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Actuarial Pricing of Health Insurance for Breast Cancer Patients Using Markov Chain Method

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Abstract

Breast cancer treatment may impose substantial financial costs on patients and their families, motivating the assessment of stage-specific insurance risk. This study presents an illustrative calculation of net single premiums using a five-state Markov model that distinguishes breast cancer stages I–IV and death as an absorbing state. The analysis uses 346 paired patient observations from Abdul Moeloek Regional General Hospital, recorded in 2023 and 2024 at a one-year interval. Observed transitions yield an annual probability matrix, from which a continuous-time generator is constructed using a first-order approximation. For an assumed nominal benefit of IDR 100,000,000, the reported five-year net single premium is IDR 15,142,651.24. The corresponding amounts for six, seven, eight, nine, and ten years are approximately 4.69, 9.09, 14.06, 19.49, and 25.27 times the five-year value. These estimates illustrate sensitivity to the policy term within the stated numerical framework, rather than validated commercial premiums. The covered payment event, insured's initial cancer stage, benefit-payment frequency, and numerical discount rate are not specified in the available calculation; consequently, independent actuarial reproduction and contract-level interpretation remain necessary before practical pricing applications.

Keywords: Breast cancer, insurance, health, markov, multistate
MSC2020: 91G05, 60J10, 60J27

Abstrak

Pengobatan kanker payudara dapat menimbulkan beban finansial yang besar bagi pasien dan keluarganya, sehingga diperlukan kajian risiko asuransi berdasarkan stadium penyakit. Penelitian ini menyajikan ilustrasi perhitungan premi tunggal bersih dengan model rantai Markov lima keadaan yang terdiri atas kanker payudara stadium I–IV dan meninggal sebagai keadaan menyerap. Analisis menggunakan 346 pengamatan pasien berpasangan dari RSUD Abdul Moeloek pada tahun 2023 dan 2024, dengan interval pengamatan satu tahun. Data perpindahan stadium digunakan untuk menyusun matriks peluang transisi tahunan, kemudian matriks generator waktu kontinu didekati menggunakan pendekatan orde pertama. Dengan asumsi nilai manfaat nominal Rp100.000.000, premi tunggal bersih lima tahun yang dilaporkan adalah Rp15.142.651,24. Nilai premi untuk jangka waktu enam, tujuh, delapan, sembilan, dan sepuluh tahun berturut-turut sekitar 4,69; 9,09; 14,06; 19,49; dan 25,27 kali nilai premi lima tahun. Hasil tersebut menunjukkan pola perubahan keluaran numerik seiring bertambahnya masa pertanggungan, tetapi belum dapat dianggap sebagai tarif premi produk asuransi yang telah tervalidasi. Kejadian pemicu pembayaran manfaat, stadium awal tertanggung, aturan frekuensi pembayaran, dan tingkat diskonto numerik belum dinyatakan secara lengkap; oleh karena itu, reproduksi aktuaria secara independen dan validasi ketentuan kontrak masih diperlukan sebelum penerapan praktis.

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Kata kunci: Kanker payudara, asuransi, kesehatan, markov, multistate
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Introduction

Risk of illness can result in sustained treatment expenses and threaten household financial stability. Insurance transfers specified financial risks from an insured individual to an insurer in exchange for a premium [1], [2]. The expected present value of an insurance benefit depends on the covered event, benefit amount, policy term, and discount assumptions [3]. For cancer coverage, defining whether a payment is triggered by diagnosis, stage progression, treatment, or death is essential to interpreting the calculated premium [4], [5].

Breast cancer is characterized by heterogeneous disease stages and clinical outcomes [6], [7]. A stage-specific survival study conducted at Abdul Moeloek Regional General Hospital illustrates the relevance of stage differentiation [8]. Recent global cancer statistics further emphasize the public-health significance of breast cancer [9]. These observations motivate the use of a multistate model rather than a single homogeneous diagnosed state.

A multistate stochastic framework represents movements between defined clinical states over time and has been used in insurance applications, including heart-disease coverage [10], [11]. In this study, the states are breast cancer stages I, II, III, IV, and death; a healthy state is not included. Under the Markov assumption, the next-state distribution is conditional on the current modeled state [12], and death is absorbing. Annual paired observations identify one-year transition proportions but do not reveal the exact times at which within-year transitions occurred.

Earlier Markov-chain insurance applications have considered heart disease [10], dengue [13], and COVID-19 [14]. More directly, Sari et al. [4] developed breast cancer insurance pricing using a Cox–Markov multistate approach, while Arık et al. [5] examined model risk and semi-Markov alternatives in cancer insurance. Thus, Markov modeling and breast cancer insurance pricing are not new. The narrower contribution of this study is an empirical, stage-specific illustration based on 346 paired annual observations from one Lampung hospital, with numerical comparisons for policy terms of five to ten years. Unlike covariate-based or semi-Markov analyses, this study uses an aggregate five-state transition structure. The objective is to describe its model-based outputs and limitations; external predictive validity and a fully specified insurance product are not established.

Methods

Study design and observations. The transition analysis uses 346 records with paired observations of breast cancer stage in 2023 and stage or death status in 2024. The interval is one year per patient. An analyzable record requires a documented stage I–IV at baseline and a documented stage I–IV or death at follow-up. Recruitment details, eligibility screening, excluded records, handling of missing observations, and ethical-approval information were not provided in the source material; these must be verified by the authors. Consequently, the 346 records constitute the analyzed sample and cannot be assumed to form a random or population-representative sample. Periodically observed multistate data also require care when estimating continuous-time transition intensities [15].

Breast Cancer

Breast cancer progression can be classified into clinically defined stages, with outcomes varying across patient groups [6], [7]. Local survival evidence from Abdul Moeloek Regional General Hospital supports the use of stage-specific categories in the present illustration [8]. The following actuarial model uses observed stage transitions; it does not estimate population-wide cancer incidence, absolute survival probabilities, or individual treatment expenditures.

Matrix diagonalization

Matrix diagonalization is a crucial technique in linear algebra used to simplify various matrix operations, such as high-power calculations or the analysis of a system's dynamic properties. A matrix is said to be diagonalizable if it is similar to a diagonal matrix. In the diagonalization process, an invertible matrix M is used to transform the original matrix A into a diagonal matrix D. Consequently, the form of matrix A can be expressed as shown in the following Equation (1) [16].

$$A = MDM^{-1} \tag{1}$$

Markov Chain

A Markov chain is a stochastic process that describes the transition of a system from one state to another with specific probabilities. In a Markov chain, the probability of a future event depends only on the current state and is not influenced by previous states; this is known as the memoryless property. This property can be expressed in the form of transition probabilities as follows[12]:

$$p_{ij} = P\{X(t+s) = j | X(t) = i\} \tag{2}$$

In this study, a Markov chain approach is used to model the health status transition process of breast cancer patients over time. The Markov chain allows us to represent the potential movement of patients between different health conditions. These transitions can be seen in Figure 1 [12].

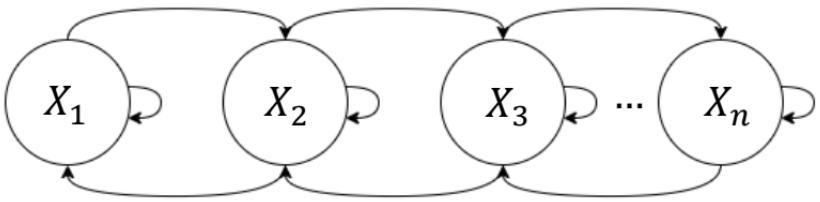


Figure 1. Multistate model

The Chapman–Kolmogorov equation is a fundamental property in Markov chains that expresses the transition probability from one state to another within a certain time interval through an intermediate state. The general form of the Chapman–Kolmogorov equation can be written as follows [17]:

$$p_{ij}(t+s) = \sum_{k=0}^{m} p_{ik}(t) p_{kj}(s) \tag{3}$$

The Chapman–Kolmogorov relation leads to the Kolmogorov forward equation, which describes how transition probabilities evolve under a time-homogeneous generator. The general expression is given in Equation (4) [17].

$$\frac{d}{dt}p_{ij}(t) \quad = \sum_{k=1}^{m} p_{ik}(t)\mu_{kj} \tag{4}$$

Transition Rates

In continuous-time Markov chains, the rate of transition between states is expressed through the transition rate μ_{ij}, which is organized within a generator matrix Q. Matrix Q plays a vital role in modeling the dynamics of transitions between states and serves as the foundation for transition probability analysis. The relationship between the elements of the generator matrix is stated as follows in Equation (5)[18]:

$$\sum_{j\neq i}^{n} \mu_{ij} \quad = \quad -\mu_{ii} \tag{5}$$

For a time-homogeneous continuous-time Markov chain, the generator Q determines the transition matrix through $P(t) = exp(Qt)$, as presented in Equation (6) [18], [17]. This equality requires the generator to be appropriately estimated from continuous-time information or consistently embedded from the annual transition matrix; it is not established merely by treating annual transition probabilities as instantaneous rates.

$$\begin{aligned} P(t) &= e^{Qt} \\ &= \sum_{n}^{\infty} \frac{t^n}{n!} Q^n = M e^{Dt}\, M^{-1}. \end{aligned} \tag{6}$$

The use of matrix diagonalization allows for the simplification of the matrix exponential calculation of $\boldsymbol{Q}$, as shown in Equation (6). Matrix $\boldsymbol{Q}$ is decomposed into MDM^{-1}, such that the calculation of $e^{\boldsymbol{Q}t}$ can be performed simply on the diagonalized matrix, which has a simpler structure. When expanded in matrix form, matrix M consists of elements a_{ik}. Matrix $e^{\boldsymbol{D}t}$ contains elements e^{dkt}, which are the exponentials of the elements d_k, and matrix M^{-1} consists of elements c_{kj}. Based on this description and Equation (7), the result can be obtained as shown in Equation (7)[16].

$$p_{ij}(t) = \sum_{k=1}^{m} a_{ik} e^{d_k t} c_{kj} \tag{7}$$

Net Single Premium

A net single premium is the expected present value of the benefit payments specified by an insurance contract, before expenses and other loadings [3]. In the numerical example in this manuscript, the benefit multiplier is IDR 100,000,000. Equation (8) is formulated for a transition-contingent benefit and includes discounting at a constant force of interest, denoted by δ. However, the source calculations do not identify the specific insured transition(s), whether more than one benefit payment is permitted, or the numerical value of δ. Therefore, the displayed amounts should be treated as the authors' numerical illustration, rather than verified premiums for a fully specified insurance product. The premium expression used in the analysis is written as follows:

$$\bar{A}^{1}_{x:\overline{n}|} = b_t \sum_{k=0}^{\infty} \frac{1}{(\delta - d_k)} (1 - e^{(d_k - \delta)n}) a_{ik} c_{kj} \mu_{ij} \quad (8)$$

Results and Discussion

The transition calculations are based on 346 paired breast cancer records from Abdul Moeloek Regional General Hospital, Lampung, recording an initial stage in 2023 and the corresponding subsequent stage or death in 2024. The observed sample represents one annual transition per patient. Records are grouped by their initial stage to estimate conditional one-year transition proportions, using the number moving from state i to state j divided by the number starting in state i. Table 1 provides an excerpt of the anonymized observations, with the first and last records shown; the complete 346-row patient-level data are not reproduced. The annual transition matrix is then used to construct an approximate continuous-time generator for the premium illustration.

Table 1. Excerpt from 346 paired breast cancer patient observations

No	Age	Gender	Entry Year	Initial Stage	Following Year	Next Stage	Final Status
1.	59	W	2023	III	2024	IV	Alive
2.	41	W	2023	II	2024	IV	Alive
3.	43	W	2023	II	2024	III	Alive
4.	53	W	2023	II	2024	III	Alive
5.	64	W	2023	II	2024	III	Alive
⋮	⋮	⋮	⋮	⋮	⋮	⋮	⋮
342	40	W	2023	II	2024	III	Alive
343	59	W	2023	III	2024	IV	Alive
344	55	W	2023	II	2024	III	Alive
345	61	W	2023	IV	2024	Death	Death
346	43	W	2023	III	2024	IV	Alive

The disease process is represented by five states: stage I (1), stage II (2), stage III (3), stage IV (4), and death (0). The allowable movements follow the observed and represented transitions, with death as an absorbing state. Figure 2 shows this breast-cancer-specific multistate structure; no healthy state is included.

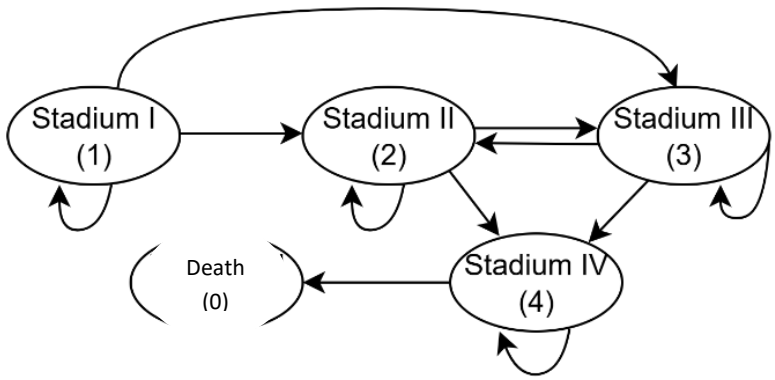


Figure 2. Multistate model for breast cancer patients

One-year conditional transition probabilities were calculated from the paired observations using the count of patients moving from initial state i to follow-up state j divided by the total count starting in state i. The initial-state denominators derived from the displayed calculations are 50 in stage I, 75 in stage II, 76 in stage III, and 145 in stage IV, for a total of 346 observations. The example calculations below illustrate this construction.

$$p_{11} = \frac{5}{50} = 0{,}1 \qquad p_{32} = \frac{3}{76} = 0{,}0395$$

$$p_{12} = \frac{25}{50} = 0{,}5 \qquad p_{33} = \frac{30}{76} = 0{,}3947$$

$$p_{13} = \frac{20}{50} = 0{,}4 \qquad p_{34} = \frac{43}{76} = 0{,}5658$$

$$p_{22} = \frac{10}{75} = 0{,}1333 \qquad p_{44} = \frac{135}{145} = 0{,}931$$

$$p_{23} = \frac{50}{75} = 0{,}6667 \qquad p_{40} = \frac{10}{145} = 0{,}069$$

$$p_{24} = \frac{15}{75} = 0{,}2$$

The value p_{11} represents the observed one-year proportion of patients in stage I (state 1) who remained in stage I at follow-up. The resulting empirical transition matrix represents probabilities over the 2023–2024 observation interval, not instantaneous rates, as displayed below:

$$P = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 \\ 0 & 0{,}1 & 0{,}5 & 0{,}4 & 0 \\ 0 & 0 & 0{,}1333 & 0{,}6667 & 0{,}2 \\ 0 & 0 & 0{,}0395 & 0{,}3947 & 0{,}5658 \\ 0{,}0690 & 0 & 0 & 0 & 0{,}9310 \end{bmatrix}$$

To use a continuous-time representation, the manuscript subsequently forms a generator matrix Q from the empirical one-year transition matrix P. The matrix shown below has the structure $Q = P - I$ when the time unit is one year, equivalently $Q \approx (P - I)/\Delta t$ for $\Delta t = 1$ year. Thus, its off-diagonal entries are treated as approximate annualized transition intensities, while its diagonal entries are set so that each row sums to zero. The resulting Q is an approximation, not a generator inferred from exact individual transition times:

$$Q = \begin{bmatrix} 0 & 0 & 0 & 0 & 0 \\ 0 & -0{,}9 & 0{,}5 & 0{,}4 & 0 \\ 0 & 0 & -0{,}8667 & 0{,}6667 & 0{,}2 \\ 0 & 0 & 0{,}0395 & -0{,}6053 & 0{,}5658 \\ 0{,}0690 & 0 & 0 & 0 & -0{,}0690 \end{bmatrix}$$

In this approximate matrix Q, off-diagonal entries are nonnegative, diagonal entries represent negative total exit intensities, and each row sums to zero [18], [19]. The death row is zero because death is absorbing. The relationship $P(t) = exp(Qt)$ is then applied to this approximation. Importantly, $exp(Q \times 1\ year)$ need not reproduce the empirical annual matrix P exactly; equating

$P - I$ with the exact generator is a first-order approximation that can be inaccurate when annual transition probabilities are substantial. Estimation by a valid matrix logarithm or by a continuous-time transition likelihood would require additional methodological verification and, where available, more detailed observation times.

$$M = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 \\ 1 & 1{,}2767 & 3{,}7144 & 1 & 87{,}7212 \\ 1 & 1{,}2068 & 1{,}9663 & 0 & -8{,}5804 \\ 1 & 1{,}1439 & 1 & 0 & 1 \\ 1 & 1 & 0 & 0 & 0 \end{bmatrix}$$

Matrix D contains the eigenvalues of matrix Q. This matrix is diagonal, meaning that it only has elements on its main diagonal, while the off-diagonal elements are zero. The values on the main diagonal are the eigenvalues calculated from matrix Q. The form of matrix D can be written as follows:

$$D = \begin{bmatrix} 0 & 0 & 0 & 0 & 0 \\ 0 & -0{,}069 & 0 & 0 & 0 \\ 0 & 0 & -0{,}5276 & 0 & 0 \\ 0 & 0 & 01 & -0{,}9 & 1 \\ 0 & 0 & 0 & 0 & -0{,}9444 \end{bmatrix}$$

Thus, having obtained matrix D and matrix M from the calculation of eigenvalues and eigenvectors, the next step is to determine matrix M^{-1}. This inverse matrix is required to construct the continuous-time transition probability matrix, which is calculated alongside matrices D and M. Matrix M^{-1} is written as follows:

$$M^{-1} = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 \\ -1 & 0 & 0 & 0 & 1 \\ 0{,}1367 & 0 & 0{,}0948 & 0{,}8136 & -1{,}0451 \\ -0{,}8642 & 1 & 7{,}9625 & -19{,}3711 & 11{,}2727 \\ 0{,}0072 & 0 & -0{,}0948 & 0{,}1864 & -0{,}0988 \end{bmatrix}$$

After obtaining the elements of matrices M, D, and M^{-1}, the continuous-time transition probability matrix can be calculated. The variable a_{ik} represents the elements of matrix M, the variable e^{dk} is the exponential of the elements of matrix D, and the variable c_{kj} represents the elements of matrix M^{-1}.

The numerical illustration evaluates the net single premium for coverage terms of 5, 6, 7, 8, 9, and 10 years using the five-state transition model and an assumed nominal benefit of IDR 100,000,000. Since the model begins with a diagnosed cancer patient, an initial policyholder state must be one of stages I–IV, not a healthy state. The original calculation is presented with generic origin and destination indices (i, j) and does not specify which initial stage or combination of stages underlies the six reported premium amounts. This missing assumption must be fixed before interpreting the values as stage-specific prices. The five-year calculation reported by the manuscript is shown below.

$$\begin{aligned} \bar{A}^{1}_{ij:\overline{5}|} &= b_t(-0{,}9340 + 0{,}2876 + 0{,}4158 + (-0{,}9626) + 0{,}6706 + \\ &\quad 0{,}4003 + 0{,}0024 + (-0{,}9109) + 1{,}3946 + 0{,}0444 + (-0{,}2567 \\ &= Rp.100.000.000{,}00 \times 0{,}1514265124 \\ &= Rp.15.142.651{,}24 \end{aligned}$$

Overall, the net single premium values for breast cancer patients based on variations in the coverage period can be seen in **Table 2**.

Table 2. Net single premiums and ratios by coverage term

Term (years)	Net single premium (IDR)	Ratio to prior term	Ratio to 5-year term
5	15,142,651.24	—	1.00
6	70,946,736.21	4.69	4.69
7	137,674,957.53	1.94	9.09
8	212,979,913.54	1.55	14.06
9	295,093,215.34	1.39	19.49
10	382,626,587.09	1.30	25.27

Table 2 distinguishes two different comparisons. The year-on-year ratio is the premium for year n divided by the premium for year $n - 1$, whereas the ratio to the five-year baseline is the premium for year n divided by the five-year premium of IDR 15,142,651.24. The six-year premium is IDR 70,946,736.21, corresponding to 4.69 times the five-year amount. The year-on-year ratios thereafter decline to 1.94, 1.55, 1.39, and 1.30 for seven through ten years; these figures are not the ratios against the five-year base. All ratios are descriptive of the numerical outputs in Table 2 and are subject to the contract and calibration assumptions discussed above.

In the specified state space, longer coverage terms provide more opportunities for transitions among stages I–IV and to the absorbing death state. However, the estimates in Table 2 require further actuarial verification. The reported seven- to ten-year net single premiums exceed the assumed nominal benefit of IDR 100,000,000; this would not be consistent with a nonnegative-discount, single-payment contract whose total benefit is capped at that amount. Multiple benefit payments or a different contract structure could alter this bound, but no such specification is available in the source calculations. Accordingly, Table 2 should be read only as a report of the existing numerical outputs, not as validated insurance tariffs. The underlying benefit trigger, payment rule, initial state, discount assumption, and arithmetic must be confirmed by the authors before publication. Figure 3 visualizes these reported amounts.

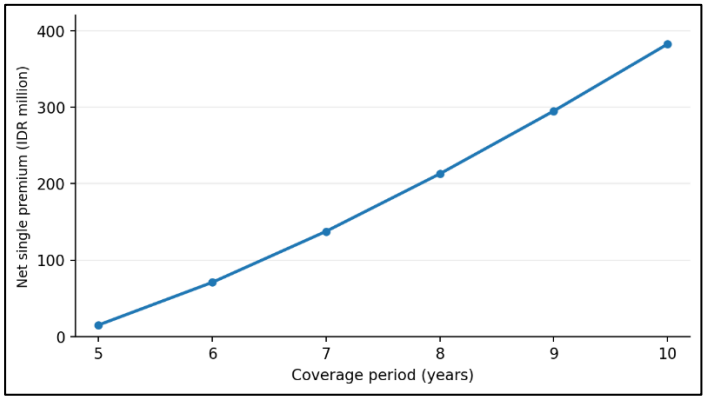


Figure 3. Reported net single premiums for 5–10-year coverage periods

Figure 3 displays the net single premiums corresponding to coverage periods of 5–10 years. The reported five-year premium is IDR 15,142,651.24. When each term is compared with this same five-year baseline, the six-, seven-, eight-, nine-, and ten-year amounts are approximately 4.69, 9.09, 14.06, 19.49, and 25.27 times as large, respectively. By contrast, the year-on-year ratios in Table 2 compare adjacent coverage terms. The increasing pattern is specific to the reported numerical model and is not independently validated as an insurer's payable premium.

Discussion of empirical evidence and model validation

The observed 2023–2024 transitions measure the proportions of patients occupying particular stages at the beginning and end of a one-year interval. They provide an informative description of this hospital sample, but they do not establish that the same transition pattern would hold in each subsequent year. Intermediate clinical changes are not identified by two annual observations, and the aggregate state table does not separately describe differences in treatment, duration since diagnosis, or other patient characteristics. Accordingly, the empirical matrix characterizes the available paired records rather than a verified long-term transition process. This qualification becomes increasingly important as the coverage term extends from five to ten years: repeated projection may accumulate not only disease-state changes but also uncertainty surrounding the assumption that the transition law remains stable.

The conversion to continuous time introduces a further modeling distinction. The reported off-diagonal entries of the generator matrix Q are based on an approximation to annual transition probabilities, while the diagonal entries are selected so that each row sums to zero. Although this has the algebraic form required of a generator, it does not guarantee that its one-year matrix exponential reproduces the observed transition matrix. An informative model check would compare the fitted and empirical probabilities for each initial stage, including probabilities of remaining at that stage and moving to the absorbing death state. Material differences would indicate that the numerical premiums may depend on the selected approximation. Estimation using a valid matrix logarithm or a continuous-time transition likelihood could be investigated as an alternative, but neither procedure has been verified as an additional result of the present analysis [15], [18].

The premium differences across the terms in Table 2 should likewise be distinguished from a measured change in an individual patient's clinical risk. An actuarial value depends on a contract's benefit trigger, the stage occupied when coverage begins, the timing of payment, and whether one or multiple payments are allowed. These provisions determine the cash flows whose discounted expected value is calculated [3], [20]. In particular, for a contract paying no more than IDR 100,000,000 once, with nonnegative discounting, the net single premium cannot exceed that benefit. The reported amounts above this threshold therefore call for a review of the contract specification or arithmetic, rather than an inference that long-term insurance is necessarily more expensive than its stated maximum benefit. No additional premiums are recalculated in this discussion.

A reproducible next stage of the study would therefore document the origin-to-destination transition counts, state ordering, method and time unit used to construct Q, initial-stage assumption, covered transition or event, payment frequency, and numerical discount rate. Mechanical checks would include the row sums of the annual transition matrix, the sign and row-sum restrictions on the generator, and the treatment of death as an absorbing state. A comparison between the empirical one-year probabilities and those implied by the continuous-time model would assess calibration. Finally, varying relevant assumptions in a sensitivity analysis could identify which inputs most strongly affect the calculated term premiums. These are recommendations for validation, not analyses claimed to have been conducted here. They provide a route toward a more transparent application while leaving the five-state model, reported data, and original research objective unchanged.

Conclusion

A five-state Markov model comprising breast cancer stages I–IV and death was applied to 346 paired annual observations from 2023 and 2024. For a nominal benefit of IDR 100,000,000, the existing numerical calculation reports a five-year net single premium of IDR 15,142,651.24. The calculated six-

to ten-year amounts are approximately 4.69, 9.09, 14.06, 19.49, and 25.27 times the five-year amount. These values are preliminary illustrations, not validated contract premiums. In particular, the transition generator was obtained through a first-order approximation, and the benefit event, initial cancer stage, payment multiplicity, and discount-rate input require verification. The fact that several reported premiums exceed the nominal benefit makes contract specification and numerical recalculation essential. Future work should evaluate a fully specified insurance contract and compare annual-panel and continuous-time estimation approaches.

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