

CONFORMITY AND INCONGRUITY OF SELECTED DISPERSION  
INDICES IN DESCRIBING THE SPATIAL DISTRIBUTION OF  
*TRICHOPLUSIA NI* (HÜBNER) IN GEOGRAPHICALLY  
SEPARATE CABBAGE PLANTINGS

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INTRODUCTION

Information on the distribution of insects within and between plants not only provides the basis for applied uses such as the development of sampling programs (IWAO, 1975; TAYLOR, 1984) and selection of optimal plant characteristics for use in breeding for resistance (LIN et al., 1987), but allows for a more fundamental understanding of insect population and community ecology. Dispersion data can provide insight into the behavior of some species (TRUMBLE and OATMAN, 1984) as well as nutritional and micro-habitat preferences (KENNEDY et al., 1950).

Unfortunately, not all methods of determining population distributions will generate equivalent dispersions from the same data set. Theoretical models of dispersion, such as Neyman type A, Poisson, and the negative binomial, may all be fitted to identical count data depending on the frequency classes chosen or the method of fitting the model (IWAO, 1970; BOSWELL and PATIL, 1970). MYERS (1978), TRUMBLE et al. (1983), and MOLLET et al. (1985) demonstrated that even alternative indices which are not based on frequency classes are subject to statistical artifacts which, if not recognized, will preclude accurate interpretation. Clearly, additional studies designed to compare dispersion indices are needed to provide the breadth of experience necessary to allow accurate interpretation of these indices over a broad range of environmental and statistical conditions.

The study reported here had three specific objectives. Our first objective was to determine if the distribution of *Trichoplusia ni* (HÜBNER) (Lepidoptera: Noctuidae) in cabbage (*Brassica oleracea* L.) varied with larval growth stage or between widely separated locations. Information of this nature provides insight into the feasibility of developing regional sampling programs or economic threshold levels using locally generated dispersion data for *T. ni*. Second, we examined potential interactions of selected indices with mean density in order to document if any of the indices varied linearly with density.

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A significant relationship between density and an index would indicate that, over the range of densities we observed, a statistical anomaly would invalidate normal interpretation of the index values. Third, we wished to investigate the statistical relationships between GREEN's coefficient (GREEN, 1966), TAYLOR's power law (TAYLOR, 1965), and both MORISITA's original (MORISITA, 1959) and Standardized indices (SMITH-GILL, 1975) using field data.

#### MATERIALS AND METHODS

The abundance and distribution of *T. ni* on cabbage was evaluated during 1984 and 1985 in widely separated regions in the southern USA. The 1984 and 1985 plantings at Weslaco, Texas (Gulf Coast) and Orange County, California (Pacific Coast) were .405 hectares in size, while the 1984 planting in Baton Rouge, Louisiana was .202 hectares. At all locations, 'Greenback' variety cabbage was planted at 25 cm intervals into double-row beds spaced 1.0 m apart. In California, cabbage was transplanted in 1984 and 1985 on 4 and 15 March, respectively. Cabbage was transplanted in Louisiana on 5 March, 1984. In Texas, cabbage was direct-seeded in 1984 and 1985 on 23 August and 2 February, respectively. All plantings were fertilized with two applications of 33.6 kg/ha of  $\text{NO}_3$  during the growing season. No insecticides were used on the plots.

In all plantings, weekly sampling was initiated when plants reached the 5–10 true-leaf stage. Since each cabbage field was a "continuum", and *T. ni* larvae do not readily migrate between plants, individual plants corresponded to the "natural habitat unit" desirable for sampling described by PATIL and STITELER (1974). Each planting was stratified into six equal blocks. From the initiation of sampling until the onset of heading (production of first wrapper leaves—terminology after THEUNISSEN and SINS, 1984), ten randomly selected plants per block per week were non-destructively monitored for the presence of small *T. ni* larvae (instars 1–2) and large larvae (instar 3 or larger). Counts of total larvae per plant were generated by adding these values. From production of wrapper leaves to harvest, five plants per block were destructively sampled per week.

Data were analyzed and dispersion indices generated using the Statistical Analysis System (available from the SAS Institute, Cary, NC). GREEN's coefficient  $C_x$  (GREEN, 1966) was calculated weekly for each larval category (small, large, and total) using the formula:

$$C_x = (s^2/m - 1) / (\Sigma n - 1),$$

where  $s^2$  = variance,  $m$  = mean number of *T. ni* larvae per plant, and  $\Sigma n$  = total number of larvae observed per week. Population dispersions are considered uniform at values near  $-1.0$ , random at  $0.0$ , and increasingly aggregated as  $+1.0$  is approached. MORISITA's original Index  $I_\delta$  (MORISITA, 1959) also was calculated weekly for each larval category using the formula:

$$I_\delta = \frac{q \sum_{i=1}^q [n_i(n_i - 1)]}{N(N - 1)}$$

where  $q$  = number of plants sampled (plant in our study = quadrat),  $n_i$  = the number of larvae in the  $i$ th quadrat (which ranges from 1 to  $q$ ), and  $N = \sum n_i$  = the total number of larvae sampled in each category. For  $I_s$ , values increase with the degree of aggregation in the population (range = 0 to  $q$ ). MORISITA (1959) proposed an  $F$ -test to determine if  $I_s$  values significantly depart from randomness, but the validity of this test has been questioned (STITELER and PATIL, 1970). By stratifying our fields into "subareas" (terminology after MORISITA 1959), and sampling small quadrats (plants) that fall entirely within the subareas (replicates in our study), we met the underlying assumptions required by this index.

MORISITA's standardized index  $I_b$ , a transformation of  $I_s$  which allows ready comparisons of samples with varying densities (and therefore unequal confidence limits), was calculated using the variable formulas listed by SMITH-GILL (1975). Like  $C_x$ ,  $I_b$  ranges from  $-1.0$  to  $+1.0$  with the same interpretation of the values. However, the significance tests for departure from randomness for  $I_b$  are incorporated into the index values, such that values above  $+0.5$  indicate significant aggregation and values below  $-0.5$  denote uniformity.

The parameters  $a$  and  $b$  of TAYLOR's power law (TPL) (TAYLOR, 1965),  $s^2 = am^b$ , were estimated from the regression equation:

$$\log(s^2) = \log(a) + (b)\log(m),$$

where  $a$  and  $b$  = the intercept and slope coefficients, respectively. Student's  $t$ -tests with 95% confidence limits were used to test the hypotheses that the "coefficient of aggregation" ( $b$ ) = 1, and if  $\log(a) = 0$ . According to Taylor, if  $b$  is significantly greater than unity, then the population is aggregated, if  $b$  is not significantly different from unity, then the population is considered to have a random dispersion. TPL values were calculated for each larval category (small and large) by location and year.

GREEN's (1970) formula was used to determine the sample sizes and related mean densities needed to estimate population density at a constant precision level:

$$\log T_q = \frac{\log(D_0^2/a)}{b-2} + \frac{b-1}{b-2} \log q$$

where  $a$  and  $b$  are from TPL,  $T_n$  = cumulative number of  $T. ni$  larvae,  $q$  = sample size (number of plants sampled) and  $D_0$  = the fixed level of precision. For the purposes of this study, we set  $D_0 = 25\%$ , a value considered sufficiently accurate for use in pest management programs (SOUTHWOOD, 1978).

Linear regression was used to determine if index values varied with mean density or each other. Relationships for each index except TPL were analyzed using weekly values for each larval category within each location and year. Plots of residuals from these regressions produced no recognizable patterns, indicating that more complex regression techniques were not required. For  $I_s$ , interactions of quadrat numbers and larval density at low population levels resulted in some statistical anomalies. These values, which were 5.00 or 10.00, were not included in the analyses. Because  $I_s$  values are utilized in  $I_b$  calculations, the corresponding  $I_b$  values also were omitted from the

analyses. The occurrence of similar artifacts at low population densities has been noted previously for TPL (TAYLOR, 1984).

#### RESULTS AND DISCUSSION

Although the mean density of *T. ni* per plant varied between larval category and year within each geographic location, two general trends were evident. First, *T. ni* populations were only slightly aggregated. Maximum and minimum values for  $C_x$  for any larval category were 0.086 and  $-0.013$ , respectively (Table 1), which suggested only minor variation from randomness. Similarly, values of  $I_s$  and  $I_b$  calculated for each week of sampling were variably significant and non-significant, with a preponderance of significant values suggesting slight aggregation (Tables 2-4). Slopes from TPL regressions produced coefficients of aggregation ranging from significantly aggregated to nearly random. However, the maximum slope of 1.188 generated in this study (small larvae in California, 1984) is not indicative of a highly aggregated population (Table 5). These results are similar to data reported by HARCOURT (1965) for theoretical distributions describing dispersion of *T. ni* larvae on crucifer plantings (primarily cauliflower) in New York. HARCOURT's data generally fit the negative binomial distribution, but some data sets (30%) adequately fit the Poisson distribution. In contrast, SHEPARD and CARNER (1976) found only 28% of the data sets for *T. ni* distribution on soybeans in South Carolina fit the negative binomial, while 53% fit the Poisson distribution.

A general reduction in aggregation for larger size class larvae as compared to smaller

Table 1. Maximum and minimum values of GREEN's coefficient, and density of *Trichoplusia ni* larvae per plant for three locations in the southern United States.

Year	Location	Number of weekly samples <sup>1</sup>	Larval Category <sup>2</sup>	Range (minimum, maximum)	
				Mean larvae/plant	GREEN's Coefficient
1984	California	12	small	0.00, 1.77	-0.002, 0.086
			large	0.00, 2.92	-0.002, 0.008
			total	0.00, 5.06	-0.002, 0.072
	Louisiana	10	small	0.00, 1.50	-0.002, 0.045
			large	0.00, 0.70	-0.002, 0.006
			total	0.08, 2.27	-0.002, 0.064
	Texas	11	small	0.07, 3.07	-0.006, 0.018
			large	0.07, 0.83	-0.013, 0.006
			total	0.23, 3.33	-0.013, 0.025
1985	California	9	small	0.02, 4.12	0.002, 0.016
			large	0.00, 4.07	-0.001, 0.003
			total	0.02, 7.76	0.001, 0.016
	Texas	8	small	0.00, 0.80	-0.004, 0.030
			large	0.00, 0.48	-0.003, 0.034
			total	0.03, 1.08	-0.004, 0.021

<sup>1</sup> larval counts were made at weekly intervals from 5-10 true leaf stage to harvest.

<sup>2</sup> small larvae=first and second instars, large larvae=third or larger instars.

Table 2. Mean larvae per plant, and original ( $I_s$ ) and standardized values ( $I_p$ ) of MORISITA'S indices for small *Trichoplusia ni* larvae in three locations in the southern United States<sup>1</sup>.

Year	Week of sampling	California			Louisiana			Texas		
		Mean	$I_s$	$I_p$	Mean	$I_s$	$I_p$	Mean	$I_s$	$I_p$
1984	1	0.00	—	—	0.00	—	—	0.07	0.000	-0.081
	2	0.02	—	—	0.00	—	—	0.65	0.972	-0.029
	3	0.03	*	*	0.00	—	—	3.07	1.636	0.504
	4	0.09	*	*	0.00	—	—	1.20	1.814	0.504
	5	0.57	5.989	0.536	0.08	*	*	0.42	0.369	-0.426
	6	0.12	0.000	-0.162	0.45	2.148	0.502	0.83	1.603	0.500
	7	0.13	8.429	0.536	1.20	3.084	0.528	2.46	1.809	0.500
	8	0.25	1.978	0.262	2.27	1.544	0.504	2.75	1.747	0.505
	9	1.33	2.297	0.508	1.73	2.756	0.525	1.45	2.299	0.505
	10	1.78	2.685	0.512	0.17	2.000	0.139	0.53	0.097	0.508
	11	1.64	1.959	0.506	—	—	—	—	—	—
	12	2.16	2.263	0.509	—	—	—	—	—	—
1985	1	0.02	*	*	—	—	—	—	—	—
	2	0.38	2.338	0.502	—	—	—	0.15	*	*
	3	1.20	1.033	0.048	—	—	—	0.21	1.564	0.138
	4	2.90	1.196	0.500	—	—	—	0.28	4.067	0.512
	5	3.60	1.239	0.501	—	—	—	0.57	1.604	0.410
	6	4.12	1.159	0.500	—	—	—	0.80	2.270	0.510
	7	3.20	1.364	0.502	—	—	—	0.35	2.095	0.503
	8	3.39	1.334	0.502	—	—	—	0.60	2.095	0.503
	9	3.70	1.304	0.502	—	—	—	—	—	—

<sup>1</sup> larval counts were made at weekly intervals from 5-10 true-leaf stage to harvest; a '\*' indicates statistical anomaly occurred due to interaction of mean and sample size.

larvae was the second trend observed. This trend was apparent for  $I_p$  and  $I_s$  for all dates and locations, and for  $C_x$  and TPL for all locations in 1984 and California in 1985. The 1985 data set from Texas produced the only exception. According to HARCOURT (1965), who also reported a decrease in aggregation with an increase in larval size, this trend was due to an increase in mortality associated with an epizootic caused by a nuclear polyhedrosis virus. Unfortunately, this explanation is not adequate for our results, as few diseased larvae were observed. Possible explanations for the decreased aggregation observed for large larvae in our study include substantial larval mortality due to predators and parasites (MARTIN et al., 1984) and/or behavioral changes (searching for pupation sites, etc.) resulting in interplant movement.

The question then arises as to whether these larval categories with variable dispersions can be combined to produce a 'total' category without causing a misleading biological interpretation. In some instances, the generally larger populations of small larvae with greater degrees of aggregation biased the total category toward higher levels of aggregation (see Table 5, 1984). However, the analyses for total larvae could not be simply additive because equal numbers of small and large larvae did not occur on each plant, the potential existed for extreme effects, where a data set having a few large larvae and greater numbers

Table 3. Mean larvae per plant, and original ( $I_s$ ) and standardized values ( $I_p$ ) of MORISITA'S indices for large *Trichoplusia ni* larvae in three locations in the southern United States<sup>1</sup>.

Year	Week of sampling	California			Louisiana			Texas		
		Mean	$I_s$	$I_p$	Mean	$I_s$	$I_p$	Mean	$I_s$	$I_p$
1984	1	0.00	—	—	0.00	—	—	0.16	0.000	-0.243
	2	0.00	—	—	0.00	—	—	0.07	*	*
	3	0.00	—	—	0.00	—	—	0.26	2.000	0.309
	4	0.03	*	*	0.00	—	—	0.70	1.488	0.409
	5	0.05	0.000	-0.054	0.00	—	—	0.33	0.702	-0.145
	6	0.07	0.000	-0.081	0.22	1.487	0.122	0.33	0.947	0.500
	7	0.24	3.371	0.506	0.37	5.800	0.558	0.32	0.702	-0.027
	8	0.25	1.978	0.262	0.77	2.253	0.508	0.37	1.558	-0.145
	9	0.27	0.000	-0.401	0.70	2.762	0.516	0.83	1.295	0.241
	10	0.90	1.551	0.501	0.07	*	*	0.30	0.980	0.295
	11	2.93	1.304	0.501	—	—	—	—	—	—
	12	2.90	1.391	0.502	—	—	—	—	—	—
1985	1	0.00	—	—	—	—	—	0.03	*	*
	2	0.00	—	—	—	—	—	0.00	—	—
	3	0.23	3.956	0.509	—	—	—	0.05	0.000	-0.053
	4	0.69	1.462	0.399	—	—	—	0.09	0.000	-0.134
	5	2.82	1.492	0.503	—	—	—	0.10	4.000	0.309
	6	3.87	1.413	0.503	—	—	—	0.10	3.393	0.304
	7	4.07	1.275	0.501	—	—	—	0.23	3.077	0.500
	8	3.37	1.909	0.507	—	—	—	0.48	3.695	0.516
	9	2.97	1.175	0.500	—	—	—	—	—	—

<sup>1</sup> larval counts were made at weekly intervals from 5–10 true leaf stage to harvest; a "\*" indicates statistical anomaly occurred due to interaction of mean and sample size.

of small larvae produced a level of aggregation for the total which was greater than that calculated for either the small or large categories (see Table 5, 1985, Texas).

Although the interpretational problems resulting from combining the small and large larval categories would probably justify not using this approach, the biological implications of the combination may be even more questionable. Combining data from larval groups which not only exhibit variable dispersions, but have differences in behavioral patterns or mortality factors and rates, would be unlikely to produce a valid estimate of a characteristic aggregation for a species.

Analyses of geographic variation in distribution produced few easily interpreted patterns. As measured by TPL, populations of *T. ni* appeared to be consistently aggregated in California, often aggregated in Texas, and never aggregated in Louisiana (Table 5). Although this trend also was evident for the data sets from California and Texas analyzed for  $I_s$  and  $I_p$ , the data from Louisiana were contrary but too limited to provide conclusive evidence (Tables 2–4).

The potential impact of geographic discrepancies in dispersion on sampling programs for small and large size class larvae has been presented in Fig. 1. Based on GREEN'S (1970) formula for determining the relationship between sample size and the number of

Table 4. Mean larvae per plant, and original ( $I_\delta$ ) and standardized values ( $I_\rho$ ) of MORISITA'S indices for total *Trichoplusia ni* larvae in three locations in the southern United States<sup>1</sup>.

Year	Week of sampling	California			Louisiana			Texas		
		Mean	$I_\delta$	$I_\rho$	Mean	$I_\delta$	$I_\rho$	Mean	$I_\delta$	$I_\rho$
1984	1	0.00	—	—	0.00	—	—	0.23	1.319	-0.017
	2	0.02	*	*	0.00	—	—	0.72	0.930	0.294
	3	0.03	*	*	0.00	—	—	3.33	1.544	0.419
	4	0.12	0.000	-0.162	0.00	—	—	1.90	1.572	0.511
	5	0.62	5.315	0.531	0.08	*	*	0.75	0.788	0.502
	6	0.19	0.000	-0.270	0.23	3.187	0.503	1.18	1.460	0.455
	7	0.37	3.032	0.508	0.83	3.190	0.527	2.78	1.757	0.513
	8	0.50	1.270	0.150	1.50	1.848	0.508	3.12	1.677	0.515
	9	1.60	1.779	0.504	1.03	3.181	0.529	2.28	1.637	0.503
	10	2.68	1.948	0.507	0.10	0.000	-0.082	0.83	0.980	-0.034
	11	4.57	1.385	0.503	—	—	—	—	—	—
	12	5.06	1.541	0.504	—	—	—	—	—	—
1985	1	0.02	*	*	—	—	—	0.03	*	*
	2	0.38	2.338	0.502	—	—	—	0.15	*	*
	3	1.43	1.033	0.048	—	—	—	0.25	1.162	0.046
	4	3.39	1.196	0.500	—	—	—	0.37	3.961	0.515
	5	6.42	1.239	0.501	—	—	—	0.67	1.385	0.309
	6	7.76	1.159	0.500	—	—	—	0.90	2.556	0.510
	7	7.50	1.364	0.502	—	—	—	0.58	1.513	0.359
	8	6.68	1.334	0.502	—	—	—	1.08	1.875	0.504
	9	6.67	1.304	0.502	—	—	—	—	—	—

<sup>1</sup> larval counts (all stages) were made at weekly intervals from 5-10 true leaf stage to harvest; a "\*" indicates statistical anomaly occurred due to interaction of mean and sample size.

larvae required for an accurate estimation of the population, a need for increasing numbers of larvae per plant followed the pattern: Louisiana < Texas < California. While this pattern suggests a trend of increasing aggregation across regions, the number of regions evaluated was too small to allow definitive conclusions. However, these data clearly show that developing a sampling program for two or more geographically distinct regions based on a single dispersion value for *T. ni* larvae would not always be valid. In addition to geographic relationships, Fig. 1 demonstrates the effect of the more aggregated dispersions of small larvae, which require two or more times as many larvae per plant as large larvae for an accurate estimate of mean density.

The calculation of dispersion indices for multiple years, locations and larval categories provided a series of data sets for naturally occurring populations which permitted further analysis of the potential for statistical anomalies within indices, as well as an analysis of the degree of relatedness between selected indices. A continuing concern expressed in the literature has been the potential presence of a relationship between some indices and mean density (SKELLAM, 1952; MYERS, 1978; MOLLET et al., 1984). In order to determine if  $C_x$ ,  $I_\delta$  or  $I_\rho$  varied detectibly with the population size, these values were regressed on mean densities of small and large larvae occurring weekly within each location and

Table 5. Dispersions of *T. ni* larvae infesting cabbage in California, Texas and Louisiana during 1984 and 1985 as measured by TAYLOR's Power Law,  $s^2=am^b$ .

Year	Larval Category <sup>1</sup>	Location	Intercept <sup>2</sup>	Slope <sup>3</sup>	Coefficient of Determination
1984	small	California	0.881*	1.188*	0.926
		Louisiana	0.456	1.097	0.882
		Texas	0.609*	1.185*	0.937
	large	California	0.318*	1.073*	0.993
		Louisiana	0.188	1.045	0.977
		Texas	0.134	1.035	0.964
	total	California	0.557*	1.124*	0.941
		Louisiana	0.319	1.061	0.886
		Texas	0.504	1.160*	0.913
1985	small	California	0.634*	1.160*	0.945
		Texas	0.223	1.050	0.852
	large	California	0.360*	1.102*	0.973
		Texas	0.334	1.079	0.929
	total	California	0.439*	1.100*	0.969
		Texas	0.349	1.093	0.914

<sup>1</sup> small larvae=first and second instars, large larvae=third or larger instars, total=all larvae.

<sup>2</sup> intercepts followed by an '\*' are significantly different from zero at  $P<0.05$ , *t*-test.

<sup>3</sup> slopes followed by an '\*' are significantly different from unity (1.00) at  $P<0.05$ , *t*-test.

year.  $C_x$  was found to be independent of mean density in nine of ten data sets. Although  $C_x$  values for small larvae were significantly and negatively related to mean density in California during 1985 ( $P=0.05$ ,  $\beta=-195$ ,  $r=0.704$ ,  $SE=1.024$ ), this result may not be reliable as regressions following removal of a single outlying data point generated a non-significant relationship.

Results of such analyses for  $I_s$  and  $I_b$  suggested that neither of these indices was closely dependent on the mean. In all ten cases,  $I_s$  did not vary significantly with mean density. These results are in agreement with those of MYERS (1978), whose computer-generated counts documented a lack of significant relationship between  $I_s$  and density when populations approached randomness. In one of ten data sets,  $I_b$  showed a significant relationship to density (small larvae, Texas, 1984;  $P=0.04$ ,  $\beta=2.074$ ,  $r=0.68$ , standard error=0.857). However, the general lack of a correlation between  $I_b$  and density concurs with previous reports by MOLLET et al. (1984) for field-collected data of *Tetranychus urticae* (KOCH) distributions in cotton.

The degree of relatedness between  $I_s$  or  $I_b$  and  $C_x$  was evaluated by linear regression. Few significant relationships were found when either  $I_s$  or  $I_b$  were regressed on  $C_x$  ( $I_s=3$  of ten data sets,  $I_b=1$  of ten data sets). Although an increase in  $C_x$  intuitively was postulated to concur with an increase in  $I_s$ , the apparent lack of agreement could be explained by variability occurring in response to differences in ranges of the indices ( $C_x=-1$  to  $+1$ ;  $I_s=$ zero to  $q$ ). However, this simple explanation clearly was



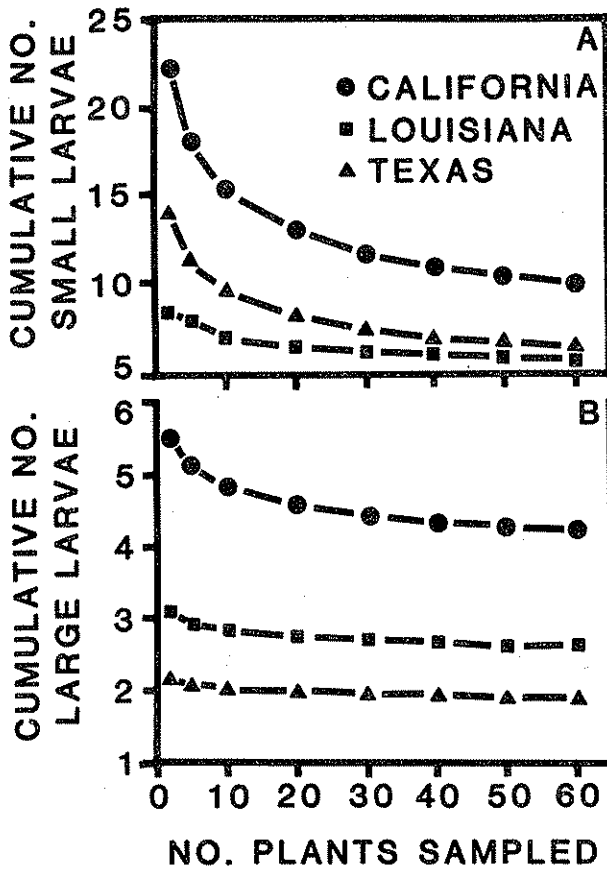


Fig. 1. Stopline for constant precision level sequential samples for a) small *T. ni* larvae (instars 1-2), and b) large larvae (instars 3+) based on a precision level of 0.25.

inadequate because 1) two out of three significant regressions between  $I_s$  and  $C_x$  (from ten data sets) could be explained as resulting from a single outlying point, 2)  $I_s$  has a range identical to that of  $C_x$ , and 3) only one data set of ten generated a significant correlation between  $I_s$  and  $C_x$ .

A more complete explanation is based upon the inadequacy of linear regression as used above as a measure of the relationship between  $I_s$  and  $C_x$ . Algebraically, these indices are related as follows:

$$I_s = \frac{q(\bar{n}-1)}{\bar{n}} C_x + \frac{q\bar{n}(\bar{n}-1)}{\bar{n}(q\bar{n}-1)}, \text{ or alternatively, } I_s = k_1(\bar{n})C_x + k_2(\bar{n}) \text{ where, } \bar{n} = \frac{1}{q} \sum_{i=q}^q m_i, \text{ and}$$

all other notation is as previously described.

The presence of this relationship implies that the lack of significant linear regression results from the variability in mean density among points used in the regression, not because of any lack of a linear relationship between  $I_s$  and  $C_x$ . In effect, nonlinearity occurs because the  $k_1$  and  $k_2$  change with density estimates, and such density estimates varied over time in our study.

We suspect that the lack of apparent agreement between these indices may also be a function of the insensitivity of  $C_x$  at the levels of dispersion found for *T. ni* larvae in our study. The maximum range of  $C_x$  within any year, location and larval category was  $-0.002$  to  $0.086$  (small larvae, 1984, California). This represents only 4.4% of the available definition of the index. In contrast, for the same year, location and larval category,  $I_p$  exhibited a range of  $-0.162$  to  $0.536$ , or 34.9% of the available definition. Thus, each value of  $C_x$  in a regression was related to a broad and variable range of  $I_p$  values, resulting in a non-significant relationship between the two methods. Additional research into the statistical and theoretical bases for this apparent loss of sensitivity of  $C_x$  near randomness would be valuable.

The lack of definition provided by  $C_x$  at the dispersion levels in our study has implications for interpretation of the biological import of this index. The data for large larvae in California during 1985 provide an example; even though TPL indicates the larvae were significantly aggregated (Table 5), and in six of seven weeks  $I_s$  and  $I_p$  documented aggregated dispersions (Tables 2 and 3), the maximum value achieved by  $C_x$  was 0.003. While this value may represent a significantly aggregated distribution, the lack of definition for weakly aggregated populations makes interpretation difficult and suggests that  $C_x$  may not be as useful as  $I_p$  for comparisons of data sets collected between species or locations.

#### SUMMARY

The study reported here had three specific objectives. Our first objective was to determine if the distribution of *Trichoplusia ni* in cabbage varied with larval growth stage or between widely separated locations. In general,  $C_x$ ,  $I_s$ ,  $I_p$  and TAYLOR'S power law were in agreement that *T. ni* larval populations were only slightly aggregated. However, small larvae (1-2 instar) were substantially more aggregated than larger larvae, creating the need for separate sampling programs for each category. Combining data from larval groups which not only exhibited variable dispersions, but had potential differences in behavioral patterns or mortality factors and rates, would be unlikely to produce a valid estimate for a characteristic aggregation of a species. Although a trend of decreasing aggregation from California>Texas>Louisiana was evident in 1984, additional regions need to be assessed before definitive conclusions can be made regarding population resource utilization patterns. However, these data clearly show that developing a sampling plan for two or more geographically distinct regions based on a single dispersion value for *T. ni* would not always be valid.

Our second objective was to examine potential interactions of selected indices with mean density in order to document if any of the indices varied linearly with density. In correlation analyses using paired weekly values in ten data sets,  $C_x$  and  $I_s$  produced no significant relationships which could not be explained by statistical artifacts. A single significant relationship was documented for  $I_p$ . Thus, these indices appeared relatively

free of dependence on the mean.

The third objective was to investigate the statistical relationships between  $C_x$ , and both  $I_s$  and  $I_p$ . Although an increase in  $C_x$  was postulated to concur with an increase in  $I_s$  or  $I_p$ , few significant relationships were found when  $I_s$  or  $I_p$  were regressed on  $C_x$ . For those which were significant, the slopes of the regression lines were nearly flat (maximum  $\beta=0.015$ ). We suspect that the lack of apparent agreement between these indices may be a function of the insensitivity of  $C_x$  at the levels of dispersion found for *T. ni* larvae in our study. The lack of definition provided by  $C_x$  at slightly aggregated dispersions makes interpretation of the biological import of this index difficult and suggests that  $C_x$  may not be as useful as  $I_p$  for comparisons of data sets collected between species or locations.

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地理的に隔たった3個所のキャベツ圃場におけるヤガの1種, *Trichoplusia ni*  
の空間分布—数種の分布指数による比較解析—

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ここで取り上げた  $C_s$ ,  $I_s$ ,  $I_p$ , および TAYLOR の乗則 (TPL) のどの分布指数による解析でも, *T. ni* の幼虫個体群の分布は集中的だがその程度は一般に弱いことが示された。ただし, 分布集中度は若令幼虫のほうが老令幼虫より概して高く, 地域間 (カリフォルニア, テキサス, ルイジアナ) でも変異がみられ, 地域や発育段階ごとに異なるサンプリング計画を立てる必要性が示唆された。  $C_s$ ,  $I_s$ ,  $I_p$  の3指数については相関分析によって平均密度への依存度を検討したところ, どの場合にも相関は弱く, 密度の影響はあまり受けていないと考えられた。この3指数間の相関関係についても統計的解析と考察を行った。