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Offspring Sex Ratio and Birth-Record-Based Reproductive Indicators of Sambar Deer (*Rusa unicolor brookei*) in an Ex-Situ Breeding Program in East Kalimantan, Indonesia

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Abstract. *Offspring sex ratio and the pace of rebreeding can be read directly from routine birth records, making them a practical first check on an ex-situ breeding program. We analysed 104 official birth-inspection records (January 2019 to April 2024) of the Vulnerable Kalimantan sambar deer, Rusa unicolor brookei, at UPTD PTHPT Api-Api, East Kalimantan. Sex ratio was tested against 1:1, among years, for a linear trend, and by recorded parity, sire and birth season, using effect sizes, 95% confidence intervals (CI), Holm-adjusted p-values and sensitivity analyses for repeated births of the same dams. The pooled ratio was 54 males to 50 females (proportion male 0.52, 95% CI 0.42–0.61; p = 0.695). Composition differed modestly among years (χ^2 = 9.73, p = 0.045; Cramér's V = 0.31), but this weakened (p = 0.055–0.075) once within-dam correlation was allowed for, and no single year departed from parity after Holm adjustment. Neither the male-ward trend (odds ratio 1.27 per year, p = 0.096) nor the parity, sire and season contrasts were significant, although their wide intervals still allow moderate effects. Eighteen of 68 identified dams (26.5%) had at least two recorded births; because late-registered dams had little time to rebreed, this is a lower bound within the observation window, not a measure of reproductive success. The mean inter-birth interval was 515 days (95% CI 462–568; median 441), longer than the 388 days previously reported for this herd. These indicators are preliminary and cannot alone establish population viability.*

Keywords: *Birth records, Ex-situ conservation, Inter-birth interval,* Rusa unicolor brookei*, Sex ratio*

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INTRODUCTION

A growing head count does not by itself secure the future of a threatened species. An ex-situ breeding program succeeds when that number becomes a viable breeding stock, with a workable sex ratio, an adequate reproductive rate, and enough demographic depth to persist without repeated restocking (Soulé et al., 1986; Lacy, 2013; IUCN/SSC, 2014; McGowan et al., 2017). Two of these criteria, sex ratio and reproductive rate, can be taken straight from routine birth records. They require neither genetic sampling nor a formal population viability analysis, so any facility that keeps basic breeding records can use them as a practical first check on program health. Their value lies in catching problems early. They cannot replace a viability analysis, which also depends on survival, age-specific fecundity and genetic data. Recent reviews of cooperatively managed zoo and conservation-breeding populations caution that reproductive performance can quietly underperform, even when resources are adequate, unless management stays tied to current, program-specific evidence rather than assumptions (Putnam et al., 2023).

The sambar deer (*Rusa unicolor*) is the largest deer species native to South and Southeast Asia and has been listed as Vulnerable on the IUCN Red List since 2008, a status reaffirmed in 2015 (Timmins et al., 2015). Its Kalimantan subspecies, *Rusa unicolor brookei*, is protected under Indonesian Law Number 5 of 1990 concerning the Conservation of Living Natural Resources and Their Ecosystems (Republic of Indonesia, 1990). UPTD Pembibitan Ternak dan Hijauan Pakan Ternak (PTHPT) Api-Api, operated by the East Kalimantan Provincial Livestock and Animal Health Service, maintains one of the province's captive breeding herds of this subspecies and issues an official birth-inspection report, recording birth date, sex, and parent identity, for every fawn born on site. Herd composition is routinely described in three age-sex classes: adult males (stags), adult females (hinds), and young (fawns), the last of which is the class directly recorded in this study's source data (Figure 1).

Offspring sex ratio in mammals is rarely a simple coin flip. The Trivers-Willard hypothesis predicts that mothers in better condition should bias investment toward the sex with the greater potential reproductive payoff, typically sons in polygynous species (Trivers & Willard, 1973), and maternal effects of this kind on sex allocation have since been documented across a range of ungulates (Hewison & Gaillard, 1999; Sheldon & West, 2004). More recent work on other deer species has kept testing this prediction with

mixed results: a multi-decade dataset on roe deer found a persistent male bias unrelated to Trivers-Willard-type maternal effects (Hagen et al., 2022), while a comparable analysis of white-tailed deer found the sex ratio tracking Fisherian frequency dependence rather than maternal condition (Dentinger et al., 2024). A parallel line of work has turned toward the sire's side of the equation, treating paternal condition and identity as sex-allocation candidates in their own right rather than as a residual to be ruled out (Pawluk & Zagalska-Neubauer, 2024). Our group has used the same Api-Api birth registry in two companion manuscripts (unpublished): one describes population dynamics and the seasonal distribution of births, and the other evaluates sire utilisation and the risk of a narrowing genetic base. The first reported that sex ratio composition differed among years but did not test what might explain the difference.

The present study draws on the same 104 records but asks different questions. Annual birth counts and sex ratios (Table 1) are repeated only as the baseline for new tests; the seasonal distribution of births, the contribution of each sire to the next generation and pedigree-based genetic indicators are not re-analysed here. The study adds three things. First, it tests a directional trend and three candidate correlates of offspring sex (recorded parity, sire identity and birth season) with effect sizes, confidence intervals, multiplicity control and sensitivity analyses for repeated births of the same dams. Second, it derives dam-level reproductive indicators, namely the proportion of dams with two or more recorded births and the inter-birth interval, and interprets them against the follow-up each dam actually had. Third, it compares these indicators with values reported for the same herd two decades earlier (Semiadi et al., 2005). A birth register sets the limits of this study. Its records cannot provide estimates of conception rate, fawn survival, age at first fawning or population growth rate. For that reason, the analysis covers a selected set of reproductive indicators and does not assess overall reproductive performance.

LITERATURE REVIEW

The departures from 1:1 that sex allocation theory predicts apply to individual mothers and may not show up at the level of a herd. The Trivers-Willard hypothesis expects mothers in good condition to produce more sons in polygynous species, because male reproductive success there is more sensitive to condition (Trivers & Willard, 1973). Evidence from ungulates is only partly supportive. Maternal rank, age and condition have each been associated with offspring sex, yet the effects are small and vary between studies

and between years (Hewison & Gaillard, 1999; Sheldon & West, 2004). Two large deer datasets illustrate why the issue remains unresolved. More than 12,000 roe deer births showed a weather-linked male bias without a maternal-condition signature (Hagen et al., 2022), whereas sex ratio variation in white-tailed deer fitted Fisherian frequency dependence (Dentinger et al., 2024). For a herd with about 100 recorded births, two consequences follow. Maternal effects, if present, are probably small, so only large effects are detectable. The variable that theory points to, maternal condition, is also absent from the register, which leaves recorded parity as a coarse proxy for maternal age and breeding experience (H2 below).



Figure 1. Age-sex classes of sambar deer (*Rusa unicolor*) referenced in this study: adult male (stag), adult female (hind), and young (fawn)

Studies of the father's contribution are fewer. In a captive primate, dominant males with higher testosterone sired litters with more sons, regardless of the female's condition (Perret, 2018). A 29-year record from a captive numbat program likewise found that paternal origin influenced offspring sex ratio after parental weight and maternal history were taken into account (Firman et al., 2023). Reviews now treat sire identity and condition as candidate mechanisms in their own right (Pawluk & Zagalska-Neubauer, 2024). In a herd where a few stags sire most fawns, a sire-specific bias could shift the annual ratio, which motivates H3. Birth timing is the third candidate. When several sex-allocation hypotheses were tested together in kangaroos, no single mechanism explained offspring sex (Toni et al., 2021). Sambar breeding is seasonal even near the equator: births

in the East Kalimantan herd peaked in June and July two decades ago (Semiadi et al., 2005), and births become more concentrated at higher latitude, where about 80% of wild hinds in Victoria, Australia, calve between April and August (Watter et al., 2020). Hinds that conceive outside the main season may differ in age or condition from those that conceive within it, so birth season is a plausible correlate of offspring sex (H4).

Conservation guidelines judge captive programs by whether they can sustain themselves. Soulé et al. (1986) proposed retaining 90% of founding genetic diversity over 100 to 200 years; Lacy (2013) defined sustainability as maintaining numbers through the population's own reproduction without new founders; and the IUCN/SSC (2014) guidelines, with the decision framework of McGowan et al. (2017), treat sex ratio and reproductive output as indicators to be reviewed regularly. Cooperatively managed zoo populations still fall short of self-sustaining reproduction despite science-based pairing recommendations (Putnam et al., 2023). Meeting these criteria requires demographic rates (survival, age-specific fecundity, growth rate) and genetic parameters that a birth register cannot supply. We therefore use sex ratio and inter-birth interval as early-warning indicators within this framework, not as tests of sustainability.

Species-specific values give these criteria a concrete reference. Semiadi et al. (1994) reported a mean gestation of 260 days and a mean calving interval of 329 days in a farmed herd in New Zealand. Dahlan and Dawend (2013) reported a gestation of 259 days and first fawning at about 32 months in Sarawak, Malaysia. For the captive herd in East Kalimantan, Semiadi et al. (2005) reported a mean calving interval of 388 days (SD 82; n = 33), a conception rate of 48.8% and births concentrated in June and July. Since then, sambar research has focused on males and genetics, including faecal testosterone and sexual behaviour (Gholib et al., 2025), testicular and sperm traits (de Matos Brandão Carneiro et al., 2025), mating behaviour in a semi-wild herd (Murad et al., 2023) and genetic diversity of captive Formosan sambar (Liang et al., 2023). Multi-year data on female output and offspring sex ratio remain scarce.

Benchmarks from other herds must be used with care. Herds differ in subspecies, latitude and photoperiod, feeding level, stocking density and mating system, and studies calculate intervals in different ways (all intervals, completed intervals only, or one mean per dam); most report a mean without a measure of spread. A gap between two herds therefore marks something to investigate, not a measured deficit. The same-site value of

Semiadi et al. (2005) removes most of these differences, although management at the facility may have changed over twenty years.

From this literature we derived four hypotheses, each tested against a null of no association: (H1) the proportion of male births follows a linear trend over 2019–2023; (H2) the proportion of male births differs between a dam's first recorded birth and her subsequent recorded births; (H3) individual sires differ in the proportion of male offspring; and (H4) the proportion of male births differs between births in October–March and April–September.

RESEARCH METHODS

Study Site and Population

The study was conducted at UPTD Pembibitan Ternak dan Hijauan Pakan Ternak (PTHPT) Api-Api, Jalan Negara Km. 32, Desa Api-Api, Kecamatan Waru, Kabupaten Penajam Paser Utara, East Kalimantan. The facility maintains a captive herd of the *Rusa unicolor brookei* subspecies in fenced paddocks under semi-natural conditions with supplementary feeding. Adult stags run with groups of hinds under natural, uncontrolled mating; the facility does not practice hand-mating or artificial insemination. The birth-inspection reports cover fawns inspected after birth and contain no information on maternal age, body condition or dominance rank.

Data Source and Quality

Data consisted of 104 Berita Acara Pemeriksaan Kelahiran (official birth-inspection reports) issued between January 2019 and April 2024, each recording the birth date, sex, and identification numbers of the fawn, sire, and dam following physical inspection by facility staff. Sire identity was recorded for all 104 births. Five fawns lacked an identification number; fawn identity is not needed for any analysis, so all 104 births were retained. Two births (1.9%) came from dams without an identification number. Both fawns were female. These two births stayed in the herd-level sex-ratio tests but were left out of the parity analysis and the dam-level indicators, so those analyses use 102 births from 68 identified dams. The conclusion of the parity test would be the same whichever parity group the two births were placed in.

Two inter-birth intervals lasted less than 200 days. The species has a reported gestation of 258 to 260 days (Semiadi et al., 1994; Dahlan & Dawend, 2013), so neither interval can be the gap between two consecutive births of the same dam. Facility records

attribute both to eartags being reassigned. We removed the two intervals from the interval analysis only, and the dams concerned stayed in every other analysis. Had we kept these intervals, the dataset would contain biologically impossible values and the mean interval would be biased downward.

Definitions of Reproductive Indicators

Recorded parity separates the first birth recorded for each dam within the window from any births recorded after it. The herd was already breeding before 2019, so the first birth on record is not necessarily a dam's first parturition. Recorded parity therefore gives a birth's order within the observation window and should not be read as lifetime parity. The documented rebreeding proportion is the percentage of identified dams with at least two recorded births in the window. We do not use the term "repeat breeder", which in livestock reproduction refers to a female that fails to conceive after repeated services. The *inter-birth interval* (IBI) is the number of days between consecutive recorded births of the same dam and corresponds to the calving interval of livestock studies. Both dam-level indicators depend on follow-up: a dam first recorded late in the window had little opportunity to give birth again before 30 April 2024 (right truncation), and an unregistered birth would merge two true intervals into one recorded interval.

Statistical Analysis

Proportions of male births are reported with Wilson 95% confidence intervals (CI). Departure from 1:1 was tested with chi-square goodness-of-fit tests overall and for each year, with exact binomial tests as a check for small yearly samples; the five yearly p-values were adjusted with the Holm procedure (Holm, 1979). Heterogeneity among the five complete years (2019–2023) was tested with a 2 × 5 chi-square test of homogeneity, checked with a Monte Carlo permutation test (200,000 permutations) and summarised with Cramér's V. H1 was tested with a binomial logistic regression of sex on year (2019 = 0), reported as an odds ratio (OR) per year, with the p-value from the equivalent Cochran-Armitage trend test. Because only five yearly proportions inform the slope, we checked the fit for extra-binomial variation with the Pearson statistic and recomputed the standard error under a quasi-binomial dispersion. H2 and H4 were tested with 2 × 2 chi-square tests with Yates' continuity correction, confirmed with Fisher's exact tests, and expressed as OR with Woolf 95% CI. For H3, each of the four sires with at least 17 recorded offspring was tested against 1:1 with a goodness-of-fit test (Holm-adjusted

across the four sires), and all sires were compared jointly in a 2 × 5 contingency table in which the remaining 22 births were pooled as "other sires". To separate statistical from biological significance, we report effect sizes with CI (Nakagawa & Cuthill, 2007) and the minimum difference in proportion male that each comparison could detect with 80% power at $\alpha = 0.05$.

Eighteen dams contributed 52 of the 102 dam-identified births, so births are not fully independent. A mixed-effects logistic model with dam as a random effect is the usual remedy, but with 50 of 68 dams contributing a single birth the dam variance component would be estimated with little precision. We used two sensitivity analyses instead. First, we repeated the overall test on first recorded births only, which gives one birth per dam. Second, we divided the homogeneity statistic by a design effect, $1 + (m - 1)\rho$, in the manner of the Rao-Scott correction (Rao & Scott, 1981), where m is the size-weighted number of births per dam (1.98 to 2.45, a plausible range given the observed distribution of births among dams) and ρ is the within-dam correlation of offspring sex, set at 0.05 and 0.10 as deliberately generous values. The same correction can only enlarge p-values, so it was not needed for the tests that were already non-significant. Reproductive indicators were summarised as the documented rebreeding proportion (Wilson 95% CI) and the mean, standard deviation, 95% CI of the mean (t distribution), median and range of IBI. We did not fit a time-to-event model; instead, we bounded the effect of truncated follow-up by recalculating the rebreeding proportion after excluding dams that could have been first recorded in 2023–2024. All analyses were run in Python 3 with SciPy and statsmodels at $\alpha = 0.05$.

RESULTS AND DISCUSSION

1. Baseline Sex Ratio

Table 1 summarises annual births and sex ratio. The pooled ratio of 54 males to 50 females (proportion male 0.52, 95% CI 0.42–0.61) did not differ from 1:1 ($\chi^2 = 0.154$, $p = 0.695$), nor did the ratio among first recorded births alone, where each dam contributes one birth (33 males, 35 females; $\chi^2 = 0.059$, $p = 0.808$). The five complete years differed in sex ratio composition ($\chi^2 = 9.727$, $df = 4$, $p = 0.045$; permutation $p = 0.045$), and the effect size was moderate (Cramér's $V = 0.31$). This evidence is weak. Tested year by year, only 2022 deviated from 1:1 (22 males, 9 females; $p = 0.020$; exact binomial $p = 0.029$), and the deviation was no longer significant after Holm adjustment for the five

yearly tests (adjusted p = 0.098). In all four scenarios we examined, allowing for possible correlation between births of the same dam raised the p-value of the homogeneity test to between 0.055 and 0.075. We therefore regard the difference among years as a signal to keep monitoring. It is not an established shift. The next two sections test whether a simple explanation accounts for it.

Table 1. Annual births and offspring sex ratio of sambar deer at UPTD PTHPT Api-Api, January 2019–April 2024 (n = 104 births).

Year	Births (n)	Males	Females	% male (95% CI)	χ^2	p	Holm p
2019	26	10	16	38.5 (22.4–57.5)	1.385	0.239	0.719
2020	12	8	4	66.7 (39.1–86.2)	1.333	0.248	0.719
2021	20	7	13	35.0 (18.1–56.7)	1.800	0.180	0.719
2022	31	22	9	71.0 (53.4–83.9)	5.452	0.020*	0.098
2023	13	7	6	53.8 (29.1–76.8)	0.077	0.782	0.782
2024[a]	2	0	2	0.0 (0.0–65.8)	–	–	–
Total	104	54	50	51.9 (42.4–61.3)	0.154	0.695	–

[a]January–April only; excluded from the homogeneity and trend tests. CI = Wilson 95% confidence interval. χ^2 = goodness-of-fit test against 1:1 (df = 1). Holm p = p-value adjusted for the five yearly tests of 2019–2023. Exact binomial p for 2022 = 0.029. Homogeneity across 2019–2023: χ^2 = 9.727, df = 4, p = 0.045 (permutation p = 0.045), Cramér's V = 0.31; p = 0.055–0.075 after design-effect correction for repeated births of the same dam. *p < 0.05 before adjustment.

2. Testing for a Directional Trend Across Years

The proportion of male births in 2019–2023 was 0.385, 0.667, 0.350, 0.710 and 0.538 (Figure 2). The logistic slope was positive (β = 0.24 per year; OR 1.27, 95% CI 0.96–1.69) but not significant (Cochran-Armitage p = 0.096), so H1 was not supported. Two features of the data limit what this test can show. The slope rests on five annual points with 12 to 31 births each, whereas resolving trends of this size in other cervids has required decades of data and thousands of births (Hagen et al., 2022). The yearly proportions also scatter around the fitted line more than binomial sampling allows (Pearson χ^2 = 7.06, df = 3, p = 0.070; dispersion 2.35). Allowing for this extra variation widens the interval of the trend (OR 95% CI 0.63–2.58; p = 0.36). The variation in Table 1 is better described as irregular fluctuation, driven mainly by 2022, than as a linear drift, although a longer series is needed to separate the two.

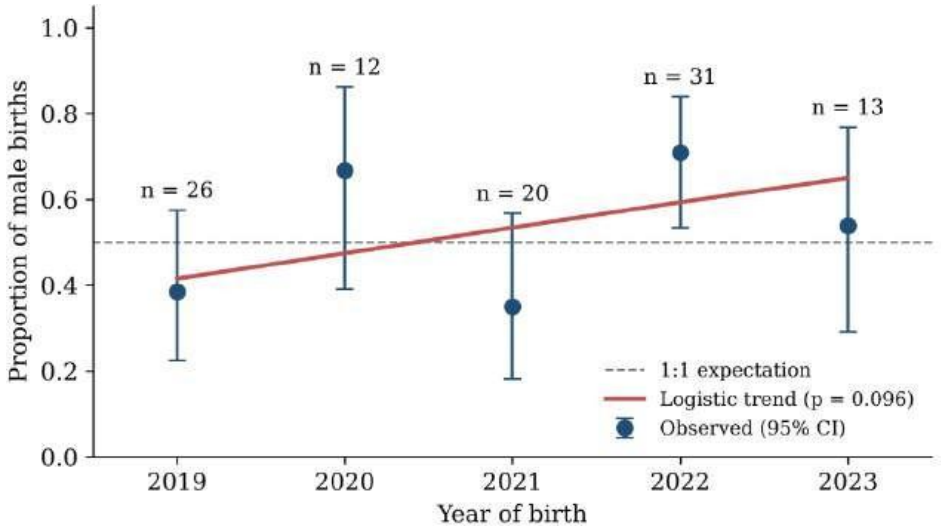


Figure 2. Proportion of male births by year at UPTD PTHPT Api-Api, 2019–2023 (n = 102 births). Points show observed proportions with Wilson 95% confidence intervals and the number of births per year; the solid line is the fitted logistic trend (OR 1.27 per year, p = 0.096); the dashed line marks the 1:1 expectation.

3. Candidate Correlates: Recorded Parity, Sire, and Birth Season

None of the three candidate correlates reached significance (Table 2, Figure 3), but the non-significant results rule out different things. Subsequent recorded births were more often male than first recorded births (61.8% versus 48.5%; OR 1.71, 95% CI 0.74–3.97; Fisher p = 0.293), and October–March births were more often male than April–September births (64.5% versus 46.6%; OR 2.09, 95% CI 0.88–4.97; Fisher p = 0.133). Each difference points the way that maternal-experience and birth-timing hypotheses would predict (Trivers & Willard, 1973; Hewison & Gaillard, 1999). Both confidence intervals range from no effect to at least a doubling of the odds of a male birth. Given these group sizes, a difference would have needed to reach roughly 28 to 29 percentage points for 80% power, while the observed differences were 13 and 18 points. H2 and H4 are therefore left unresolved. They have not been rejected. A change of 10 to 15 percentage points in the proportion of sons would affect the planning of future breeding stock, but a sample of this size could not detect it. Large studies of other cervids have not found a consistent parity or maternal-condition signal either (Hagen et al., 2022; Dentinger et al., 2024), so any such effect is probably small.

For the four leading sires, between 40.9% and 66.7% of offspring were male. After Holm adjustment none of them deviated from 1:1 (adjusted $p \geq 0.629$), and comparing all

sires together also gave no significant result ($\chi^2 = 5.25$, df = 4, p = 0.262; Cramér's V = 0.22). With 17 to 25 offspring per sire, the tests could pick up a bias only if a sire produced more than roughly 78% to 84% sons. The data therefore make a strong paternal effect unlikely, such as the effects reported in captive primates and numbats (Perret, 2018; Firman et al., 2023), but they leave room for a moderate one (Pawluk & Zagalska-Neubauer, 2024).

Two further caveats apply to the parity comparison. First, 18 dams had births in both groups, so part of the comparison is within dams, yet the chi-square test handles these paired observations as if they were independent. Second, the herd was already breeding before 2019, so the first-recorded group contains both primiparous hinds and multiparous hinds whose earlier births happened before the window. Each of these features would pull a real parity effect toward zero.

Season raises a different issue. Two decades ago, births at this facility peaked in June and July (Semiadi et al., 2005). The October–March group is therefore made up of off-peak births, and their mothers may differ in age or condition from hinds that calve at the peak. A test of several birth-timing hypotheses in kangaroos examined a similar pattern (Toni et al., 2021). The birth register does not allow this possibility to be tested here.

Table 2. Offspring sex ratio at UPTD PTHPT Api-Api by recorded parity, birth season and sire, January 2019–April 2024, with reference groups, test used and effect sizes.

Factor	Group	n	Males	% male (95% CI)	χ^2	p	OR (95% CI)
Recorded parity	First recorded birth (ref.)	68	33	48.5 (37.1–60.2)	1.107	0.293 [0.293]	1.00
	Subsequent recorded births	34	21	61.8 (45.0–76.1)			1.71 (0.74–3.97)
Birth season	April–September (ref.)	73	34	46.6 (35.6–57.9)	2.133	0.144 [0.133]	1.00
	October–March	31	20	64.5 (46.9–78.9)			2.09 (0.88–4.97)
Sire vs 1:1	0095	25	16	64.0 (44.5–79.8)	1.960	0.162 [0.629]	–
	0082	22	9	40.9 (23.3–61.3)	0.727	0.394 [0.788]	–
	0101	18	12	66.7 (43.7–83.7)	2.000	0.157 [0.629]	–
	0118	17	7	41.2 (21.6–64.0)	0.529	0.467 [0.788]	–
	Other sires (pooled)	22	10	45.5 (26.9–65.3)	–	–	–
Sire, joint test	Five sire groups	104	54	51.9 (42.4–61.3)	5.252	0.262	V = 0.22

Recorded parity gives the order of births within the 2019–April 2024 window and does not indicate lifetime parity; n = 102, as two births had no dam identification. Ref. = reference group. OR = odds of a male birth compared with the reference group, with Woolf 95% CI. Confidence intervals for proportions are Wilson intervals. Parity and season were compared with 2 × 2 χ^2 tests using Yates' correction, and the values in brackets are Fisher's exact p. Each sire was compared with the expected 1:1 ratio by a goodness-of-fit test (df = 1), and the values in brackets are p-values Holm-adjusted across the four sires. The joint sire test is a 2 × 5 χ^2 test of the four leading sires against all remaining sires pooled (n = 104; df = 4; Cramér's V = 0.22).

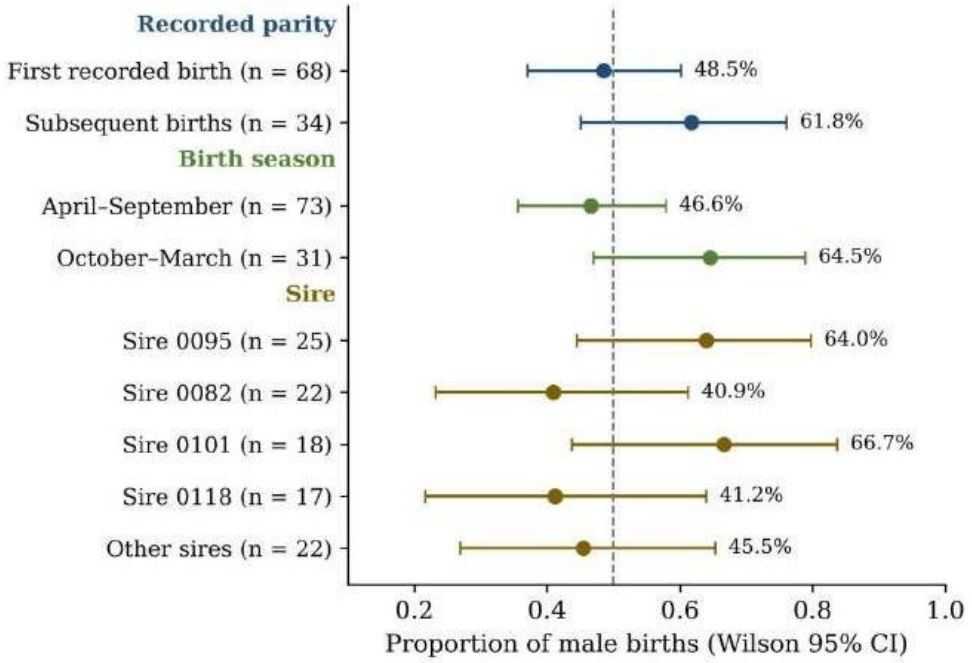


Figure 3. Proportion of male births (Wilson 95% confidence interval) by recorded parity, birth season and sire at UPTD PTHPT Api-Api, January 2019–April 2024. The number of births in each group is shown in brackets; the dashed line marks the 1:1 expectation.

Taken together, the three tests leave the among-year variation in Table 1 without an identified driver. The conclusion is limited: the tests show that no single factor produced a large effect, not that parity, sire or season have no effect. The register lacks maternal body condition, age and dominance rank, the variables that Trivers-Willard-type and paternal-effect hypotheses address directly (Sheldon & West, 2004; Pawluk & Zagalska-Neubauer, 2024). Recording a body condition score and an estimated age for each hind once a year, alongside the existing birth reports, would let a future analysis test the maternal-condition hypothesis directly instead of through recorded parity.

4. Reproductive Indicators: Documented Rebreeding and Inter-Birth Interval

Of 68 identified dams, 18 (26.5%, 95% CI 17.4–38.0) had two or more recorded births in the window (Table 3). This figure describes documented rebreeding within a 64-month window and does not measure reproductive success. Two features of the records bias it downward. Dams first recorded late in the window had little time to give birth again: the shortest observed interval was 330 days, so dams first recorded after mid-2023 could not have contributed a second birth. If all 15 births of 2023 and early 2024 belonged to newly recorded dams and those dams were excluded, the proportion would rise to 34.0% (18 of 53). Births that went unregistered, for instance a fawn that died before inspection, would also hide rebreeding that did occur. The rebreeding proportion within the window is therefore at least 26.5%; its exact value requires a time-to-event analysis that censors each dam at the end of her follow-up.

The 18 rebreeding dams supplied 32 analysable intervals after the two intervals shorter than 200 days were excluded (see Research Methods). The mean IBI was 515 days (SD 148; 95% CI 462–568), the median 441 days and the range 330 to 978 days. The mean exceeds the median by 74 days, so the distribution is right-skewed: a minority of long intervals, which may include missed births, pulls the mean upward. Truncation works in the opposite direction, because intervals that would have ended after April 2024 were never observed.

Table 3. Dam-level reproductive indicators for sambar deer at UPTD PTHPT Api-Api, January 2019–April 2024.

Indicator	Value	Note
Births with identified dam	102 of 104	Two births lacked dam identification
Identified dams	68	Denominator for dam-level indicators
Dams with one recorded birth	50 (73.5%)	Includes dams with short follow-up
Dams with ≥2 recorded births (documented rebreeding)	18 (26.5%; 95% CI 17.4–38.0)	Lower bound within the window
Documented rebreeding, restricted denominator	≤34.0% (18 of 53)	Extreme scenario; see note
IBI analysed	32 intervals from 18 dams	Two intervals <200 days excluded
Mean ± SD IBI	515 ± 148 days (95% CI 462–568)	Right-skewed distribution
Median (range) IBI	441 (330–978) days	–

IBI = inter-birth interval; CI = Wilson 95% interval for proportions and t-based 95% interval for the mean. The restricted denominator is an extreme scenario in which all 15 births of 2023–2024 came from newly recorded dams, each removed from the denominator.

Comparisons with published values must use the same statistic and allow for different conditions. Semiadi et al. (1994; 2005) reported means, so we compare means. The present mean of 515 days is 127 days longer than the 388 days reported for this herd two decades earlier (Semiadi et al., 2005) and 186 days longer than the 329 days of a farmed herd in New Zealand (Semiadi et al., 1994); the 95% CI of the present mean (462–568 days) excludes both values (Table 4). The same-site comparison is the more informative, because subspecies, climate and the broad husbandry system are shared. The New Zealand herd differed in latitude, photoperiod, pasture and management intensity, so that gap should not be read as a deficit of the same size. Neither comparison shows why the interval is longer. The data are consistent with three explanations. Lactational anoestrus may last longer than usual, the nutritional plane may be too low to support rapid rebreeding, or unregistered births may have joined two true intervals into a single recorded one. These remain hypotheses because the register cannot tell them apart. The low conception rate already reported for this herd (48.8%; Semiadi et al., 2005) suggests that rebreeding may have been slow for a long time. Each explanation can be checked in a specific way: score the body condition of lactating hinds, watch for estrus in hinds that have not calved within about 12 months of their last recorded birth, and register every parturition, including fawns that die before inspection. Direct behavioural observation, as Murad et al. (2023) used in semi-wild sambar, would provide the mating data needed to distinguish failure to conceive from loss of a pregnancy or a fawn.

5. Interpreting the Indicators Against Ex-Situ Breeding Criteria

Read against the criteria in the Literature Review, the indicators give a mixed and preliminary picture (Table 4). The pooled sex ratio lies close to the 1:1 expectation, so the herd is not producing a lopsided cohort of future breeders, although the among-year variation means the ratio needs annual review. Rebreeding is documented for at least a quarter of identified dams, so the herd is reproducing, but the mean interval is longer than both the herd's own earlier value and a captive benchmark from New Zealand, which suggests that reproduction is slower than it could be. These are early-warning indicators. They do not show whether the population is viable or self-sustaining in the sense of Lacy (2013), the IUCN/SSC (2014) guidelines and McGowan et al. (2017), because that judgement requires age-specific survival and fecundity, population growth rate and genetic parameters (Putnam et al., 2023), none of which a birth register provides. Genetic

diversity and the spread of reproduction among breeders are examined in the companion analyses of this herd; work on captive Formosan sambar shows that this dimension can limit a program even where reproduction appears adequate (Liang et al., 2023). A demographic model built from individual life histories, including deaths and transfers, would be the next step toward a formal viability assessment.

Table 4. Observed indicators at UPTD PTHPT Api-Api compared with the theoretical sex-ratio expectation (Panel A), published empirical benchmarks (Panel B) and reference values not measured in this study (Panel C).

Indicator	This study (observed)	Reference value	Type and source of reference	Comparability
A. Theoretical expectation				
Offspring sex ratio (proportion male)	0.52 (54 M : 50 F; 95% CI 0.42–0.61)	0.50	Fisherian expectation (theoretical)	Direct test; p = 0.695
B. Published empirical benchmarks				
Mean IBI	515 days (95% CI 462–568; n = 32)	388 days (SD 82; n = 33)	Same herd, earlier period; Semiadi et al. (2005)	High: same site; management may have changed
Mean IBI	515 days	329 days	Farmed herd, New Zealand; Semiadi et al. (1994)	Low: latitude, feeding and management differ
Median IBI	441 days	Not reported	–	No median-based benchmark available
Documented rebreeding proportion	≥26.5% of 68 dams	No equivalent	Conception rate of 48.8% in Semiadi et al. (2005) is a different measure	Not comparable
C. Reference values not measured here				
Gestation length	Not measured	258–260 days	Semiadi et al. (1994); Dahlan & Dawend (2013)	Used to screen implausible intervals
Age at first fawning	Not measured	≈32 months; ≈694 days	Dahlan & Dawend (2013); Semiadi et al. (2005)	Context only

Panel A compares an observed value with a theoretical expectation; Panel B compares observed values with empirical values from other studies, whose conditions differ as noted; Panel C lists values used only for data screening or context. IBI = inter-birth interval.

CONCLUSION

From January 2019 to April 2024, the overall offspring sex ratio of the sambar deer herd at UPTD PTHPT Api-Api stayed close to 1:1. The ratio varied modestly among years, largely because 2022 produced more males, but this variation weakened once the analysis was adjusted for multiple testing and for repeated births of the same dams, and no linear trend was confirmed. Offspring sex had no significant association with recorded parity, sire identity or birth season. The confidence intervals are still wide enough for moderate effects that a larger dataset might reveal. At least 26.5% of identified dams

rebred within the observation window, and the mean inter-birth interval of 515 days was longer than the 388 days previously reported for this herd. These indicators come from birth records and are preliminary. They do not settle whether the program has succeeded in conservation terms, because the study did not measure population viability, genetic diversity or long-term population growth. The findings lead to three practical steps: keep monitoring the sex ratio every year; record each hind's body condition and estimated age once a year together with the birth reports; and follow up hinds that have not calved within about 12 months of their last birth, making sure every parturition is registered. Together with the companion demographic and genetic analyses and, eventually, a complete life-history database, these records would allow a formal population viability analysis.

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