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Research Article

Outage Analysis of Partial Relay Selection Schemes with Feedback Delay and Channel Estimation Errors in Nonidentical Rayleigh Fading Channels

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This paper investigates the impact of the channel estimation error and outdated channel state information (CSI) on the outage performances of partial relay selection (PRS) and efficient partial relay selection (EPRS). Considering imperfect channel estimation and outdated CSI with decode-and-forward (DF) relaying strategy, closed-form expressions for exact outage probabilities and asymptotic outage probabilities for PRS and EPRS are provided assuming independent and nonidentically distributed Rayleigh fading channels. Numerical investigations verify the analytical expression for outage probability and show how much performance is degraded by the channel estimation errors and the feedback delay that causes the outdated CSI.

1. Introduction

Cooperative relaying has attracted great attention because it offers an excellent performance at low cost by forming a virtual multiple-input multiple-output (MIMO) system using all available nodes as relays. In industry, the cooperative relaying system has been included in both worldwide interoperability for microwave access (WiMAX) [1] and long-term evolution (LTE) [2] to extend the cell coverage and improve the cell-edge user throughput. Among various cooperative diversity techniques, the best relay selection (BRS) is one of the most promising schemes since it can achieve full spatial diversity from multiple relays with low complexity [3], even in the presence of interference [4]. To further reduce overhead and complexity of BRS, partial relay selection (PRS) was introduced in [5], in which a single relay is selected based on only the first-hop channel state information (CSI). Hence, PRS can prolong lifetime of energy-constrained relay node. Despite such advantages, achievable performance of PRS is severely bounded because partial CSI cannot sufficiently represent the end-to-end channel quality [6]. To improve the performance of PRS with a small additional overhead, efficient partial relay selection (EPRS) was introduced in [7]. In EPRS, a link with the smaller average channel power between the first and the second hops is chosen at each end-to-end path, and the CSI for the links chosen at every end-to-end path is used for a single relay selection. Since the CSI used for EPRS is more correlated with the end-to-end channel quality than for PRS, EPRS can attain better performance than PRS [7, 8].

The assumption of perfect CSI for selecting a single relay and decoding a received signal may be impractical because the acquired CSI generally contains estimation errors because of noise, and also it can be outdated because of a feedback delay time. Therefore, performance evaluation and system design considering the impacts of both channel estimation errors and outdated CSI may be critical and necessary to satisfy performance requirements in practical communication environments. In [9], the outage probability and the average error rate of decode-and-forward (DF) relaying systems with BRS were investigated under identically distributed Rayleigh fading channels in the presence of both feedback delay and

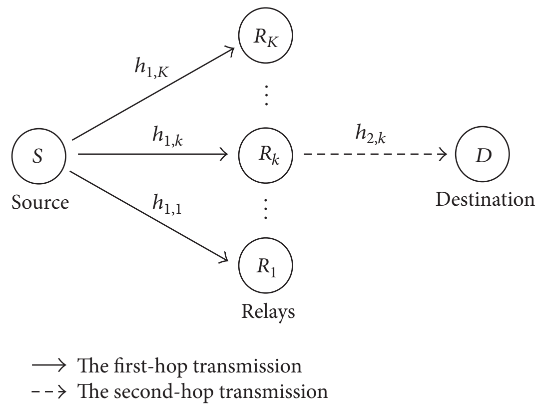

Figure 1: Dual-hop relaying system using relay selection where relay k is selected.

channel estimation errors. In [10, 11], the outage probability and the average error rate of amplify-and-forward relaying systems with PRS and BRS were presented for identically distributed Rayleigh fading channels considering feedback delay. In [9–11], fading channels for all the links of each hop are assumed to be statistically identical. However, when relays are distributed over a sufficiently large area, the assumption of identically distributed fading channels can be impractical. In this paper, hence, we assume nonidentically distributed fading channels for all the links. In addition, to the best of our knowledge, outage performances of PRS and EPRS in DF relaying systems have not been studied in the presence of both feedback delay and channel estimation errors. Therefore, this paper focuses on investigation into the impact of both feedback delay and the channel estimation errors on the outage performances of PRS and EPRS in DF relaying systems. Considering outdated CSI and imperfect channel estimation with a DF relaying strategy, exact outage probabilities and asymptotic outage probabilities for PRS and EPRS are provided in closed form under independent and nonidentically distributed Rayleigh fading channels. Numerical investigations verify that the analytic results are perfectly matched with the simulated ones and show how much performance is degraded by the channel estimation errors and the feedback delay that induces the outdated CSI.

2. System Model

We consider a dual-hop DF relaying system using relay selection that consists of a source, K relays, and a destination, as shown in Figure 1. All nodes are equipped with a single antenna. In this paper, we assume that the direct link between the source and the destination is unavailable due to high path loss and shadowing effect. The DF relays operate in a half-duplex mode, and PRS and EPRS are employed to select a single relay. As a relay selection protocol, *proactive protocol* [3] is considered, where a single relay is selected before data transmission. Because only the selected relay tries to receive and decode the signal from the source, the proactive protocol is more energy-efficient than *reactive protocol*, in which a single relay is selected after decoding the received signal at all relays. In the proactive protocol with DF relaying, the selected relay reencodes the decoded signal and forwards it to the destination only when the decoding succeeds.

Let the channel coefficient for relay k ($k \in \{1, \dots, K\}$) at hop i ($i \in \{1, 2\}$) be denoted as $h_{i,k}$, where the channels for relay k at the first and the second hops mean those between the source and relay k and between relay k and the destination, respectively. The channels $h_{i,k}$'s are assumed to be independent and nonidentically distributed complex Gaussian random variables with zero mean and variance $\beta_{i,k}$; that is, $h_{i,k} \sim \mathcal{CN}(0, \beta_{i,k})$. Also, considering block fading, the channels are assumed to be constant during a block length.

2.1. Channel Estimation Errors and Feedback Delay. The source and the relays transmit training signals orthogonally, and then the minimum-mean-square-error- (MMSE-) estimated channels at the relays and the destination are given by [12–14]

$$h_{i,k} = \hat{h}_{i,k} + e_{i,k}, \tag{1}$$

where $e_{i,k} \sim \mathcal{CN}(0, \sigma_e^2)$ denotes the estimation error and $\hat{h}_{i,k} \sim \mathcal{CN}(0, \beta_{i,k} - \sigma_e^2)$ represents the MMSE-estimated channel. In this paper, we assume that the channel estimation error is independent of the received signal-to-noise ratio (SNR) for a data signal, and the estimation error distribution is identical at all the relays and the destination.

Since the estimated CSI is fed back for relay selection, the CSI may be outdated when the source and the selected relay transmit a data signal. Letting $\hat{\beta}_{i,k} \triangleq \beta_{i,k} - \sigma_e^2$, the relation between the previously estimated channel $\hat{h}_{i,k}$ and the currently estimated channel $\tilde{h}_{i,k}$ can be modelled as [13, 15]

$$\tilde{h}_{i,k} = \rho_d \hat{h}_{i,k} + \sqrt{1 - \rho_d^2} u_{i,k}, \tag{2}$$

where $u_{i,k} \sim \mathcal{CN}(0, \hat{\beta}_{i,k})$, $\tilde{h}_{i,k} \sim \mathcal{CN}(0, \hat{\beta}_{i,k})$, and ρ_d denotes the correlation coefficient between $\hat{h}_{i,k}$ and $\tilde{h}_{i,k}$, which can be defined as $\rho_d \triangleq J_0(2\pi f_d \tau)$ [16]. $J_0(\cdot)$ is a Bessel function of the first kind of zero order, f_d is the maximum Doppler frequency, and τ is a feedback delay time. It is noted that $\hat{h}_{i,k}$ and $\tilde{h}_{i,k}$ are the channels used for relay selection and decoding, respectively. Hereafter, let $\hat{g}_{i,k} \triangleq |\hat{h}_{i,k}|^2$ and $\tilde{g}_{i,k} \triangleq |\tilde{h}_{i,k}|^2$.

2.2. Partial Relay Selection Schemes. For PRS, the first-hop CSI is used to select a single relay. Thus, each relay feeds back the estimated CSI for the first hop to the source. Then, the source broadcasts the index of a selected relay to all the relays. The relay selected by PRS is expressed as [5]

$$k_p^* = \arg \max_{k=1,\dots,K} \left\{ \hat{g}_{1,k} \right\}. \tag{3}$$

For EPRS, the first-hop estimated CSI for relay k is used if $\hat{\beta}_{1,k} < \hat{\beta}_{2,k}$; otherwise the second-hop estimated CSI for relay k is used. Every relay can know the average channel powers for the first and the second hops using the received signals from the source and the destination. Thus, each relay determines either to feed back the estimated CSI for the first hop to the destination or transmit the training signals to the destination for the second-hop channel estimation. Then, the destination selects a single relay and broadcasts the index of a selected relay to all the relays. The relay selected by EPRS is expressed as [7]

$$k_e^* = \arg\max_{k=1,\ldots,K} \{\widehat{w}_k\}, \quad (4)$$

where $\widehat{w}_k = \widehat{g}_{1,k}$ for $\widehat{\beta}_{1,k} < \widehat{\beta}_{2,k}$; otherwise $\widehat{w}_k = \widehat{g}_{2,k}$.

3. Outage Performance Analysis

3.1. Received SNR with Channel Estimation Error. In this paper, we assume equal transmit powers at the source and the relays and equal noise powers at the relays and the destination, denoted as P and σ_n^2, respectively, but it is straightforward to extend to different transmit powers. When the source or the selected relay (that succeeds in decoding) transmits x, the received signal at the relay or the destination is given by

$$r_{i,k} = h_{i,k}\sqrt{P}x + n_{i,k} = \widetilde{h}_{i,k}\sqrt{P}x + e_{i,k}^d\sqrt{P}x + n_{i,k}, \quad (5)$$

where $E[|x|^2] = 1$, $n_{i,k} \sim \mathcal{CN}(0,\sigma_n^2)$ is an additive white Gaussian noise, and $e_{i,k}^d \sim \mathcal{CN}(0,\sigma_e^2)$ denotes the error of MMSE-estimated channel in decoding. Let $\rho_t \triangleq P/\sigma_n^2$ and $\zeta \triangleq \rho_t/(1+\rho_t\sigma_e^2)$, where ρ_t denotes the average transmit SNR. Then the received SNR with channel estimation error is obtained as $\widetilde{\gamma}_{i,k} = \zeta\widetilde{g}_{i,k}$, where it is noted that $\widetilde{h}_{i,k}$ is used instead of $\widehat{h}_{i,k}$ because of the received signal for decoding, not relay selection. It is also noted that there is no received signal at the destination when the selected relay fails to decode the signal received from the source.

3.2. Exact Outage Probability Analysis. An outage probability expression of PRS can be easily obtained from that of EPRS. Thus, in this paper, we show only derivations of outage probability of EPRS.

Using [17, equations (2) and (3)] and (4), for a given target data rate R in bps/Hz, the outage probability of EPRS in dual-hop DF relaying systems in the presence of channel estimation error and feedback delay is obtained by

$$\begin{aligned} P_o(R) = &\sum_{k=1}^{K} \Pr\left\{\frac{1}{2}\log_2\left(1+\widetilde{\gamma}_{2,k}\right) < R, \frac{1}{2}\log_2\left(1+\widetilde{\gamma}_{1,k}\right) > R,\ \widehat{w}_k > \max_{\substack{j=1,\ldots,K\\ j\neq k}}\{\widehat{w}_j\}\right\} \\ &+ \sum_{k=1}^{K} \Pr\left\{\frac{1}{2}\log_2\left(1+\widetilde{\gamma}_{1,k}\right) < R,\ \widehat{w}_k > \max_{\substack{j=1,\ldots,K\\ j\neq k}}\{\widehat{w}_j\}\right\}, \end{aligned} \quad (6)$$

where it is assumed that the decoding at relay k succeeds when the achievable data rate between the source and relay k exceeds R [17]. In (6), the first and the second parts mean the outage probability in case of decoding success and failure at the selected relay, respectively. Letting $R_t \triangleq (2^{2R}-1)/\zeta$, (6) can be rewritten as

$$\begin{aligned} P_o(R) &= \sum_{k=1}^{K} \Pr\left\{\widetilde{g}_{2,k} < R_t,\ \widetilde{g}_{1,k} > R_t,\ \widehat{w}_k > \max_{\substack{j=1,\ldots,K\\ j\neq k}}\{\widehat{w}_j\}\right\} \\ &+ \sum_{k=1}^{K} \Pr\left\{\widetilde{g}_{1,k} < R_t,\ \widehat{w}_k > \max_{\substack{j=1,\ldots,K\\ j\neq k}}\{\widehat{w}_j\}\right\}. \end{aligned} \quad (7)$$

If $\widehat{w}_k = \widehat{g}_{1,k}$, then the first and the second parts in (7) are, respectively, expressed as

$$\Pr\left\{\widetilde{g}_{2,k} < R_t\right\}\Pr\left\{\widetilde{g}_{1,k} > R_t,\ \widehat{g}_{1,k} > \max_{\substack{j=1,\ldots,K\\ j\neq k}}\{\widehat{w}_j\}\right\}, \quad (8)$$

$$\Pr\left\{\widetilde{g}_{1,k} < R_t,\ \widehat{g}_{1,k} > \max_{\substack{j=1,\ldots,K\\ j\neq k}}\{\widehat{w}_j\}\right\}. \quad (9)$$

On the other hand, if $\widehat{w}_k = \widehat{g}_{2,k}$, then the first and the second parts in (7) are expressed as

$$\Pr\left\{\widetilde{g}_{1,k} > R_t\right\}\Pr\left\{\widetilde{g}_{2,k} < R_t,\ \widehat{g}_{2,k} > \max_{\substack{j=1,\ldots,K\\ j\neq k}}\{\widehat{w}_j\}\right\}, \quad (10)$$

$$\Pr\left\{\widetilde{g}_{1,k} < R_t\right\}\Pr\left\{\widehat{g}_{2,k} > \max_{\substack{j=1,\ldots,K\\ j\neq k}}\{\widehat{w}_j\}\right\}. \quad (11)$$

Conditioned on $\widehat{g}_{i,k} = z$, $\widetilde{g}_{i,k}$ is a noncentral Chi-square distributed random variable with two degrees of freedom, and its probability density function (PDF) is obtained using [18, eq. (2-1-118)] as follows:

$$\begin{aligned} \Pr\left\{\widetilde{g}_{i,k} = x \mid \widehat{g}_{i,k} = z\right\} &= \frac{1}{\widehat{\beta}_{i,k}(1-\rho_d^2)} \\ &\cdot e^{-(\rho_d^2 z + x)/(\widehat{\beta}_{i,k}(1-\rho_d^2))} I_0\left(\sqrt{\frac{4\rho_d^2 z x}{\widehat{\beta}_{i,k}^2(1-\rho_d^2)^2}}\right), \end{aligned} \quad (12)$$

where $I_0(\cdot)$ is a modified Bessel function of the first kind of order zero. Using [18, eq. (2-1-120)], (12) is rewritten as

$$\Pr\{\tilde{g}_{i,k} = x \mid \hat{g}_{i,k} = z\} = \frac{1}{\hat{\beta}_{i,k}(1-\rho_d^2)} \cdot e^{-(\rho_d^2 z + x)/(\hat{\beta}_{i,k}(1-\rho_d^2))} \sum_{m=0}^{\infty} \frac{1}{(m!)^2} \left(\frac{\rho_d^2 z x}{\hat{\beta}_{i,k}^2 (1-\rho_d^2)^2} \right)^m. \quad (13)$$

Using (13), the cumulative distribution function (CDF) and complementary CDF of $\tilde{g}_{i,k}$ conditioned on $\hat{g}_{i,k} = z$ are, respectively, obtained by

$$\begin{aligned}
&\Pr\{\tilde{g}_{i,k} < x \mid \hat{g}_{i,k} = z\} \\
&= \int_0^x \Pr\{\tilde{g}_{i,k} = y \mid \hat{g}_{i,k} = z\} dy = \frac{1}{\hat{\beta}_{i,k}(1-\rho_d^2)} \\
&\cdot e^{-\rho_d^2 z/(\hat{\beta}_{i,k}(1-\rho_d^2))} \sum_{m=0}^{\infty} \frac{1}{(m!)^2} \left(\frac{\rho_d^2 z}{\hat{\beta}_{i,k}(1-\rho_d^2)} \right)^m \\
&\cdot \int_0^x \left(\frac{y}{\hat{\beta}_{i,k}(1-\rho_d^2)} \right)^m e^{-y/(\hat{\beta}_{i,k}(1-\rho_d^2))} dy \\
&= e^{-\rho_d^2 z/(\hat{\beta}_{i,k}(1-\rho_d^2))} \sum_{m=0}^{\infty} \frac{1}{(m!)^2} \left(\frac{\rho_d^2 z}{\hat{\beta}_{i,k}(1-\rho_d^2)} \right)^m \\
&\cdot \gamma\left(m+1, \frac{x}{\hat{\beta}_{i,k}(1-\rho_d^2)}\right),
\end{aligned} \quad (14)$$

$$\begin{aligned}
&\Pr\{\tilde{g}_{i,k} > x \mid \hat{g}_{i,k} = z\} \\
&= \int_x^{\infty} \Pr\{\tilde{g}_{i,k} = y \mid \hat{g}_{i,k} = z\} dy = \frac{1}{\hat{\beta}_{i,k}(1-\rho_d^2)} \\
&\cdot e^{-\rho_d^2 z/(\hat{\beta}_{i,k}(1-\rho_d^2))} \sum_{m=0}^{\infty} \frac{1}{(m!)^2} \left(\frac{\rho_d^2 z}{\hat{\beta}_{i,k}(1-\rho_d^2)} \right)^m \\
&\cdot \int_x^{\infty} \left(\frac{y}{\hat{\beta}_{i,k}(1-\rho_d^2)} \right)^m e^{-y/(\hat{\beta}_{i,k}(1-\rho_d^2))} dy \\
&= e^{-\rho_d^2 z/(\hat{\beta}_{i,k}(1-\rho_d^2))} \sum_{m=0}^{\infty} \frac{1}{(m!)^2} \left(\frac{\rho_d^2 z}{\hat{\beta}_{i,k}(1-\rho_d^2)} \right)^m \\
&\cdot \left(m! - \gamma\left(m+1, \frac{x}{\hat{\beta}_{i,k}(1-\rho_d^2)}\right) \right),
\end{aligned} \quad (15)$$

where $\gamma(\cdot,\cdot)$ denotes the incomplete gamma function and we used $\int_x^{\infty} t^{m-1} e^{-t} dt = (m-1)! - \gamma(m, x)$ to derive the complementary CDF.

Let $\hat{y}_k \triangleq \max_{j=1,\ldots,K, j\neq k}\{\hat{w}_j\}$. Then, the CDF of $\hat{y}_k$ is obtained by

$$\begin{aligned}
\Pr\{\hat{y}_k < x\} &= \prod_{j=1, j\neq k}^{K} \Pr\{\hat{w}_j < x\} = \prod_{j=1, j\neq k}^{K} \left(1 - e^{-x/\lambda_j}\right) \\
&= 1 + \sum_{j=1}^{K-1} \sum_{L_j \subseteq S_k} (-1)^j e^{-x \sum_{q\in L_j} 1/\lambda_q},
\end{aligned} \quad (16)$$

where $S_k = \{1, \ldots, k-1, k+1, \ldots, K\}$, L_j represents all possible subsets of S_k with the cardinality of j, and $\lambda_q = \hat{\beta}_{1,q}$ if $\hat{\beta}_{1,q} < \hat{\beta}_{1,q}$; otherwise $\lambda_q = \hat{\beta}_{2,q}$.

Using (15), (16), and $\Pr\{\hat{g}_{i,k} = z\} = (1/\hat{\beta}_{i,k}) e^{-z/\hat{\beta}_{i,k}}$, $\Pr\{\tilde{g}_{i,k} > R_t, \hat{g}_{i,k} > \max_{j=1,\ldots,K, j\neq k}\{\hat{w}_j\}\}$ in (8) is derived as follows:

$$\begin{aligned}
&\int_0^{\infty} \Pr\{\tilde{g}_{i,k} > R_t \mid \hat{g}_{i,k} = z\} \Pr\{\hat{y}_k < z \mid \hat{g}_{i,k} = z\} \Pr\{\hat{g}_{i,k} = z\} dz \\
&= \int_0^{\infty} \left[e^{-\rho_d^2 z/(\hat{\beta}_{i,k}(1-\rho_d^2))} \sum_{m=0}^{\infty} \frac{1}{(m!)^2} \left(\frac{\rho_d^2 z}{\hat{\beta}_{i,k}(1-\rho_d^2)} \right)^m \right. \\
&\cdot \left(m! - \gamma\left(m+1, \frac{R_t}{\hat{\beta}_{i,k}(1-\rho_d^2)}\right) \right) \\
&\cdot \left. \left(1 + \sum_{j=1}^{K-1} \sum_{L_j \subseteq S_k} (-1)^j e^{-z \sum_{q\in L_j} 1/\lambda_q} \right) \frac{1}{\hat{\beta}_{i,k}} e^{-z/\hat{\beta}_{i,k}} \right] dz \\
&= \sum_{m=0}^{\infty} \left(m! - \gamma\left(m+1, \frac{R_t}{\hat{\beta}_{i,k}(1-\rho_d^2)}\right) \right) \\
&\cdot \frac{(\rho_d^2)^m (1-\rho_d^2)}{m!} \\
&+ \sum_{j=1}^{K-1} \sum_{L_j \subseteq S_k} \sum_{m=0}^{\infty} \left(m! - \gamma\left(m+1, \frac{R_t}{\hat{\beta}_{i,k}(1-\rho_d^2)}\right) \right) \\
&\cdot \frac{(-1)^j}{m!} \left(\frac{\rho_d^2}{1-\rho_d^2} \right)^m \left(\frac{1}{1-\rho_d^2} + \hat{\beta}_{i,k} \sum_{q\in L_j} \frac{1}{\lambda_q} \right)^{-m-1}.
\end{aligned} \quad (17)$$

Analogous to (17), using (14), (16), and $\Pr\{\hat{g}_{i,k} = z\} = (1/\hat{\beta}_{i,k}) e^{-z/\hat{\beta}_{i,k}}$, $\Pr\{\tilde{g}_{i,k} < R_t,\ \hat{g}_{i,k} > \max_{j=1,\ldots,K, j\neq k}\{\hat{w}_j\}\}$ in (9) and (10) is derived as follows:

$$\begin{aligned}
&\int_0^{\infty} \Pr\{\tilde{g}_{i,k} < R_t \mid \hat{g}_{i,k} = z\} \Pr\{\hat{y}_k < z \mid \hat{g}_{i,k} = z\} \Pr\{\hat{g}_{i,k} = z\} dz \\
&= \int_0^{\infty} \left[e^{-\rho_d^2 z/(\hat{\beta}_{i,k}(1-\rho_d^2))} \sum_{m=0}^{\infty} \frac{1}{(m!)^2} \left(\frac{\rho_d^2 z}{\hat{\beta}_{i,k}(1-\rho_d^2)} \right)^m \right. \\
&\cdot \gamma\left(m+1, \frac{R_t}{\hat{\beta}_{i,k}(1-\rho_d^2)}\right)
\end{aligned}$$

$$\cdot\left(1+\sum_{j=1}^{K-1}\sum_{L_j\subseteq S_k}(-1)^j e^{-z\sum_{q\in L_j}1/\lambda_q}\right)\frac{1}{\widehat{\beta}_{i,k}}e^{-z/\widehat{\beta}_{i,k}}\Bigg]dz$$
$$=\sum_{m=0}^{\infty}\gamma\left(m+1,\frac{R_t}{\widehat{\beta}_{i,k}(1-\rho_d^2)}\right)\frac{(\rho_d^2)^m(1-\rho_d^2)}{m!}$$
$$+\sum_{j=1}^{K-1}\sum_{L_j\subseteq S_k}\sum_{m=0}^{\infty}\gamma\left(m+1,\frac{R_t}{\widehat{\beta}_{i,k}(1-\rho_d^2)}\right)\frac{(-1)^j}{m!}$$
$$\cdot\left(\frac{\rho_d^2}{1-\rho_d^2}\right)^m\left(\frac{1}{1-\rho_d^2}+\widehat{\beta}_{i,k}\sum_{q\in L_j}\frac{1}{\lambda_q}\right)^{-m-1}. \tag{18}$$

Then, using (16), $\Pr\{\widetilde{g}_{2,k} > \max_{j=1,\dots,K,j\neq k}\{\widetilde{w}_j\}\}$ in (11) is obtained as follows:

$$\int_0^{\infty}\Pr\left\{\widetilde{y}_k<z\mid\widetilde{g}_{2,k}=z\right\}\Pr\left\{\widetilde{g}_{2,k}=z\right\}dz$$
$$=\int_0^{\infty}\left(1+\sum_{j=1}^{K-1}\sum_{L_j\subseteq S_k}(-1)^j e^{-z\sum_{q\in L_j}1/\lambda_q}\right)$$
$$\cdot\frac{1}{\widehat{\beta}_{2,k}}e^{-z/\widehat{\beta}_{2,k}}dz=1 \tag{19}$$
$$+\sum_{j=1}^{K-1}\sum_{L_j\subseteq S_k}\frac{(-1)^j}{\widehat{\beta}_{2,k}}\left(\frac{1}{\widehat{\beta}_{2,k}}+\sum_{q\in L_j}\frac{1}{\lambda_q}\right)^{-1}.$$

Finally, we obtain the outage probability by substituting (17)–(19), $\Pr\{\widetilde{g}_{i,k} < R_t\} = 1 - e^{-R_t/\widehat{\beta}_{i,k}}$, and $\Pr\{\widetilde{g}_{i,k} > R_t\} = e^{-R_t/\widehat{\beta}_{i,k}}$ into (8)–(11) and inserting (8)–(11) into (7). Also, the outage probability of PRS can be easily obtained by inserting (17) and (18) with $\lambda_q = \widehat{\beta}_{1,q}$ into (8) and (9), respectively, and substituting (8) and (9) into (7).

3.3. Asymptotic Outage Probability Analysis. In this section, different asymptotic analysis is performed according to the condition of σ_e^2 because when $\sigma_e^2 = 0$ (i.e., perfect channel estimation), $R_t = (2^{2R}-1)/\rho_t$, whereas when $\sigma_e^2 \neq 0$, $R_t = (2^{2R}-1)(1+\rho_t\sigma_e^2)/\rho_t \approx (2^{2R}-1)\sigma_e^2$ by high SNR approximation. It is noted that when $\sigma_e^2 \neq 0$, R_t does not depend upon ρ_t.

When $\sigma_e^2 = 0$, using $R_t = (2^{2R}-1)/\rho_t$ and high SNR approximation, the following approximated equation is obtained:

$$\gamma\left(m+1,\frac{R_t}{\widehat{\beta}_{i,k}(1-\rho_d^2)}\right)$$
$$=\gamma\left(m+1,\frac{2^{2R}-1}{\rho_t\widehat{\beta}_{i,k}(1-\rho_d^2)}\right) \tag{20}$$
$$\overset{\rho_t\to\infty}{\approx}\frac{1}{m+1}\left(\frac{2^{2R}-1}{\rho_t\widehat{\beta}_{i,k}(1-\rho_d^2)}\right)^{m+1}.$$

Using (20), (17) and (18) are, respectively, approximated as

$$\sum_{m=0}^{\infty}(\rho_d^2)^m(1-\rho_d^2)+\sum_{j=1}^{K-1}\sum_{L_j\subseteq S_k}\sum_{m=0}^{\infty}(-1)^j\left(\frac{\rho_d^2}{1-\rho_d^2}\right)^m$$
$$\cdot\left(\frac{1}{1-\rho_d^2}+\widehat{\beta}_{i,k}\sum_{q\in L_j}\frac{1}{\lambda_q}\right)^{-m-1}, \tag{21}$$

$$\frac{2^{2R}-1}{\rho_t\widehat{\beta}_{i,k}}\left\{1+\left(\frac{1}{1-\rho_d^2}\right)\sum_{j=1}^{K-1}\sum_{L_j\subseteq S_k}(-1)^j\right.$$
$$\left.\cdot\left(\frac{1}{1-\rho_d^2}+\widehat{\beta}_{i,k}\sum_{q\in L_j}\frac{1}{\lambda_q}\right)^{-1}\right\}. \tag{22}$$

Using (21), (22), $\Pr\{\widetilde{g}_{i,k} < R_t\} \approx (2^{2R}-1)/(\rho_t\widehat{\beta}_{i,k})$, and $\Pr\{\widetilde{g}_{i,k} > R_t\} \approx 1$ by high SNR approximation, (8)–(11) are, respectively, approximated as

$$\left(\frac{2^{2R}-1}{\rho_t\widehat{\beta}_{2,k}}\right)\left\{\sum_{m=0}^{\infty}(\rho_d^2)^m(1-\rho_d^2)\right.$$
$$+\sum_{j=1}^{K-1}\sum_{L_j\subseteq S_k}\sum_{m=0}^{\infty}(-1)^j\left(\frac{\rho_d^2}{1-\rho_d^2}\right)^m \tag{23}$$
$$\left.\cdot\left(\frac{1}{1-\rho_d^2}+\widehat{\beta}_{1,k}\sum_{q\in L_j}\frac{1}{\lambda_q}\right)^{-m-1}\right\},$$

$$\left(\frac{2^{2R}-1}{\rho_t\widehat{\beta}_{1,k}}\right)\left\{1+\left(\frac{1}{1-\rho_d^2}\right)\right.$$
$$\left.\cdot\sum_{j=1}^{K-1}\sum_{L_j\subseteq S_k}(-1)^j\left(\frac{1}{1-\rho_d^2}+\widehat{\beta}_{1,k}\sum_{q\in L_j}\frac{1}{\lambda_q}\right)^{-1}\right\}, \tag{24}$$

$$\left(\frac{2^{2R}-1}{\rho_t\widehat{\beta}_{2,k}}\right)\left\{1+\left(\frac{1}{1-\rho_d^2}\right)\right.$$
$$\left.\cdot\sum_{j=1}^{K-1}\sum_{L_j\subseteq S_k}(-1)^j\left(\frac{1}{1-\rho_d^2}+\widehat{\beta}_{2,k}\sum_{q\in L_j}\frac{1}{\lambda_q}\right)^{-1}\right\}, \tag{25}$$

$$\left(\frac{2^{2R}-1}{\rho_t\widehat{\beta}_{1,k}}\right)\left\{1\right.$$
$$\left.+\sum_{j=1}^{K-1}\sum_{L_j\subseteq S_k}\frac{(-1)^j}{\widehat{\beta}_{2,k}}\left(\frac{1}{\widehat{\beta}_{2,k}}+\sum_{q\in L_j}\frac{1}{\lambda_q}\right)^{-1}\right\}. \tag{26}$$

Finally, the asymptotic outage probability of EPRS for $\sigma_e^2 = 0$ can be obtained by substituting (23)–(26) into (7). In addition, the asymptotic outage probability of PRS for $\sigma_e^2 = 0$ can be obtained by inserting (23) and (24) into (7).

Table 1: Description of simulation cases.

Cases	R	K	Average channel powers
I	1 bps/Hz	2	$\beta_{1,1} = 1.1$, $\beta_{1,2} = 12$ $\beta_{2,1} = 11$, $\beta_{2,2} = 1.2$
II	1 bps/Hz	3	$\beta_{1,1} = 1.1$, $\beta_{1,2} = 12$, $\beta_{1,3} = 1.3$ $\beta_{2,1} = 11$, $\beta_{2,2} = 1.2$, $\beta_{2,3} = 13$
III	1 bps/Hz	3	$\beta_{1,1} = 1.1$, $\beta_{1,2} = 4.8$, $\beta_{1,3} = 1.3$ $\beta_{2,1} = 4.4$, $\beta_{2,2} = 1.2$, $\beta_{2,3} = 5.2$

From (23)–(26), it is observed that the order of $1/\rho_t$ in the asymptotic outage probability expression is one, and hence the diversity orders of EPRS and PRS are one.

In contrast to the asymptotic analysis for $\sigma_e^2 = 0$, the asymptotic outage probability of EPRS for $\sigma_e^2 \neq 0$ is easily obtained by inserting $R_t \approx (2^{2R} - 1)\sigma_e^2$ in (8)–(11), (17), and (18). Also, the asymptotic outage probability of PRS for $\sigma_e^2 \neq 0$ is simply obtained by replacing with $R_t \approx (2^{2R} - 1)\sigma_e^2$ in (8), (9), (17), and (18). Therefore, the asymptotic outage performance for EPRS and PRS is not affected by ρ_t, which results in the diversity order of zero.

4. Numerical Results

To verify the analysis presented in this paper and evaluate the outage performance, we consider three simulation cases as shown in Table 1, where it is noted that $\widehat{\beta}_{i,k}$'s in the outage probability expression are obtained by $\widehat{\beta}_{i,k} = \beta_{i,k} - \sigma_e^2$. In Cases I and II, $|10\log_{10}(\beta_{1,k}/\beta_{2,k})| = 10$ dB for all k, and $K = 2$ and 3, respectively, whereas in Case III, $|10\log_{10}(\beta_{1,k}/\beta_{2,k})| = 6$ dB for all k, and $K = 3$, where $|10\log_{10}(\beta_{1,k}/\beta_{2,k})|$ means a difference between the average channel powers for the first and the second hops at relay k. It is noted that Cases I and II are better scenarios than Case III, since EPRS and PRS provide better outage performance as the gap between the average channel powers for the first and the second hops becomes larger. Table 2 illustrates the values of correlation coefficient $\rho_d = J_0(2\pi f_d \tau)$ for various conditions. It is noted that the carrier frequency of 2.0 GHz has been used for 3GPP LTE system simulation as shown in [19]. From Table 2, we choose $\rho_d = 0.5, 0.8, 0.9$, and 1.0 for simulations, where $\rho_d = 1.0$ means that there is no feedback delay. It is noted that for $\rho_d = 1.0$ only simulated results are shown, since it is impossible to solve the outage probability expression when $\rho_d = 1.0$. Figures 2–7 show the outage probabilities of EPRS and PRS with various correlation coefficients (dependent on feedback delay time) and channel estimation errors for Cases I–III. All the figures demonstrate that the theoretical

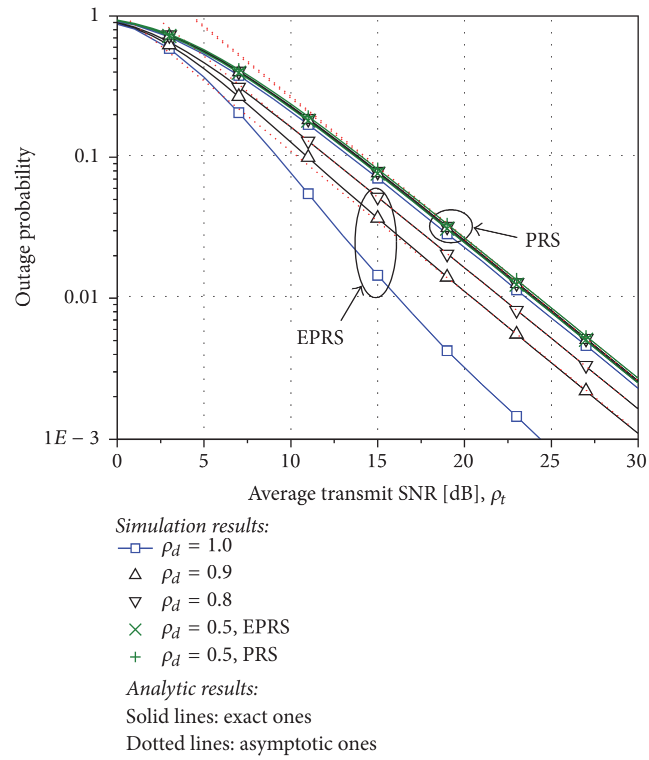

Figure 2: Outage probabilities of EPRS and PRS with perfect channel estimation for Case I when $\rho_d = 0.5, 0.8, 0.9, 1.0$.

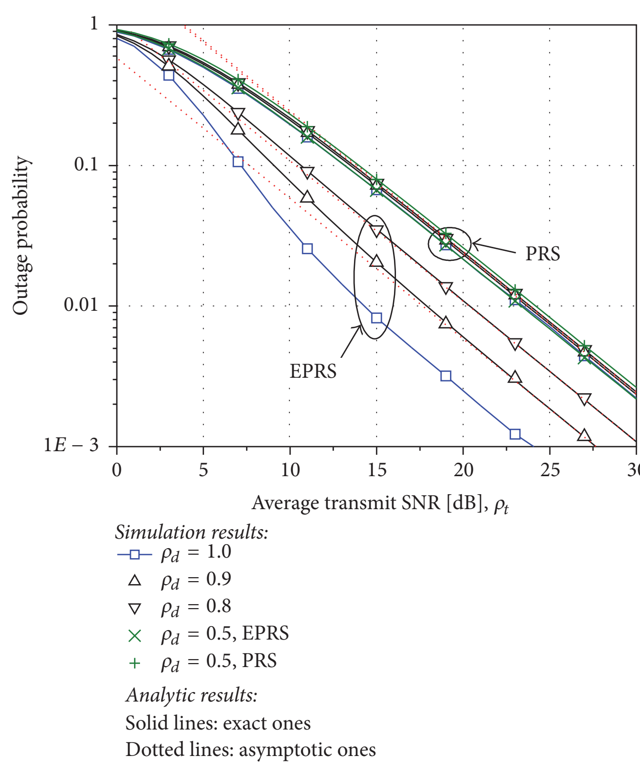

Figure 3: Outage probabilities of EPRS and PRS with perfect channel estimation for Case II when $\rho_d = 0.5, 0.8, 0.9, 1.0$.

Table 2: Correlation coefficient $\rho_d = J_0(2\pi f_d \tau)$, where f_c denotes the carrier frequency.

f_c [GHz]	τ [ms]	4.5 km/h	10 km/h	30 km/h	60 km/h
2.0	0.5	0.9998	0.9992	0.9924	0.9698
2.0	1	0.9993	0.9966	0.9698	0.8818
2.0	1.5	0.9985	0.9924	0.9326	0.7441
0.8	1.5	0.9998	0.9988	0.9891	0.9566
3.0	1.5	0.9965	0.9829	0.8516	0.4720

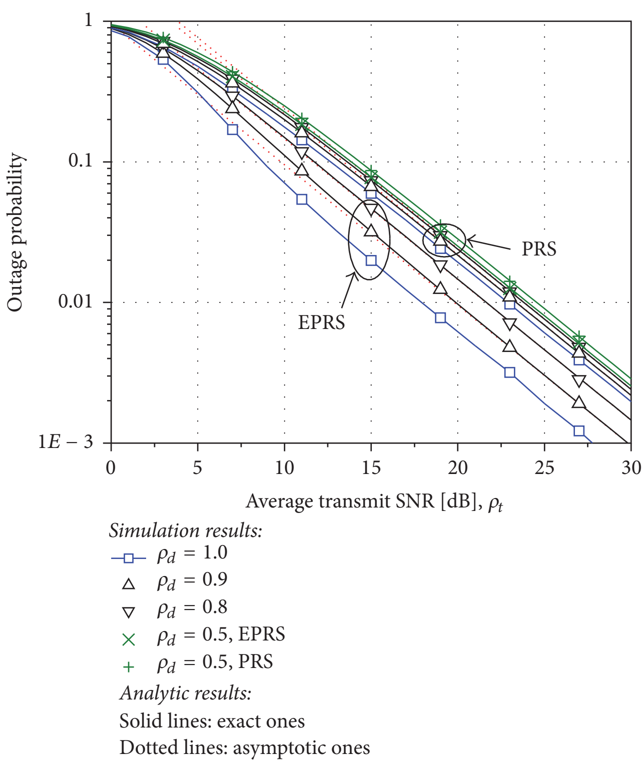

Figure 4: Outage probabilities of EPRS and PRS with perfect channel estimation for Case III when $\rho_d = 0.5, 0.8, 0.9, 1.0$.

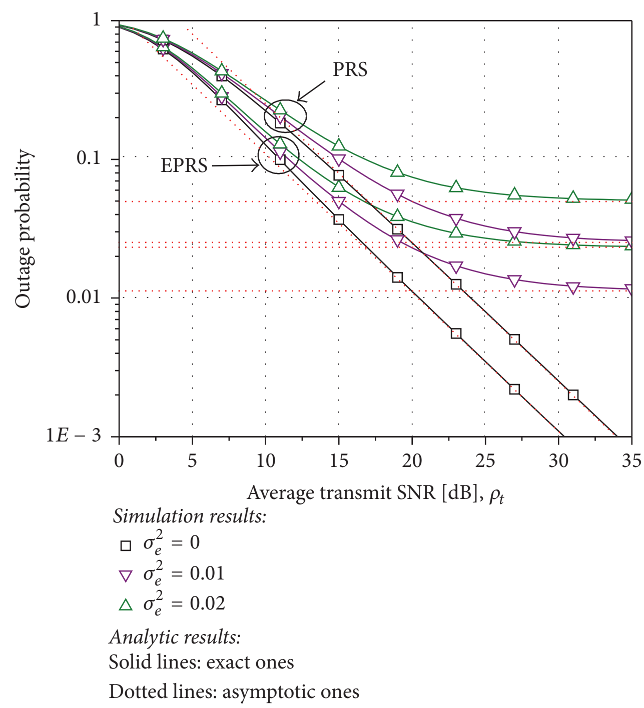

Figure 5: Outage probabilities of EPRS and PRS with $\rho_d = 0.9$ for Case I when $\sigma_e^2 = 0, 0.01, 0.02$.

analysis of exact outage probabilities is in perfect agreement with simulation results, and the asymptotic results are well matched with the simulated ones in the high SNR regime.

Figures 2–4 show the outage probabilities of EPRS and PRS with $\rho_d = 0.5, 0.8, 0.9$, and 1.0 for Cases I–III, respectively, when $\sigma_e^2 = 0$ (i.e., perfect channel estimation). From the figures, it is observed that the outage performance of EPRS is significantly degraded as ρ_d decreases (i.e., the feedback delay time increases), whereas the outage performance of PRS is nearly impervious to ρ_d. Also, it is indicated that the diversity orders of EPRS and PRS are not changed by ρ_d, since their diversity orders are one regardless of ρ_d. In comparing the outage performances for Cases I and II in Figures 2 and 3, respectively, EPRS achieves better outage performance than PRS as K increases, but the outage performance of EPRS becomes closer to that of PRS as both K and ρ_d diminish. The reason is that partial CSI used for EPRS can well represent the end-to-end channel quality for high ρ_d, but the accuracy of the partial CSI becomes worse as ρ_d decreases. In comparing the outage performances for Cases II and III in Figures 3 and 4, respectively, the outage performance of EPRS is degraded and is close to that of PRS as an average channel power gap between the first and the second hops decreases, since the accuracy of the partial CSI for EPRS is poor when an average channel power gap between the first and the second hops is small.

Figures 5–7 show the outage probabilities of EPRS and PRS with $\sigma_e^2 = 0$, 0.01, and 0.02 for Cases I–III, respectively, when $\rho_d = 0.9$. The figures indicate that the outage performances of both EPRS and PRS become worse and saturated as σ_e^2 increases, and their diversity orders are considerably reduced even for low σ_e^2. It is noted that, for $\sigma_e^2 \neq 0$, all the asymptotic results are constant with respect to ρ_t; that is, the diversity orders are zero. In addition, it is remarkable that when σ_e^2 increases from 0.01 to 0.02, a level of performance degradation of EPRS and PRS is similar for Cases I–III. In the outage performances for Cases I–III in Figures 5–7, respectively, analogous to the results in Figures 2–4, EPRS attains better performance than PRS as either K or an average

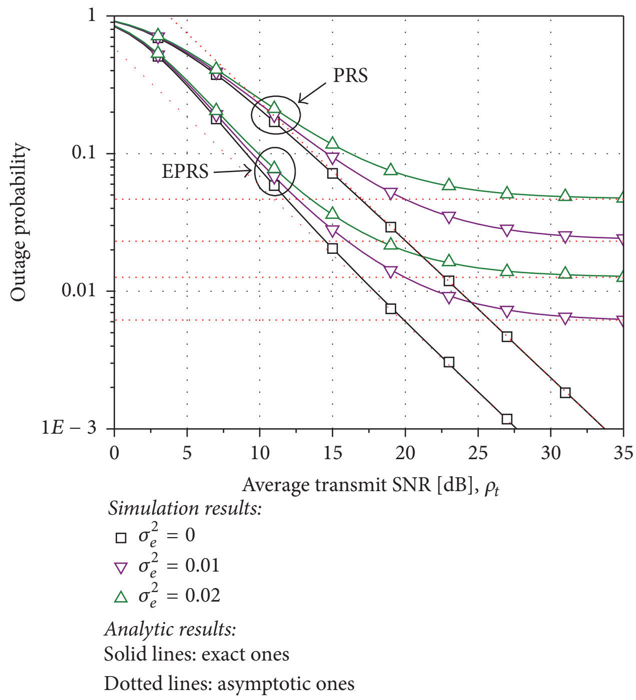

FIGURE 6: Outage probabilities of EPRS and PRS with $\rho_d = 0.9$ for Case II when $\sigma_e^2 = 0, 0.01, 0.02$.

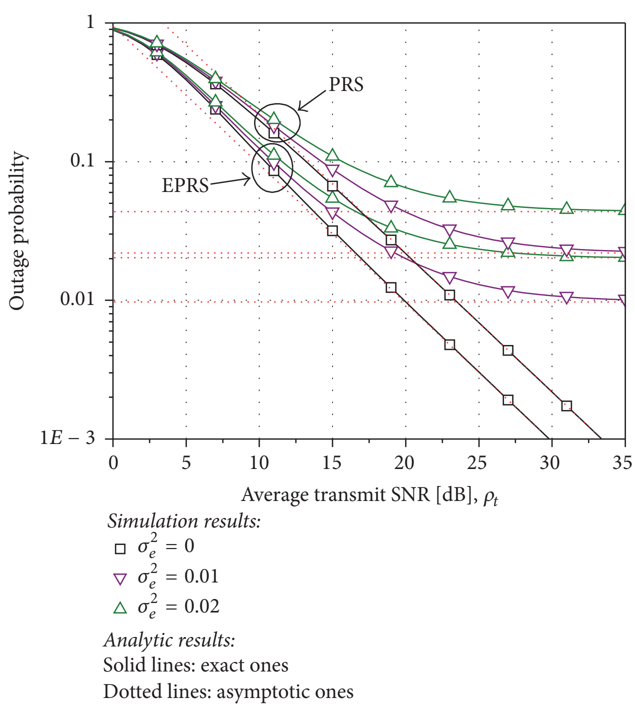

FIGURE 7: Outage probabilities of EPRS and PRS with $\rho_d = 0.9$ for Case III when $\sigma_e^2 = 0, 0.01, 0.02$.

channel power gap between the first and the second hops increases. Also, such a performance aspect is not affected by σ_e^2.

5. Conclusions

This paper presents the exact and closed-form expressions for outage probabilities of PRS and EPRS in dual-hop DF relaying systems with channel estimation errors and outdated CSI under nonidentical Rayleigh fading channels. In addition, the expressions for their asymptotic outage probabilities are presented. Numerical results verify the analytic expressions and show that the performance improvement of EPRS over PRS becomes smaller as the feedback delay time increases, but it can be better as the number of relays and the average channel power gap between the first and the second hops increase. Furthermore, the impact of channel estimation errors on the outage performance is much more serious than feedback delay, since the channel estimation errors induce a considerable reduction in the diversity order. Finally, we recognize that a CSI feedback design is much more important for EPRS than PRS, and an advanced channel estimation scheme is necessarily required for both EPRS and PRS in order to maintain the diversity order. Furthermore, when the multiantenna relays are considered, an impact of channel estimation errors and outdated CSI on the system performance may be more serious than the single-antenna scenario. Therefore, further study of the performance analysis of the cooperative relaying system with MIMO configuration is required in the presence of channel estimation errors and outdated CSI.

Competing Interests

The authors declare that they have no competing interests.

Acknowledgments

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