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## 2015-2016 PALIA ABUNDANCE ESTIMATES

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## ABSTRACT

The palila (*Loxioides bailleui*) population was surveyed annually during 1998–2016 on Mauna Kea Volcano to determine abundance, population trend, and spatial distribution. In the latest surveys, the 2015 population was estimated at 852–1,406 birds (point estimate: 1,116) and the 2016 population was estimated at 1,494–2,385 (point estimate: 1,934). Similar numbers of palila were detected during the first and subsequent counts within each year during 2012–2016; the proportion of the total annual detections in each count ranged from 46% to 56%; and there was no difference in the detection probability due to count sequence. Furthermore, conducting repeat counts improved the abundance estimates by reducing the width of the confidence intervals between 9% and 32% annually. This suggests that multiple counts do not affect bird or observer behavior and can be continued in the future to improve the precision of abundance estimates. Five palila were detected on supplemental survey stations in the Ka'ohē restoration area, outside the core survey area but still within Palila Critical Habitat (one in 2015 and four in 2016), suggesting that palila are present in habitat that is recovering from cattle grazing on the southwest slope. The average rate of decline during 1998–2016 was 150 birds per year. Over the 18-year monitoring period, the estimated rate of change equated to a 58% decline in the population.

## INTRODUCTION

The palila (*Loxioides bailleui*) is an endangered, seed-eating, finch-billed Hawaiian honeycreeper (a distinct group with family Fringillidae: subfamily Cardulinae) found only on Hawai'i Island. Once occurring on the islands of Kaua'i and O'ahu, as well as Mauna Loa and Hualālai volcanoes of Hawai'i, palila are now found only in subalpine, dry-forest habitats on Mauna Kea (Banko *et al.* 2002a). Previous analyses showed that palila numbers fluctuated throughout the 1980s and 1990s, but since 1998 palila have declined and they appear to have declined steadily since 2003 (Jacobi *et al.* 1996, Johnson *et al.* 2006, Leonard *et al.* 2008, Banko *et al.* 2009, Gorresen *et al.* 2009, Banko *et al.* 2013, Camp *et al.* 2014).

Palila tend to move up and down the western slope of Mauna Kea seasonally as they track the availability of their main food, seeds of the endemic māmane (*Sophora chrysophylla*) tree (Hess *et al.* 2001). During population surveys, usually in late January, māmane seedpods are most abundant at higher elevations, but seedpod abundance increases at lower elevations by May (Banko *et al.* 2002b). Although the distribution of palila shifts in response to food availability, the areas that are occupied seasonally overlap extensively, and the area that is surveyed each winter provides a stable and representative basis for evaluating population abundance and trends.

The aim of this report is to update abundance estimates and population trends for the palila since 1998, based on the 2015 and 2016 surveys. Additional transects were added to the original Hawaiian Forest Birds Survey (HFBS) transects in 1998 to produce a more precise population estimate and provide more complete coverage of the palila distribution during the survey period (Johnson *et al.* 2006).

## METHODS

### **Bird Sampling**

Since 1980, 95% of the palila population has occurred in a 64.4 km<sup>2</sup> area on the southwestern slope of Mauna Kea (Scott et al. 1984, Banko et al. 2013; Figure 1). We refer to this area hereafter as the “core survey area.” During 27 January–18 February 2015 and 2–9 February 2016, point-transect sampling was conducted on Mauna Kea to estimate palila abundance and range. In 2015 and 2016, 13 bird survey transects inside the 64.4 km<sup>2</sup> palila core survey area (transects 101–108, 122–126, Figure 1) were surveyed one or more times. Prior to 2008 surveys were conducted mountain-wide, but lack of detections outside the southwest slope led to focusing survey effort to the core palila habitat, with the intent to survey the entire mountain every five years, starting in 2012 (D. Leonard, pers. comm.). In addition to surveying the core habitat, supplemental stations adjacent to it were surveyed in 2015 and 2016 to look for possible range expansion. These included transects 109, 110 and 116 in 2015, and 109, 110, and 114 in 2016, plus additional stations in the lower portions of transects 101, 102, 124 and 125 in the Ka’ohe restoration area (Banko *et al.* 2009) in both years. Within the core survey area, the 2015 survey consisted of 839 counts at 420 stations, while in 2016 the survey consisted of 837 counts at 420 stations (Figure 1, Table 1).

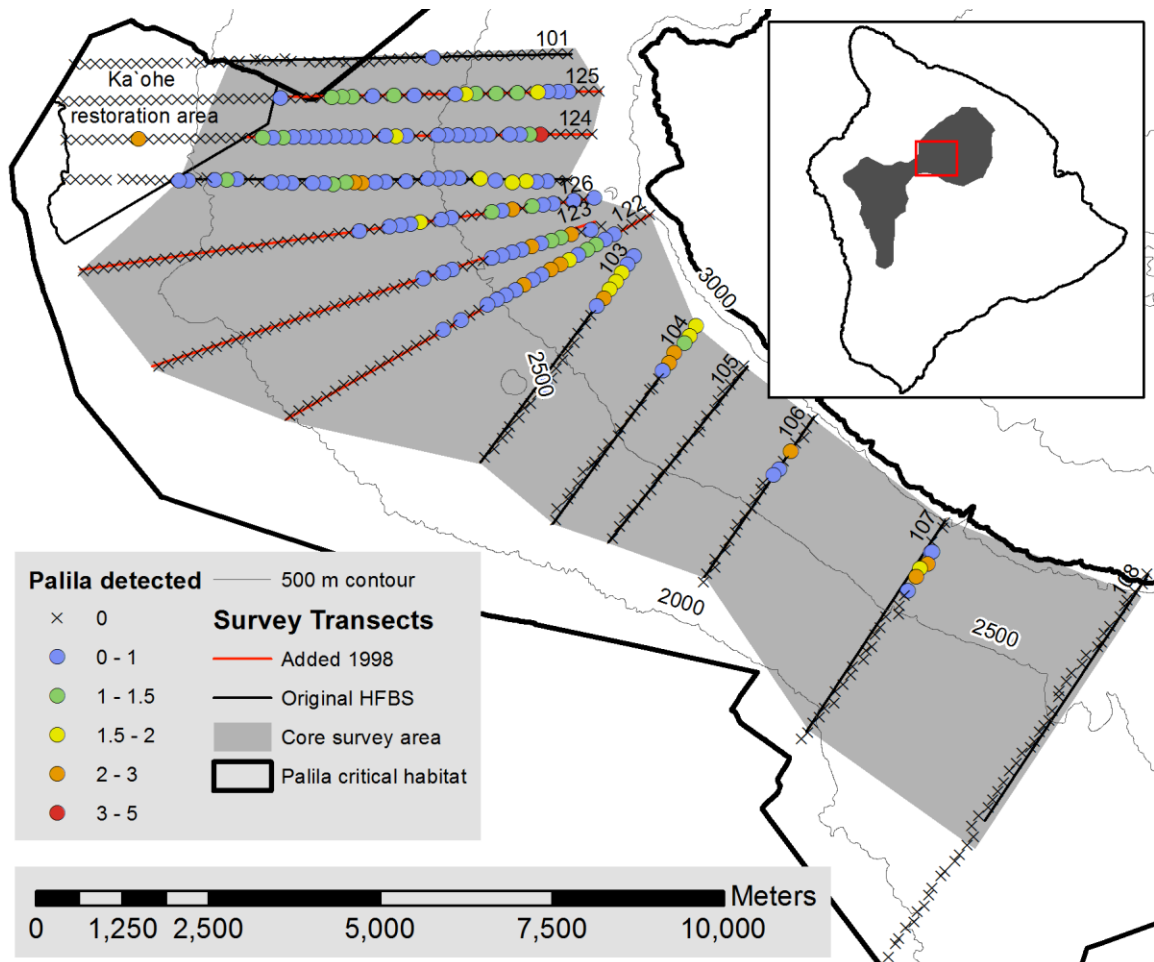


Figure 1. Palila detected per count during 2015–2016 surveys (mean 2.7 counts/occupied station, minimum 1, maximum 4). × symbols mark stations where no palila were detected regardless of survey effort. The total survey area of the palila population is demarcated in the light gray shaded region. Black lines depict the original HFBS (Hawai'i Forest Bird Survey) transects, with red lines indicating the five transects added in 1998. Transects are numbered by date of creation, not geographically. The inset map shows the historical palila range on Hawai'i Island.

In 2015, a majority of the stations ( $n = 415$ ) within the core survey area were counted twice, while one station was counted once and four stations were counted three times. In 2016, a majority of the stations ( $n = 416$ ) were counted twice, while three stations were counted once and one station was counted three times. At the supplemental stations, a majority (62 of 63 stations; 125 counts) were counted twice in 2015. In 2016 only 18 of the 79 supplemental stations were counted twice for a total of 97 counts.

Most Hawaiian forest bird surveys last eight minutes (Camp *et al.* 2009), however six minutes is used for Mauna Kea counts because their woodland habitat is more open than mesic and wet forest habitats, allowing for easier and more rapid detection. Counts commenced at sunrise and continued up to four hours (approximately 11:00 HST). During each count, trained and

calibrated observers recorded the species, detection type (heard, seen, or both), and distance of each bird from the survey station center. Time of sampling and weather conditions (cloud cover, rain, wind, and wind gust [hereafter gust]) were also recorded, and surveying was postponed when conditions hindered the ability to detect birds (wind and gust >20 kph or heavy rain).

Table 1. Number of transects and stations sampled by year inside the core survey area and supplemental to it.

Year	Inside core survey area			Supplemental survey effort	
	Transects	Stations	Counts	Stations	Counts
1998	12	355	357	186	186
1999	13	414	418	192	212
2000	13	418	424	224	224
2001	13	414	416	221	223
2002	13	416	417	270	271
2003	13	403	403	258	258
2004	13	397	397	240	251
2005	13	402	428	340	351
2006	13	386	398	323	356
2007	12	387	412	256	256
2008	12	386	432	0	0
2009	13	416	416	0	0
2010	13	415	420	0	0
2011	13	411	432	0	0
2012	13	486	909	360	360
2013	13	418	889	0	0
2014	13	443	887	79	99
2015	13	420	839	63	125
2016	13	420	837	79	97

### Abundance Estimation

Distance analysis fits a detection function to estimate the probability of detecting a bird at a given distance from the observer. This detection function is fitted with covariates, accounting for the effect of the observer, detection type, weather conditions, and year. With each additional year of data, estimates of these effects become more precise, and the improved detection function may cause population estimates of previous years to change slightly.

Density estimates (birds/km<sup>2</sup>) were calculated from point-transect sampling data using program DISTANCE, version 6.0, release 2 (Thomas *et al.* 2010). The 2015 and 2016 data were pooled with detections from previous surveys since 1998. Candidate models were limited to half-normal

and hazard-rate detection functions with expansion series of order two (Buckland *et al.* 2001: 361, 365). Survey effort was adjusted by the number of times the station was counted. To improve model precision, potential sampling covariates were evaluated with the multiple covariate distance sampling engine of DISTANCE (Thomas *et al.* 2010). Covariates included the weather conditions, time of sampling, type of detection, observer, and year of survey. Right-tail truncation was set at 85.5 m, the distance where the detection probability was approximately 10%. This procedure facilitates modeling by deleting outliers and reducing the number of adjustment parameters needed to modify the detection function. The detection probability model selected was the one having the lowest corrected Akaike's Information Criterion (AICc; Buckland *et al.* 2001, Burnham and Anderson 2002; Figure 2, Table 2). Annual population densities for each survey were calculated using the global detection function, and the pooled data were post-stratified by year and location (inside/outside core survey area). The 95% confidence intervals for the annual density estimates were derived from the 2.5 and 97.5 percentiles using bootstrap methods in DISTANCE for 999 iterations (Buckland *et al.* 2001, Thomas *et al.* 2010). Population abundance estimates were the product of the density estimate times the area of the core survey area (64.4 km<sup>2</sup>).

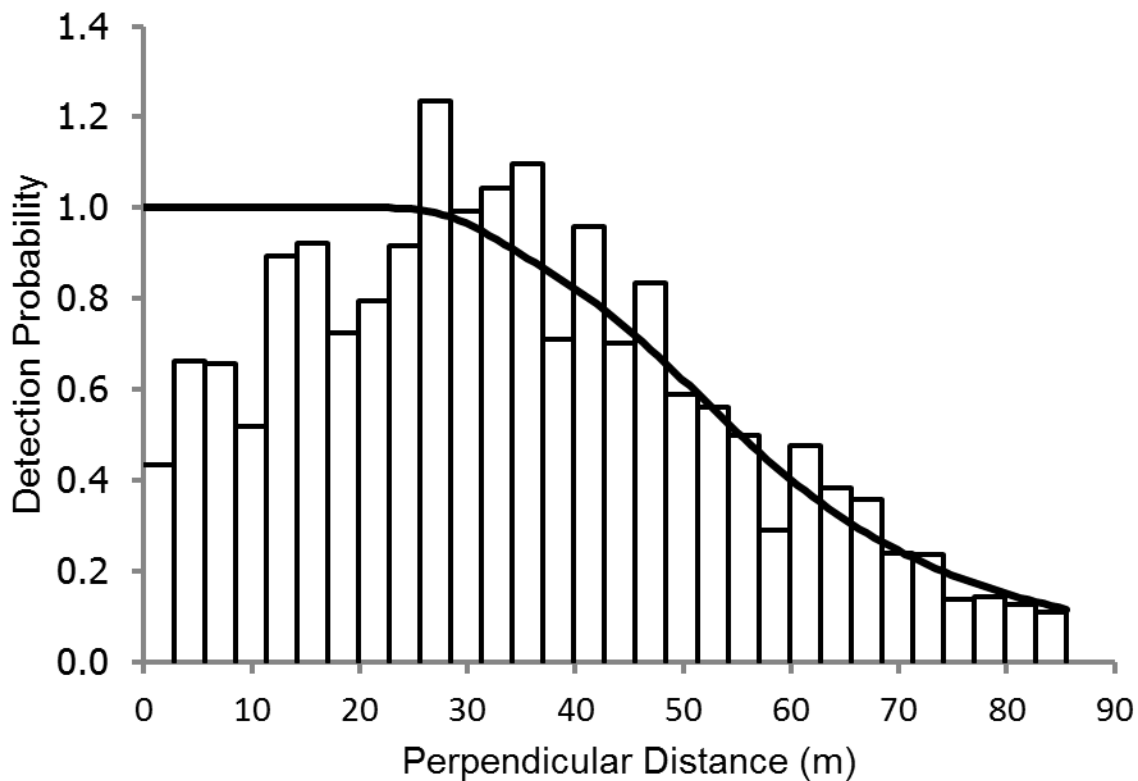


Figure 2. Hazard-rate detection function, with no expansion series and including the covariate detection type (heard versus pooled visual; line) and palila distance data (histogram) pooled across all surveys from 1998 to 2016. Data were truncated at 85.5 m.

Table 2. Results of fitting 17 detection function models to the 1998–2016 palila distance histogram.  $\Delta\text{AICc}$  is the difference in AICc scores between each model and the overall best-fit model, and  $w_i$  is the discrete model probability. The hazard-rate detection function with no adjustment terms and detection type (heard only versus pooled visual) as a covariate was chosen (bold) over the same base model but with the unpooled detection type covariate because the former was within 2 AICc units and more parsimonious.

Model <sup>1,2</sup>	# parameters	Ln(Likelihood)	AICc	$\Delta\text{AICc}$	w
H-rate Key DetType	4	21512.99	43033.98	0	0.41812
<b>H-rate Key VBDetType</b>	3	21514.04	43034.08	0.10	0.39773
H-rate Key Year(f)	5	21515.37	43035.62	1.64	0.18415
H-rate Key HBDetType	3	21628.84	43263.69	229.71	0.00000
H-rate Key Gust	5	21645.60	43303.21	269.23	0.00000
H-rate Key Wind	4	21651.16	43312.62	278.64	0.00000
H-rate Key Obs	5	21654.22	43318.20	284.22	0.00000
H-rate Key	2	21663.24	43330.49	296.51	0.00000
H-rate Key Time	3	21662.60	43331.20	297.22	0.00000
H-rate Key Rain	3	21662.51	43331.30	297.32	0.00000
H-rate Key Cloud	4	21661.65	43331.31	297.33	0.00000
H-norm Key	1	21728.96	43459.92	425.94	0.00000

<sup>1</sup> Models are hazard-rate (H-rate) and half normal (H-norm); adjustment terms are cosine (Cos), simple polynomial (S-poly) and hermite polynomial (H-poly); and covariates are cloud cover (Cloud), detection type (DetType = heard, visual or both, HBDetType = pooled heard and both, VBDetType = pooled visual and both), gust strength (Gust), observer (Obs), time of detection (Time), wind strength (Wind), and year (Year; continuous = c and factor = f).

<sup>2</sup> Models H-norm Cos, H-norm H-poly, H-rate Cos, H-rate S-poly, and H-rate Key Year(c) failed to converge.

## Trend detection

Trend in palila abundance was assessed with a log-linear regression model in a Bayesian framework. The long-term trend from 1998–2016 was derived from the posterior probability of the slope, following Camp *et al.* (2008). Diagnostics included visual inspection of residual plots, revealing a pattern in the residuals due to a possible step change in abundance; but a change point analysis (Erdman and Emerson 2007) did not identify any change point with greater than 32% posterior probability. A Shapiro-Wilk test of normality did not reject the null hypothesis of a normal distribution in the log-scale residuals ( $W = 0.94$ ,  $P = 0.25$ ). There was significant lag-1 autocorrelation (common in population time series), but comparing corrected AIC values showed an increase of 2.29 units from simple regression to the AR1 model. Given the borderline AIC difference and a biological expectation of autocorrelation, we chose to use the AR1 regression model.

The AR1 log-linear regression model was fit using Stan (Stan Development Team 2015) running from an R environment (R Core Team 2015). Regression parameters for slope, intercept, and auto-correlation were given diffuse normal priors, and regression error a diffuse uniform prior. Model parameters were estimated from 2,500 iterations for each of four chains after first

discarding 1,000 iterations as a “warm-up” period. The four chains were pooled (10,000 total samples) to estimate the posterior distributions. Gelman-Rubin convergence statistics for all estimated parameters were less than the 1.1 threshold indicating convergence (Gelman et al. 2004)

We then used the posterior distribution of the slope to make an inference about long-term population trend. We used an equivalence testing approach with a population change of 25% over 25 years as biologically significant thresholds. This corresponds to an annual rate of change of -0.0119 or 0.0093 on the log-scale. A biologically meaningful change occurs when the posterior distribution of the slope occurs outside the equivalence region, whereas a negligible trend occurs when the slope is within the equivalence region. A statistically significant trend that occurred within the threshold levels was deemed a biologically negligible trend. An inconclusive result occurs when small sample size and high variation in estimates results in the posterior distribution of the slope providing weak evidence in the three outcomes (i.e., less than 50% probability in any of the three outcomes, respectively; Camp *et al.* 2008).

### **Repeat Surveys**

Most stations within the core survey area were sampled twice between 2012 and 2016. Multiple counts increase the numbers of detections, thereby reducing the total uncertainty in the abundance estimates and improving the overall power to detect population changes. Using AIC, the model with covariates of count number (i.e., first versus subsequent counts) and year was compared to the model with year alone. This approach compared the fit of each model to the data to determine whether the detections on the first *vs.* subsequent counts differed.

## **RESULTS**

### **Abundance**

Within the 64.4 km<sup>2</sup> core survey area of Palila Critical Habitat on the southwestern flank of Mauna Kea, 40% more palila were detected in 2016 than in 2015 (192 in 2015 and 319 in 2016; Table 3). An additional five palila were detected on the lower extensions of transects 124 and 125 (one in 2015 and four in 2016; Figure 1, Table 3). In 2015, the palila population in the core survey area was estimated at 852–1,406 birds (point estimate of 1,116; Figure 3, Table 3). In 2016, the palila population in the core area was estimated at 1,494–2,385 (point estimate of 1,934). The model that best fit the distance histogram was a hazard-rate detection function with no adjustment terms and detection type (heard only versus pooled visual [seen only or heard first but later visually confirmed]) as a covariate (Figure 2 and Table 2).

### **Trend**

The posterior distribution of the log-scale slope had a 95% credible interval of -0.101 to -0.043, with a mean of -0.074. The posterior distribution of the slope was overwhelmingly below the decline threshold (-0.012) with 99.97% probability in the decline region, and only 0.03% of posterior probability falling into the stable region. Therefore, there was very strong evidence that the palila population experienced a long-term decline in abundance between 1998 and 2016.

Table 3. Annual palila detections and population estimate parameters. Detections are given for palila recorded inside the core survey area and for supplemental adjacent stations during six-minute counts. Population parameters include the population estimate, % coefficient of variation (CV), standard error (SE), and lower and upper limits of the 95% confidence interval inside the core survey area.

Year	# Detections Inside	# Supplemental detections	Estimate	%CV	SE	Lower Limit	Upper Limit
1998	313	2	4,634	9.82	455	3,786	5,565
1999	388	1	5,310	9.42	500	4,397	6,370
2000	234	14	3,079	10.30	317	2,445	3,741
2001	345	4	4,589	9.60	440	3,738	5,472
2002	339	9	4,616	9.56	441	3,794	5,560
2003	439	7	5,851	8.86	518	4,930	6,922
2004	371	9	5,013	8.55	429	4,182	5,889
2005	315	1	4,214	9.86	415	3,436	5,077
2006	271	16	3,909	10.18	398	3,161	4,734
2007	210	3	2,932	10.51	308	2,385	3,570
2008	186	na	2,659	10.40	277	2,172	3,262
2009	189	na	2,411	11.57	279	1,881	3,033
2010	151	na	1,610	12.06	194	1,267	2,044
2011	119	na	1,338	14.25	191	989	1,745
2012 <sup>1</sup>	362	0	2,133	10.89	232	1,698	2,619
2013 <sup>2</sup>	337	na	1,756	9.02	158	1,454	2,080
2014 <sup>3</sup>	351	0	2,017	9.67	195	1,654	2,428
2015 <sup>4</sup>	192	1	1,116	12.63	141	852	1,406
2016 <sup>5</sup>	319	4	1,934	11.53	223	1,494	2,385

<sup>1</sup> Of 362 total detections, 194 recorded on first count, 168 recorded on subsequent counts.

<sup>2</sup> Of 337 total detections, 178 recorded on first count, 159 recorded on subsequent counts.

<sup>3</sup> Of 351 total detections, 163 recorded on first count, 188 recorded on subsequent counts.

<sup>4</sup> Of 192 total detections, 99 recorded on first count, 93 recorded on subsequent counts.

<sup>5</sup> Of 319 total detections, 178 recorded on first count, 141 recorded on subsequent counts.

### Repeat Surveys

The numbers of palila detected during first and subsequent counts varied, but the proportions of the total detections on first counts were similar (Table 3): 54% (2012), 53% (2013), 46% (2014), 52% (2015), and 56% (2016). Using AIC statistics to assess the fit of detection probabilities from models with and without the covariate year, the reference model incorporating year as a covariate had an AICc value of 43035.62 (Table 2). Including count (first vs. subsequent) as an additional covariate inflated the AICc value by 171 units, indicating a decrease in model fit. Thus, there was no evidence that detection functions for birds detected in the first vs. subsequent counts were different. Furthermore, conducting repeat counts improved the abundance estimates by reducing the width of the confidence intervals by 9–32%: 10% (2012), 17% (2013), 21% (2014), 32% (2015), and 9% (2016).

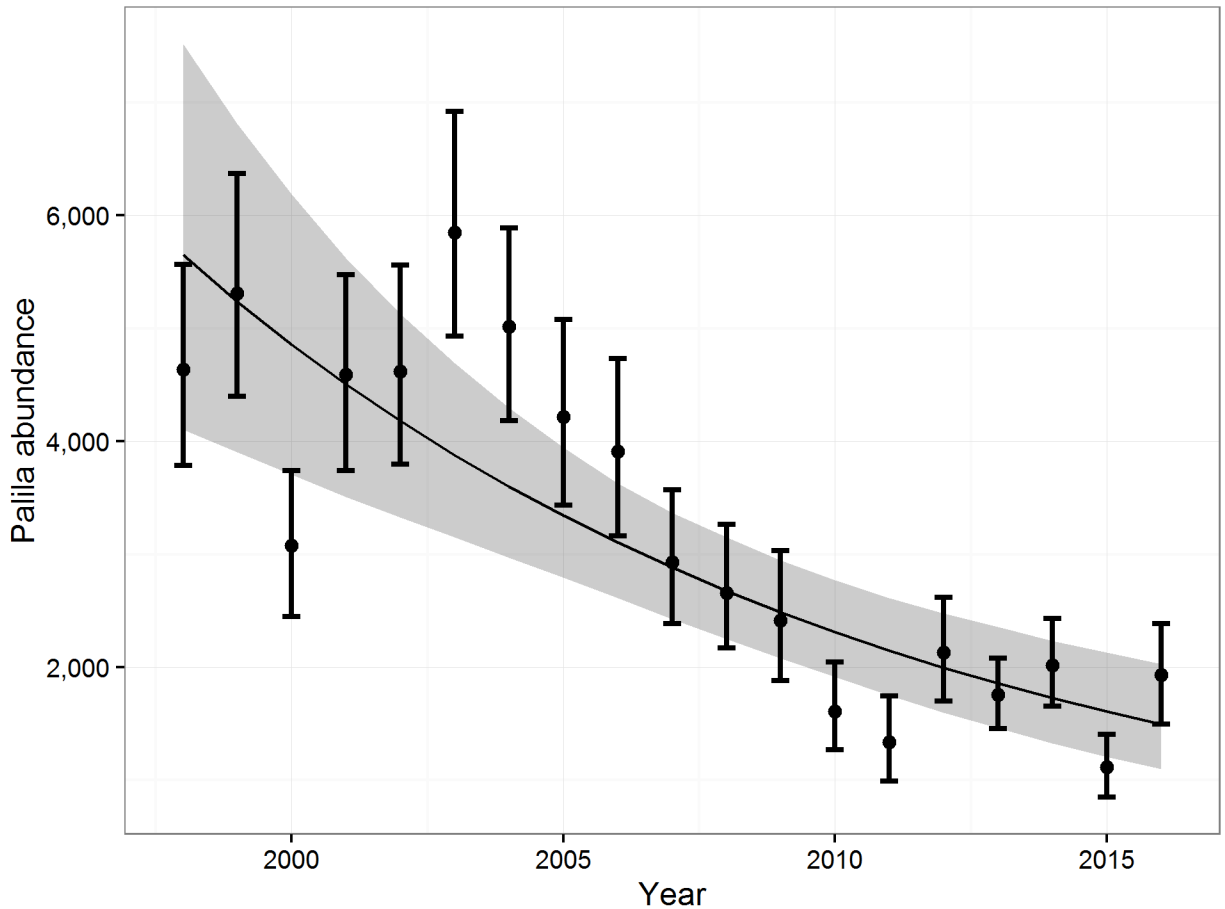


Figure 3. Annual palila population estimates and 95% confidence intervals inside the core survey area on the western slope of Mauna Kea.

### CONCLUSIONS

The 2016 palila population was estimated at 1,494–2,385 birds (point estimate of 1,934). There was very strong evidence that the palila population declined after 1998, with the greatest decline occurring after 2003. The average rate of decline during 1998–2016 was 150 birds per year, resulting in a 58% decline in the population over the 18-year period.

Increasing the number of counts at a station, going from one to two counts for example, increases the numbers of birds detected. This is advantageous because increased numbers of detections tend to remove spikes and dips in the detection histogram and reduces estimator uncertainty. This process can be negated if there are marked differences between the counts, for example if birds are frightened away from the survey area during the first count or if survey conditions are strikingly different. However, similar numbers of palila were detected during the first and subsequent counts during 2012–2016. In addition, the detection model was not

improved by including the count covariate, indicating that there was no difference in the numbers of palila detected or in their detection probability due to repeat counts. Thus, greater precision in population estimates can be gained if surveys continue to include repeat counts.

### **MANAGEMENT IMPLICATIONS**

Repeat counts reduced the width of the confidence intervals by 200 birds with the greatest improvements occurring in years with some of the lowest abundances. Repeated counts will become increasingly important if the palila population declines further and numbers of detections per count are reduced. If the population continues to decline, modelling the threshold at which two repeated counts would be insufficient to provide comparable or acceptable confidence in the precision of the estimate could provide useful guidance for planning future counts.

Compared to the 2013 and 2014 surveys there were no palila detected along the southeast and southwest peripheries (Figure 1). Similar year-to-year changes in distribution have been seen before, and the pattern seen in the 2015 and 2016 surveys would not change the conclusion that the range has not contracted (see Figure 4 and Table 3 in Camp *et al.* 2014). However, current palila range is only about 5% of its historical extent (Figure 1 inset; Banko *et al.* 2013).

The survey plan for Mauna Kea schedules a mountain-wide survey of palila critical habitat every 5 years, with the next one due in 2017. Carrying out the scheduled mountain-wide survey may detect palila range expansion beyond the core habitat. For example, in 2015 and 2016, no palila were detected on transect 109, which is adjacent to the core survey area on its southeastern edge, but one of the observers reported an incidental palila detection on transect 109 near station 5 a week after the 2016 survey (B. Rownd, pers. comm.).

Additionally there have been anecdotal observations of palila in the Ka'ohē restoration area (K. Brinck, pers. obs.), but 2016 was the first year that palila were recorded on the Ka'ohē parcel during the survey. Continued surveys of the Ka'ohē area would identify if palila are expanding there, and if so, the area could be added to annual abundance estimates as a separate stratum.

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