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PALILA ABUNDANCE ESTIMATES AND TREND

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ABSTRACT

The palila (*Loxioides bailleui*) population was surveyed annually during 1998−2014 on Mauna Kea Volcano to determine abundance, population trend, and spatial distribution. In the latest surveys, the 2013 population was estimated at 1,492−2,132 birds (point estimate: 1,799) and the 2014 population was estimated at 1,697−2,508 (point estimate: 2,070). Similar numbers of palila were detected during the first and subsequent counts within each year during 2012−2014, and there was no difference in their detection probability due to count sequence. This suggests that greater precision in population estimates can be achieved if future surveys include repeat visits. No palila were detected outside the core survey area in 2013 or 2014, suggesting that most if not all palila inhabit the western slope during the survey period. Since 2003, the size of the area containing all annual palila detections do not indicate a significant change among years, suggesting that the range of the species has remained stable; although this area represents only about 5% of its historical extent. During 1998−2003, palila numbers fluctuated moderately (coefficient of variation [CV] = 0.21). After peaking in 2003, population estimates declined steadily through 2011; since 2010, estimates have fluctuated moderately above the 2011 minimum (CV = 0.18). The average rate of decline during 1998−2014 was 167 birds per year with very strong statistical support for an overall declining trend in abundance. Over the 16-year monitoring period, the estimated rate of change equated to a 68% decline in the population.

INTRODUCTION

The palila (*Loxioides bailleui*) is an endangered, seed-eating, finch-billed honeycreeper found only on Hawaiʻi Island. Once occurring on the islands of Kauaʻi and Oʻahu, as well as Mauna Loa and Hualālai volcanoes of Hawaiʻi, palila are now found only in subalpine, dry-forest habitats on Mauna Kea Volcano (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, Leonard *et al.* 2008, Banko *et al.* 2009, Gorresen *et al.* 2009, Banko *et al.* 2013).

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 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 for the palila based on the 2013 and 2014 surveys. We assess the palila long-term trend during 1998–2014. This period covers the population trend for the entire time series since additional transects were established (Johnson *et al.* 2006). These additional transects were established to produce a more precise population estimate and provide more complete coverage of the palila distribution during the survey period. Stations in the palila habitat within the survey area were sampled twice during the 2012–2014 surveys, thus allowing us to address the question of how repeat samples improve estimate precision.

Bird Sampling

During 7 January–7 February 2013 and 14–24 January 2014, point-transect sampling was conducted on Mauna Kea to estimate palila abundance and range. In 2013, 13 bird survey transects inside the 64.4 km^2 palila core survey area (transects 101–108, 122–126) were surveyed one or more times. In 2014, the same 13 transects within and an additional 5 transects outside (transects 109–121, 127–133) the core survey area were surveyed one or more times. The 2013 survey consisted of 889 counts at 418 stations, while in 2014 the survey consisted of 986 counts at 522 stations (Figure 1, Table 1).
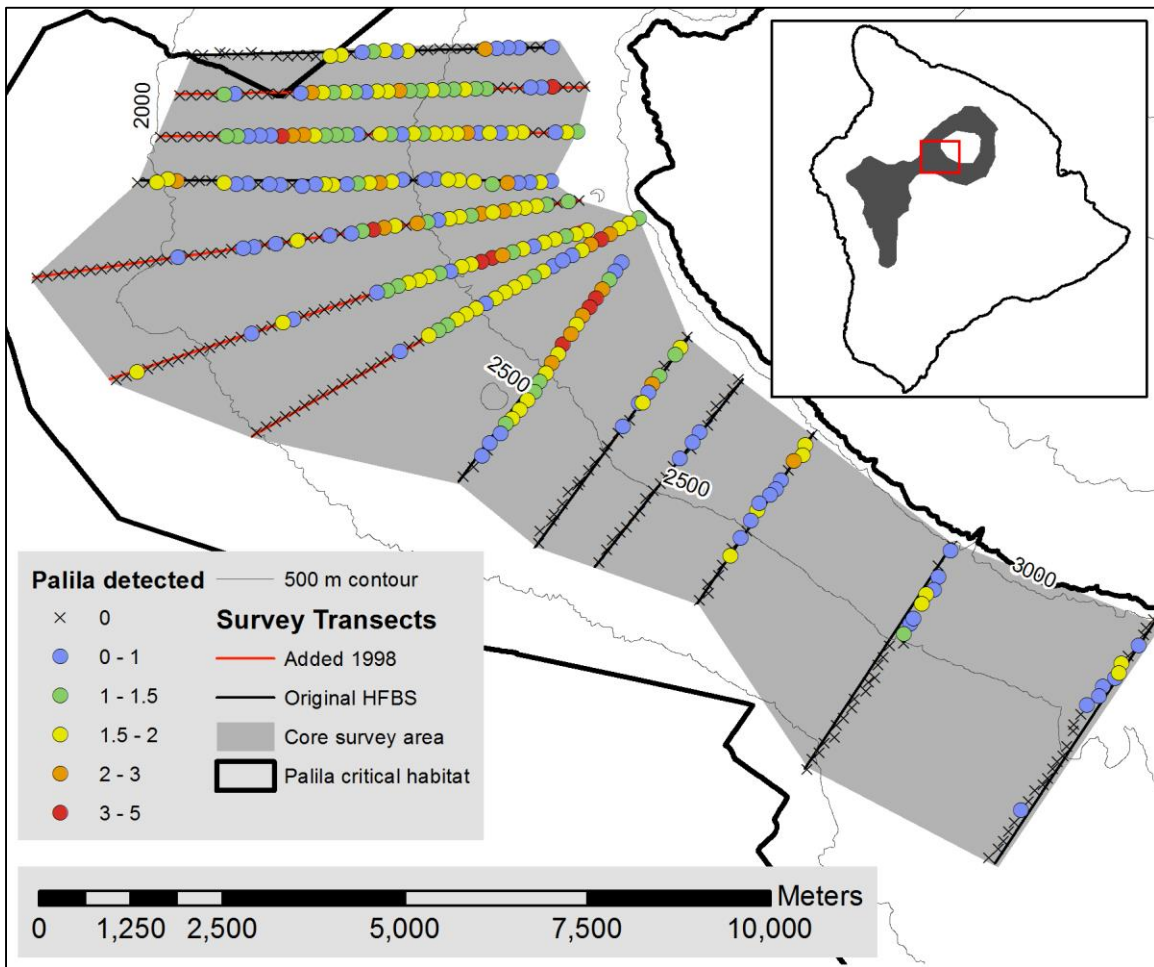
Figure 1. Palila detected per each visit during 2013–2014 surveys (mean 2.2 visits/station, minimum 2, maximum 6). X symbols mark stations where no palila were detected regardless of survey effort. The total survey area of the palila population is demarcated in the shaded region. Lines depict the original HFBS (Hawaiʻi Forest Bird Survey) transects (black) plus those added in 1998 (red). The inset map shows the historical palila range on Hawaiʻi Island.

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

Year	Inside core survey area			Outside core survey area		
	Transects	Stations	Counts	Transects	Stations	Counts
1998	12	355	357	14	186	186
1999	13	414	418	14	192	212
2000	13	418	424	17	224	224
2001	13	414	416	17	221	223
2002	13	416	417	20	270	271
2003	13	403	403	20	258	258
2004	13	397	397	18	240	251
2005	13	402	428	20	340	351
2006	13	386	398	20	323	356
2007	12	387	412	20	256	256
2008	12	386	432	0	0	0
2009	13	416	416	0	0	0
2010	13	415	420	0	0	0
2011	13	411	432	0	0	0
2012	13	486	909	20	360	360
2013	13	418	889	0	0	0
2014	13	443	887	5	79	99

In 2013, a majority of the stations were counted twice (357 stations), while 4 stations were counted once and 57 stations were counted three times. In 2014, a majority of the stations were counted twice (456 stations), while 62 stations were counted once and 4 stations were counted three times. Within the core survey area, a majority of the stations were counted two or more times (443 stations; 887 counts) in 2014. A total of 99 counts at 79 stations was conducted on transects outside the core survey area.

Most Hawaiʻi Forest Bird Surveys last eight minutes (Camp *et al.* 2009), however, six minutes is used for palila 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).

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^2) were calculated from point-transect sampling data using program DISTANCE, version 6.0, release 2 (Thomas *et al.* 2010). The 2013 and 2014 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). The uniform detection function was not considered. Survey effort was adjusted by the number of times the station was counted. To improve model precision, sampling covariates were incorporated in 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 Akaike's Information Criterion corrected for small sample size (AICc; Buckland *et al.* 2001, Burnham and Anderson 2002). 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^2).

Trend Detection
The trend in palila abundance was assessed in the time series using log-linear regression models. In a Bayesian framework, regression was used to assess the long-term population trend (1998−2014), where the evidence of a trend was derived from the posterior probability of the slope using a log-link regression model, following Camp *et al.* (2008). Diagnostics demonstrated that the log-linear regressions of trends met all model assumptions (visual inspection of residual plots; Shapiro-Wilk normality test, W = 0.9378, *P* = 0.29), except that temporal autocorrelation was evident (Figure 2). The lag-one autoregressive model had the lowest AIC value and was a substantially better fit than a model that assumed temporal independence (lag-zero; 11.2 AIC units greater). The abundance at time *t* was dependent on the previous population abundance at time *t-1*, and the simple linear regression model was modified to account for temporal dependence using an autoregressive moving average model of lag one (ARMA (1, 1)). Standard autoregressive models require that the time series is stationary (i.e., no trend exists). Our ARMA model allows for non-stationary time series where the trend incorporates correlation between observations in subsequent years.

We used log-linear regression with ARMA (1, 1) adjustments for autocorrelation in a Bayesian framework to assess population trends in abundance. In the regression model we estimate four parameters: the intercept (α), the slope (β), the log-normal error of standard deviation (σ), and a measure of the correlation between abundance in successive years (ϕ). The posterior of the slope was estimated using Stan (Stan Development Team 2013) running from an R environment (R Development Team 2014). The parameters α, β, and ϕ were given uninformative normal priors, and an uninformative uniform prior was given for σ. In Bayesian analysis, uninformative prior distributions are chosen so the posterior distributions reflect only the observed data. The trends were centered on the year 2006 to improve model convergence. The model parameters were estimated from 500 iterations for each of four chains (i.e., model runs) after first discarding 500 iterations as a "warm-up" period. The four chains were pooled (2,000 total samples) to calculate the posterior distribution. Gelman-Rubin convergence statistics for all estimated parameters were below 1.001; less than the 1.1 threshold that indicates convergence (Gelman and Rubin 1992).
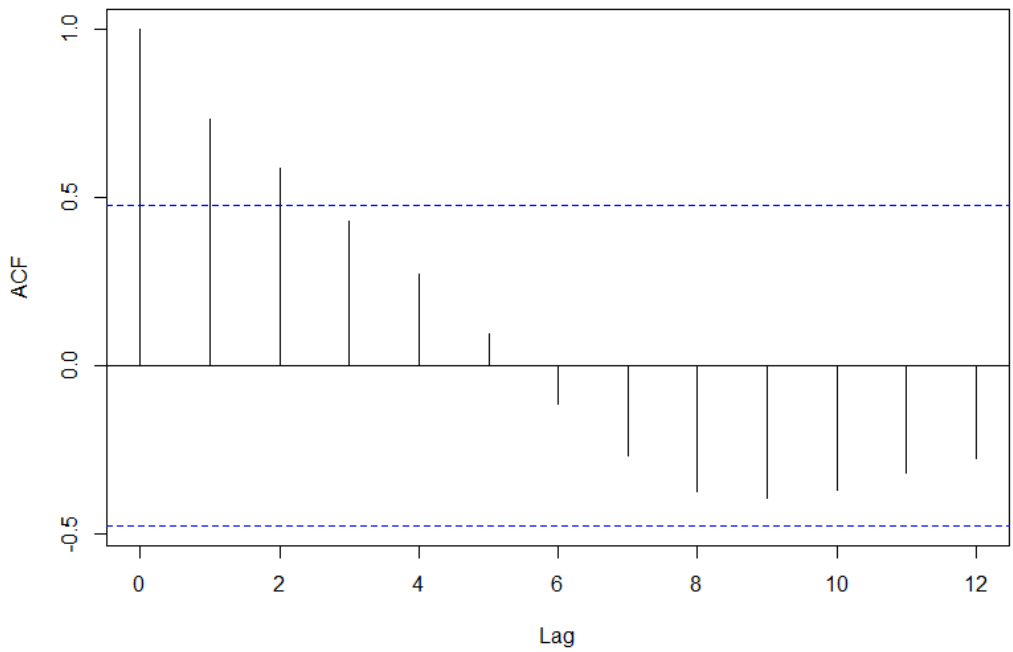
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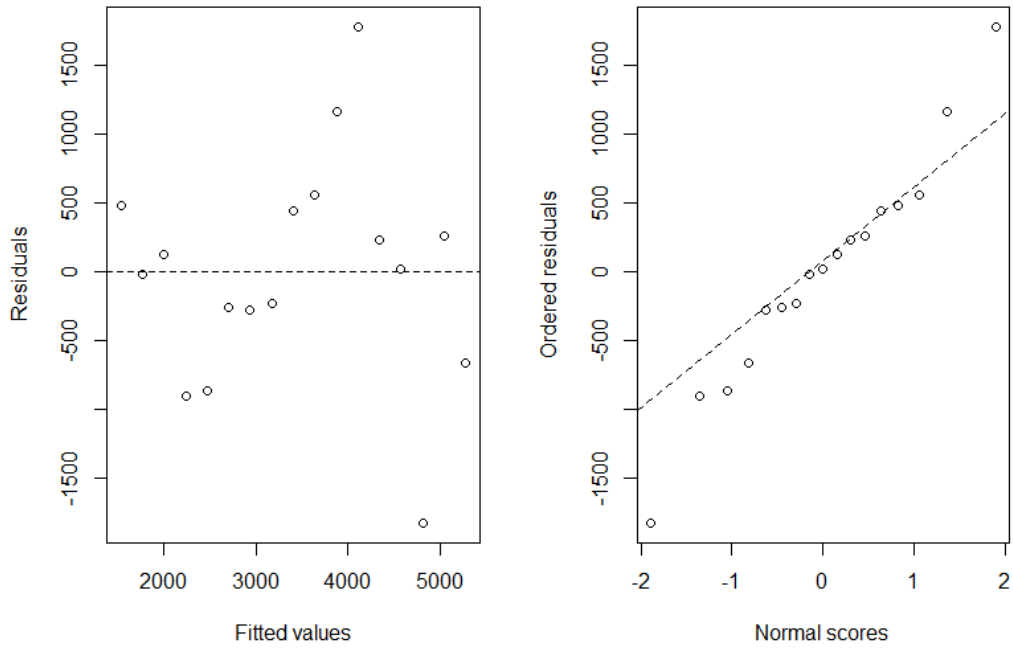
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Figure 2. Diagnostic analyses to assess regression model assumptions: panel A) residuals and panel B) autocorrelation (ACF: autocorrelation function).

An equivalence-testing approach was applied to the regression model to account for natural variability in the estimates. We chose biologically meaningful thresholds for the overall population trend as a 25% change in the population over a 25-year period (annual rate of change equal to -0.0119 and 0.0093 on the log-scale). A biologically meaningful trend occurs when the posterior probability distribution of the slope lies outside the equivalence region, whereas a negligible trend occurs when the slope is within the equivalence region. 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 (Camp *et al.* 2008).

Adaptive Winter Use Area
Although palila have been surveyed annually since 1980, beginning in 1998 additional transects were added within their most concentrated population area in an effort to improve knowledge of their range and precision of abundance estimates (Johnson *et al.* 2006). Historically palila were found in māmane-dominated forests of Mauna Kea, western Mauna Loa, and Hualālai volcanoes on the Island of Hawaiʻi, although more recently they have been confined to the subalpine, dry forest of Mauna Kea (Figure 1, inset). Palila are detected infrequently during winter surveys outside the west slope, therefore we limit inference on abundance to this region.

To examine the possibility that the palila's range may be contracting, we analyzed the spatial distribution of palila detections during the standardized winter surveys. For each year beginning in 2003, we identified the highest and lowest elevations on each transect where palila were recorded during the preceding five years (see Figure 1). We created a polygon around these endpoints with a 200 m buffer to determine the adaptive winter use area for each year during 2003–2014. These adaptive winter use area polygons were all within the 64.4 km^2 core survey area encompassing the standard polygon used to infer palila abundance and trend.

We used DISTANCE to estimate palila abundance within each annual winter use polygon during 2003–2014 using only those survey stations that fell within the polygon defined by the previous five years of surveys, using the same model chosen by the procedure to estimate the abundance described above. The density of palila in each year was extrapolated only to that year's polygon. This resulted in an annual estimate that excluded many stations, most—but not all—of which yielded zero palila detections. We used linear regression to determine if the size of the adaptive winter use area changed over time.

Repeat Surveys
Most stations within the winter use area were sampled twice between 2012 and 2014. 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, where second and third counts were pooled) 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 distributions of the first and subsequent counts differed.

<center>RESULTS</center>

Abundance
No palila were detected outside the winter use area in 2013 or 2014. In 2013, the palila population was estimated at 1,492–2,132 birds (point estimate: 1,799; Figure 3, Table 2). In
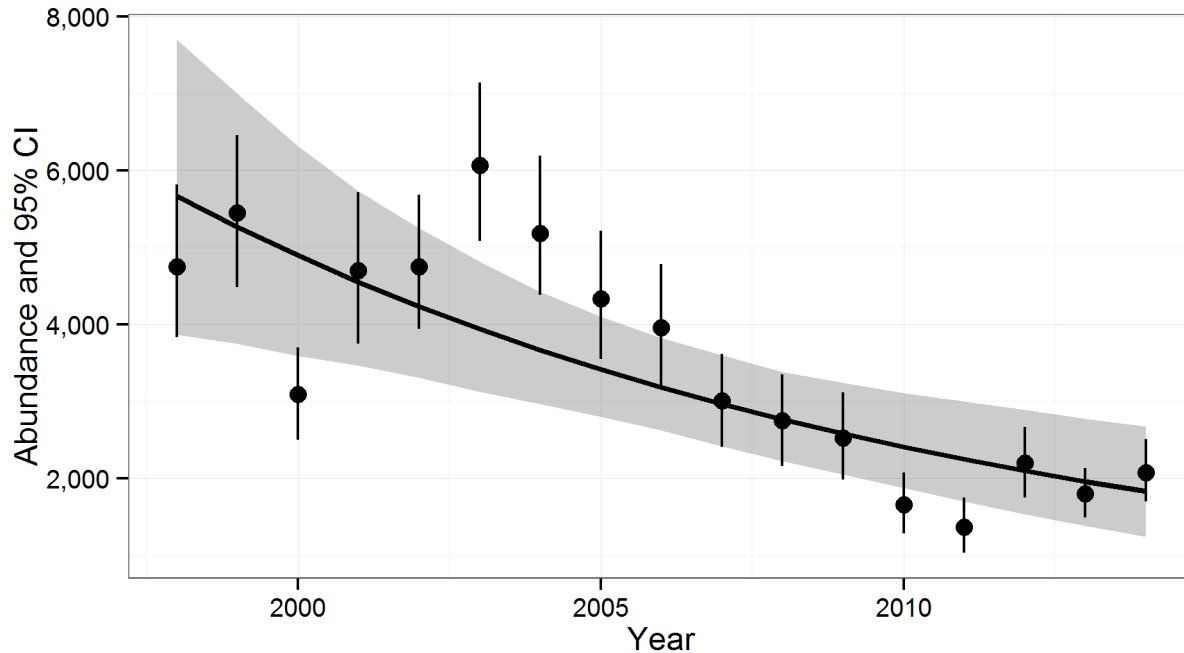
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Figure 3. Annual palila population estimates and 95% confidence intervals inside the winter use area on the western slope of Mauna Kea. Line represents the best fit log-linear regression (with lag-one autocorrelation) and associated 95% confidence band.

2014, the palila population was estimated at 1,697−2,508 (point estimate of 2,070). These estimates were for the 64.4 km² core survey area on the western flank of Mauna Kea and included 403 and 426 palila observations (337 and 355 on-count detections during 2013 and 2014, respectively). An additional 66 and 71 palila were detected outside of the six-minute count period and were used to fit the detection function but were not used to estimate population abundance. 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 [heard first but later visually confirmed, and seen only] detections) as a covariate (Appendix).

Trend
Between 1998 and 2003, palila numbers fluctuated moderately (coefficient of variation [CV] = 0.21 change from local minimum to maximum value; Figure 3). After peaking in 2003, palila population estimates declined steadily through 2011. During 2010−2014, estimates also fluctuated moderately above the 2011 level (CV = 0.18) with a local maximum peak in 2012. The average rate of decline during 1998−2014 has been 167 birds per year with very strong evidence (posterior probability; P = 0.998) of an overall declining trend in palila abundance. A lag-one ARMA model to account for temporal autocorrelation in the annual estimates yielded a steeply declining slope (-0.071, 95% credible interval = -0.102 to -0.034). This slope was well below the cautionary lower threshold level of a 25% change over a 25-year period (-0.0119), and there was virtually no support for a negligible (P = 0.002) or increasing population trend (P < 0.001). Over the 16-year monitoring period, the estimated rate of change equated to a

Table 2. Annual palila detections and population estimate parameters. Detections are given for palila recorded inside and outside the core survey area 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	# Detections outside	Estimate	%CV	SE	Lower limit	Upper limit
1998	313	2	4,746	10.46	496	3,832	5,811
1999	388	1	5,445	9.36	510	4,476	6,459
2000	234	14	3,089	10.16	314	2,499	3,698
2001	345	4	4,693	10.33	485	3,743	5,713
2002	339	9	4,750	9.25	439	3,936	5,677
2003	439	7	6,067	8.80	534	5,078	7,141
2004	371	9	5,182	8.90	461	4,379	6,184
2005	315	1	4,329	10.05	435	3,547	5,214
2006	271	16	3,958	10.21	404	3,176	4,778
2007	210	3	3,004	10.51	316	2,406	3,616
2008	186	na	2,744	11.12	305	2,158	3,346
2009	189	na	2,518	11.63	293	1,983	3,113
2010	151	na	1,654	12.37	205	1,276	2,069
2011	119	na	1,367	14.11	193	1,031	1,749
2012[1]	362	0	2,194	10.87	238	1,745	2,660
2013[2]	337	na	1,799	9.00	162	1,492	2,132
2014[3]	355	0	2,070	10.10	209	1,697	2,508

[1] Of the 362 detections made during the 2012 counts, 194 were observed on the first count and 168 during subsequent counts.
[2] Of the 337 detections made during the 2013 counts, 174 were observed on the first count and 163 during subsequent counts.
[3] Of the 355 detections made during the 2014 counts, 165 were observed on the first count and 190 during subsequent counts.

68% decline in the palila population. The regression model fit the observed data with a Bayesian R^2 of 0.69, and simulated model runs with log-normal random error were greater than observed errors 52.5% of the time, where a value of 50% would indicate the observed error fit a log-normal distribution exactly (Gelman *et al.* 2004).

Adaptive Winter Use Area
A linear regression fit to the adaptive winter use areas during 2003−2014 showed a non-significant ($P = 0.17$, $R^2 = 0.10$) decline of 0.3 km^2 per year (Figure 4, Table 3). We used the same detection functional form (hazard rate with detection type [visual detections pooled] as a covariate, truncated at 85.5 m; Appendix, Table 4) to estimate abundance in program DISTANCE, using only stations that fell into each annual use polygon and restricting inference to that same polygon. Population estimates of abundance during that 11-year period ranged from 1,185 to 5,182 (Table 3). Population estimates associated with the smaller winter use area polygons closely paralleled estimates derived from the larger non-adaptive winter use area of
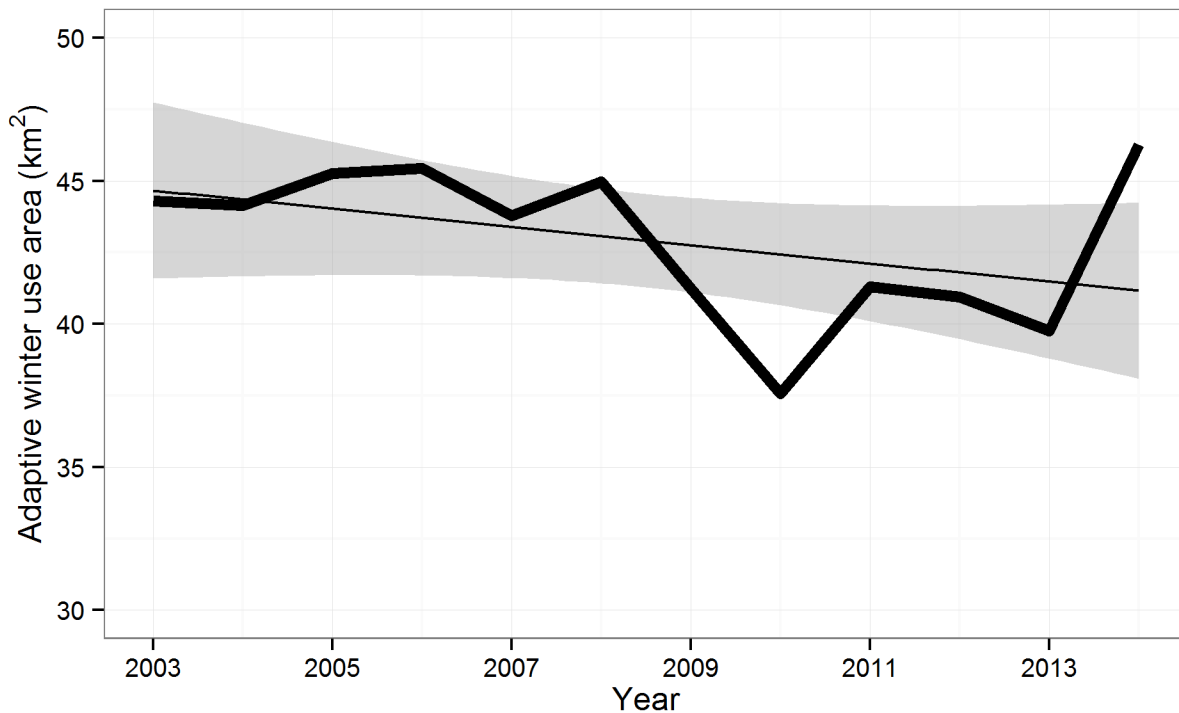
Figure 4. Size of winter use area polygons determined from annual palila surveys on Mauna Kea during 2003–2014 and associated 95% confidence band. Annual estimates of area (km^2) were based on the distribution of palila detections in the previous five years.

64.4 km^2 (R^2 = 0.997). Nevertheless, in all years the population estimate associated with the adaptive winter use area was lower than the estimate associated with the larger area. The coefficient of variation for the adaptive area estimates were slightly lower, but that difference was not statistically significant (Wilcoxon test, P = 0.18).

Repeat Surveys
The number of palila detected during first and subsequent counts varied, but the proportion of total detections on first and subsequent counts was similar; 54%, 52%, and 46% of all detections occurred on the first count in 2012, 2013, and 2014, respectively (Table 2). We used AIC statistics to assess the fit of detection probabilities from models with and without the covariate year. The model incorporating year as a covariate had an AICc value of 38643.12 (Appendix, Table 4). Nevertheless, including count with the year covariate inflated the AICc value by 151.20 units over the year-only model and thus did not improve the model fit. There was a small difference in the coefficients for the first and subsequent counts (0.0014 ± 0.1801 and 0.0046 ± 0.1849, respectively), but these differences were not statistically significant (two-sample z-test: z-score = -0.012, P = 0.99). Thus, there was no evidence that the numbers and distributions of birds detected in the first and subsequent counts were different.

Table 3. Palila population estimates for adaptive winter use area polygons. The size of the adaptive winter use area (Area; km^2) was based on the distribution of detections during annual surveys from the previous five years. Survey transects were added in 1998 (see Figure 1), so adaptive winter use areas begin in 2003, the first year with five years of preceding surveys on all transects. Abundance and % coefficient of variation ("core area" columns) for the 64.4 km^2 core survey area are reproduced from Table 2 for comparison.

| Year | Adaptive area | | | | | Core area | |
	Area	Abundance	%CV	CI	Density	Abundance	%CV
2003	44.3	5,182	8.95	4,304–6,121	117.0	6,067	8.80
2004	44.1	4,318	8.10	3,693–5,064	97.9	5,182	8.90
2005	45.3	3,530	9.04	2,937–4,188	77.9	4,329	10.05
2006	45.4	3,250	9.50	2,679–3,889	71.6	3,958	10.21
2007	43.8	2,523	9.97	2,036–3,022	57.6	3,004	10.51
2008	45.0	2,234	10.77	1,799–2,742	49.6	2,744	11.12
2009	41.2	2,158	11.49	1,717–2,689	52.4	2,518	11.63
2010	37.6	1,290	11.92	987–1,590	34.3	1,654	12.37
2011	41.3	1,185	13.88	883–1,528	28.7	1,367	14.11
2012	40.9	1,656	8.66	1,389–1,951	40.5	2,194	10.87
2013	39.8	1,261	7.93	1,074–1,466	31.7	1,799	9.00
2014	46.3	1,549	7.80	1,328–1,802	33.5	2,070	10.10

CONCLUSIONS

There is very strong evidence that the palila population has declined since 1998. No palila were detected on surveys outside the core survey area on the western slope of Mauna Kea in 2013 or 2014. Although there has been a general downward trend in the adaptive estimates of the palila range since 2003, this contraction was not significant, suggesting that the range of the species has remained stable since 2003, albeit at only about 5% of its historical extent (Banko *et al.* 2013). Abundance estimates restricting each year's area of inference to a polygon based on the distribution of palila during the previous five years of surveys were consistently lower than estimates based on the entire 64.4 km^2 core survey area. This difference implies that estimates based on the core survey area may be inflated, though by how much depends upon the delineation of the species' range. As an example, in 2014 the abundance point estimate based on samples in the 46.3 km^2 area (defined by detections in the 2009–2013 surveys) was only 72% of the estimate using the entire survey area. The area occupied, however, changes from year-to-year. Thus, palila density and range are confounded when estimating population status and trends. To control for this problem, trends are assessed in the core survey area, which holds the area constant and produces an accurate population rate of change. In contrast, estimating the population size in the adaptive winter use area produces a more accurate estimate of the annual abundance. Any change in the survey protocol would benefit from consultation between resource managers and wildlife biologists to decide upon the relevant area of inference to estimate population abundance and change in the species' range.

Delineating the extent of the palila geographical range during winter surveys, and monitoring how that area varies through time is a quantitative method to assess changes in the species' range. Both spatial and temporal dynamics complicate characterizing the species' range. We delimited the outer species' border based on the palila distribution during the previous five years; simple presence and absence data. The 2010 winter use area was 37.6 km^2, smaller than in all other years because no palila were detected on the southernmost transect during the previous five years. This example illustrates the importance of sampling beyond the adaptive winter use area. Based on the 2010 delimited range palila could have been considered locally extinct at the southern edge of the survey area based on a lack of occurrence between 2005 and 2009. However, in 2011 and subsequent surveys palila were again detected on the southern transect. Further, the 2010 anomaly highlights the discrete nature of adaptive winter use areas caused by the greater distances between transects (400–2,700 m at transect midpoints) than between survey stations along transect (150 m).

We sought to minimize the spatial-temporal fluctuations by delimiting the adaptive winter use area using occurrences from the previous five years. Five years is an arbitrary interval, but it is long enough to average out some temporal stochasticity, approaches the species' mean lifespan, and still provides several estimates from which to assess changes in the species' range in the relatively short 11-year time series. With continued annual surveys of the core survey area and periodic surveys of the transects on the other slopes of Mauna Kea, more advanced and nuanced techniques to quantify the spatial patterns in palila geographic range and border can be applied (Fortin *et al.* 2005).

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 detection tends 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. However, similar numbers of palila were detected during the first and subsequent counts during 2012−2014. In addition, the detection model was not improved by including time of detection or count covariates, indicating that there was no difference in the numbers of palila detected or in their detection probability due to time of day or count sequence. Thus, greater precision in population estimates can be gained if future surveys include repeat visits.

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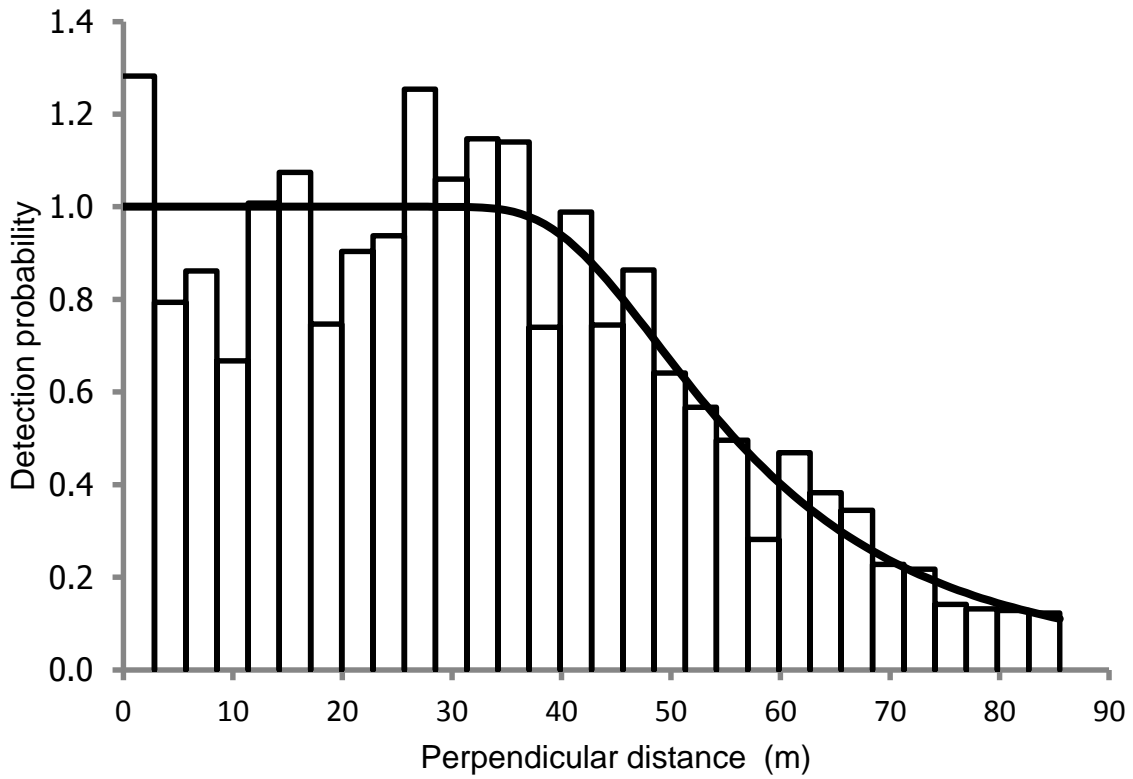
Figure 5. Hazard-rate detection function, with no expansion series and including the covariate detection type (heard versus pooled visual; line) and palila distance data (histogram) pooled across all surveys from 1998 to 2014. Data were truncated at 85.5 m.

Appendix, continued

Table 4. Results of fitting 17 detection function models to the 1998–2014 palila distance histogram. ΔAICc is the difference in AICc scores between each model and the overall best-fit model, and w_i is the discrete model probability.

Model[1,2]	# parameters	Ln(-likelihood)	AICc	ΔAICc	w_i
H-rate Key VBDetType	3	19302.85	38611.70	0	0.73595
H-rate Key DetType	4	19302.87	38613.75	2.05	0.26405
H-rate Key Year(f)	18	19303.49	38643.12	31.42	0.00000
H-rate Key Obs	25	19370.16	38790.61	178.91	0.00000
H-rate Key HBDetType	3	19418.90	38843.80	232.10	0.00000
H-rate Key Gust	5	19441.62	38893.25	281.55	0.00000
H-rate Key Wind	4	19444.84	38897.69	285.99	0.00000
H-rate Key	2	19453.55	38911.11	299.41	0.00000
H-rate Key Rain	3	19452.81	38911.63	299.93	0.00000
H-rate Key Time	3	19453.15	38912.30	300.60	0.00000
H-rate Key Year(c)	3	19453.16	38912.33	300.63	0.00000
H-rate Key Cloud	4	19453.01	38914.02	302.32	0.00000
H-norm Key	1	19519.62	39041.23	429.53	0.00000

[1] 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, 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).

[2] Models H-norm Cos, H-norm H-poly, H-rate Cos, and H-rate S-poly failed to converge.