoS at 10 a.m. and 4 p.m. as a function of the sample size.