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LARS Technical Report 030178

CR 151660

Bayesian Classification in a Time-varying Environment

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(NASA-CR-151660) BAYESIAN CLASSIFICATION IN
A TIME-VARYING ENVIRONMENT (Purdue Univ.)
21 p HC A02/MF A01 CSCI 12E

N78-20571

Unclas
G3/43 11816



The Laboratory for Applications of Remote Sensing
Purdue University West Lafayette, Indiana

1978

1. Report No.	2. Government Accession No.	3. Recipient's Catalog No.	
4. Title and Subtitle Bayesian Classification in a Time-Varying Environment		5. Report Date March 1, 1978	6. Performing Organization Code
		8. Performing Organization Report No. 030178	
7. Author(s) Philip H. Swain		10. Work Unit No.	
9. Performing Organization Name and Address Laboratory for Applications of Remote Sensing 1220 Potter Drive West Lafayette, Indiana 47906		11. Contract or Grant No. NAS9-14970	
		13. Type of Report and Period Covered	
12. Sponsoring Agency Name and Address NASA		14. Sponsoring Agency Code	
		15. Supplementary Notes	
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17. Key Words (Suggested by Author(s)) pattern classification, multitemporal observations, remote sensing.		18. Distribution Statement	
19. Security Classif. (of this report)	20. Security Classif. (of this page)	21. No. of Pages 18	22. Price*

*For sale by the National Technical Information Service, Springfield, Virginia 22161

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Abstract

This paper deals with the problem of classifying a pattern based on multiple observations made in a time-varying environment. The identity of the pattern may itself change. A Bayesian solution is derived, after which the conditions of the physical situation are invoked to produce a "Cascade" classifier model. Experimental results based on remote sensing data demonstrate the effectiveness of the classifier.

Key words: pattern classification, multitemporal observations, remote sensing

BAYESIAN CLASSIFICATION IN A
TIME-VARYING ENVIRONMENT

Philip H. Swain¹

Introduction

We pose the following pattern classification problem:

A series of observations is made on a pattern in a time-varying environment. The identity of the pattern itself may change. It is desired to classify the pattern after the current observation is made, drawing on information derived from earlier observations plus knowledge about the statistical behavior of the environment.

An example of such a situation arises in remote sensing applications in which the sensor system can make multiple passes over the same ground area [1]. The identity of the ground cover may change between passes. In general it is desired to determine the current identity of the ground cover, but past observations can be helpful in accomplishing the identification.

Approach

The classification strategy we shall develop is a Bayes optimal (minimum risk) strategy [2]. In the ordinary single

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observation case, the approach is to select a decision rule so as to minimize the conditional average loss

$$L_X(\omega_i) = \sum_{j=1}^m \lambda_{ij} p(\omega_j | X) \quad (1)$$

where

X is an n -variate observation (feature) vector

$\{\omega_j, j=1, 2, \dots, m\}$ is the set of m classes

λ_{ij} is the cost resulting from classifying into class i a pattern actually from class j

$p(\omega_j | X)$ is the conditional probability that, given observation X , its class is ω_j

That is, $L_X(\omega_i)$ is the expected loss incurred if an observation X is classified as ω_i . Commonly [2] λ_{ij} is taken to be the "0-1 loss function," i.e.,

$$\begin{aligned} \lambda_{ij} &= 0, \quad i = j && \text{(no cost for correct classification)} \\ &= 1, \quad i \neq j && \text{(unit cost for an error)} \end{aligned}$$

Then Eq. (1) becomes

$$L_X(\omega_i) = 1 - p(\omega_i | X) \quad (2)$$

and an appropriate decision rule which will minimize $L_X(\omega_i)$ is:

Decide $X \in \omega_i$ if and only if

$$p(X | \omega_i) p(\omega_i) = \max_j p(X | \omega_j) p(\omega_j) \quad (3)$$

where $p(X | \omega_i)$ is the probability density function for the obser-

vations associated with class ω_i and $p(\omega_i)$ is the a priori probability of class ω_i . Thus the set of products $\{p(X|\omega_i)p(\omega_i), i=1, 2, \dots, m\}$ is a set of discriminant functions for the classification problem.

We now generalize this Bayes optimal approach to the case of a series of observations. It will be convenient to assume that observations are made at two times. Generalization to a larger number of observation times is straightforward.

Let $X_1 = X(t_1)$ and $X_2 = X(t_2)$ be n -variate random vectors, the pattern observations at times t_1 and t_2 , respectively.

Let $\{\omega_i = \omega_i(t_1) | i=1, 2, \dots, m_1\}$ be the set of possible classes at time t_1 , and let $\{\omega_i = \omega_i(t_2) | i=1, 2, \dots, m_2\}$ be the set of possible classes at time t_2 .

We define a compound conditional average loss

$$L_{X_1 X_2}(\omega_i) = \sum_{j=1}^{m_2} \lambda_{ij} p(\omega_j | X_1, X_2) \quad (4)$$

where λ_{ij} is the cost resulting from classifying into class i , at time t_2 , a pattern actually from class j . In this case $p(\omega_j | X_1, X_2)$ is the a posteriori probability that, given the observations X_1 at time t_1 and X_2 at time t_2 , the class of the pattern at time t_2 is ω_j .

Once again assuming a "0-1 loss function," Eq. (4) becomes

$$L_{X_1 X_2}(\omega_i) = 1 - p(\omega_i | X_1, X_2) \quad (5)$$

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which is minimized if we choose ω_i to maximize the a posteriori probability $p(\omega_i | X_1, X_2)$. Thus an appropriate set of discriminant functions for a Bayes optimal classification strategy is the set of a posteriori probabilities; i.e.

$$\left\{ p(\omega_i | X_1, X_2), \quad i = 1, 2, \dots, m_2 \right\}$$

As usual, however, we wish to derive a set of equivalent discriminant functions expressed in terms of class-conditional density functions and a priori probabilities as in Eq. (3). This may be accomplished proceeding as follows. First we write:

$$p(\omega | X_1, X_2) = \frac{p(\omega, X_1, X_2)}{p(X_1, X_2)} \quad (6)$$

For fixed X_1 and X_2 , the denominator in Eq. (6) is constant.

Let $c = 1/p(X_1, X_2)$ and write Eq. (6) as

$$\begin{aligned} p(\omega | X_1, X_2) &= cp(\omega, X_1, X_2) \\ &= c \sum_{\nu} p(X_1, X_2, \nu, \omega) \\ &= c \sum_{\nu} p(X_1, X_2 | \nu, \omega) p(\nu, \omega) \\ &= c \sum_{\nu} p(X_1, X_2 | \nu, \omega) p(\omega | \nu) p(\nu) \quad (7) \end{aligned}$$

The summation is over the classes which can occur at time t_1 . The factor $p(X_1, X_2 | \nu, \omega)$ is a joint class-conditional density; $p(\omega | \nu)$ may be interpreted as a transition probability (the probability that the class is ω at time t_2 given the class was ν at time t_1); and $p(\nu)$ is an a priori probability.

Thus, the multiobservational decision rule analogous to Eq. (3) is:

$$\begin{aligned} & \text{Decide } X_2 \in \omega_i \text{ if and only if} \\ & \sum_{k=1}^{m_1} p(X_1, X_2 | v_k, \omega_i) p(\omega_i | v_k) p(v_k) \\ & = \max_j \sum_{k=1}^{m_1} p(X_1, X_2 | v_k, \omega_j) p(\omega_j | v_k) p(v_k) \quad (8) \end{aligned}$$

and the set of discriminant functions is the set of sums of products:

$$\left\{ \sum_{k=1}^{m_1} p(X_1, X_2 | v_k, \omega_i) p(\omega_i | v_k) p(v_k), \quad i=1, 2, \dots, m_2 \right\}. \quad (9)$$

A "Cascade" Implementation

In practice, the terms in the discriminant functions must be estimated from "training samples." The most formidable job is estimating the $m_1 \cdot m_2$ joint class-conditional densities $p(X_1, X_2 | v_k, \omega_i)$, each of which is of dimension $2n$.² Clearly a large number of training samples will be required. When certain approximations can be justified, the situation is eased considerably. We shall now show that these approximations lead to a rather attractive model for a multitemporal classifier.

²The observation vectors need not be of the same dimensionality.

If X_1 has n_1 components and X_2 has n_2 components, the $p(X_1, X_2 | v, \omega)$ is N -variate, where $N = n_1 + n_2$.

We are accustomed to assuming class-conditional independence in the spatial domain; i.e., given the class at a particular point, the random variable which is the measurement vector at that point is independent of the class or measurement vector at any other point. Applying this same idea to multitemporal measurements at a given point, we say that given the classes v_k at t_1 and ω_i at t_2 , the random variables X_1 and X_2 are independent. Then we can write

$$p(X_1, X_2 | v_k, \omega_i) = p(X_1 | v_k, \omega_i) p(X_2 | v_k, \omega_i) \quad (10)$$

and furthermore

$$\begin{aligned} p(X_1 | v_k, \omega_i) &\cong p(X_1 | v_k) \\ p(X_2 | v_k, \omega_i) &\cong p(X_2 | \omega_i) \end{aligned} \quad (11)$$

Imposing these conditions, it follows that

$$p(X_1, X_2 | v_k, \omega_i) = p(X_1 | v_k) p(X_2 | \omega_i).$$

The discriminant functions, Eq. (9), then become

$$\left\{ \begin{aligned} &\sum_{k=1}^{m_1} p(X_1 | v_k) p(X_2 | \omega_i) p(\omega_i | v_k) p(v_k), \\ & i=1, 2, \dots, m_2 \end{aligned} \right\} \quad (12)$$

From Eq. (12) we can model the discriminant function calculations as indicated in Figure 1, from which we derive the term "cascade classifier" to describe this multistage classifier.

Simulation and Experimental Results

The cascade classifier model was programmed and applied to the analysis of a set of Landsat multispectral data. The data,

collected by the satellite on two successive passes, eighteen days apart, over Fayette County, Illinois (see Table 1), were geometrically registered at Purdue University's Laboratory for Applications of Remote Sensing. The objective of the analysis was to discriminate among the ground cover classes "corn", "soybeans", "woods", and "other", where the last category was simply a catch-all consisting of water, pasture, fallow and other relatively minor ground covers. Each class was actually decomposed in the analysis process into a union of subclasses, each having a data distribution describable as approximately multivariate normal.³

To provide a baseline for comparison, the data from each of the passes was first analyzed separately. The a priori probabilities of the classes were approximated as being equal, and 557 test samples, independent of the training samples, were used to evaluate the results. As shown in Table 1(a) and (b), the performance of this conventional maximum likelihood classifier was 68% correct for the June 29, 1973 data, and 72% correct for the July 17, 1973 data.

To implement the cascade analysis, it was assumed unlikely that the ground cover would change identity over so short a time span. Accordingly, the transition probabilities were estimated as follows:

$$p(\omega_i | v_k) = 0.8 \quad \text{for } \omega_i = v_k, \quad (13a)$$

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and all other transition probabilities were set equal and such that

³All probability densities were assumed to be multivariate normal (Gaussian), characterized by mean vector and covariance matrix.

$$\sum_{\substack{i \\ \omega_i \neq v_k}} p(\omega_i | v_k) = 0.2. \quad (13b)$$

Again the a priori probabilities were assumed equal and the same test samples were used to evaluate the results.

The results of this multitemporal classification, Table 1(c), were substantially better than either of the unitemporal analyses. The overall results were 84% correct. In addition, the performance for each class was better than the best attained for the class in either of the unitemporal analyses. The unitemporal and multitemporal results are compared in Figure 2.

The results can be sensitive, however, to the specification of the transition probabilities and a priori probabilities. This is demonstrated in the following experiment.

Landsat data from two passes over Grant County, Kansas, were analyzed in a manner similar to that used for the Fayette County data. In this case, the two passes were separated by more than two months and a different set of classes was involved (Table 2). The transition probabilities were specified as in Eq. (13a) and (13b); equal a priori probabilities were assumed.

As shown in Table 2 and Figure 3, in this case the overall performance of the multitemporal cascade classifier was only marginally better than the best unitemporal result. A closer look at the class-by-class results is revealing. The largest detractors from the multitemporal results were the classes "alfalfa" and "pasture." In both of these cases, the unitemporal results for the second pass were substantially lower than those obtained in the first pass. (There are physical explanations for why this is reasonable, but this is not germane to our exploration of classifier behavior.)

Let us examine the impact that the relatively arbitrary assignment of transition probabilities has on the classification results. In case the actual transition probabilities are not known (which was true for the cited examples), the assignment can be made anywhere between two extremes. On the one hand, it could be assumed that

$$p(\omega_i | v_k) = \frac{1}{m_1}, \quad k = 1, 2, \dots, m_1$$

i.e., equiprobable transitions. Then the discriminant functions have the form

$$\begin{aligned} & \sum_{k=1}^{m_1} p(X_1 | v_k) p(X_2 | \omega_i) \frac{1}{m_1} p(v_k) \\ &= \frac{1}{m_1} p(X_2 | \omega_i) \sum_{k=1}^{m_1} p(X_1 | v_k) p(v_k) \\ &= \frac{1}{m_1} p(X_2 | \omega_i) p(X_1). \end{aligned}$$

Since $\frac{1}{m_1}$ and $p(X_1)$ will be common to each of the discriminant functions, the decision will depend only on $p(X_2 | \omega_i)$ and will be independent of the first-stage results.

On the other hand we could make $p(\omega_i | v_i) = 1$ and $p(\omega_i | v_j) = 0$, $j \neq i$. Then the discriminant functions become

$$p(X_1 | v_i) p(X_2 | \omega_i) p(v_i).$$

Thus, in a sense, the contributions from the two stages are weighted equally.

There is no way to make the first stage input dominate the second stage.

In view of these considerations, another classification of the Grant County data was performed. In this case, the transition probabilities $p(w_i|v_i)$ were set equal to unity for the "alfalfa" and "pasture" classes in order to give as much strength as possible to the first stage results. Table 3 and Figure 3 show the outcome of this classification. The confusing influence resulting from the second stage data has been reduced.

It is interesting to compare the results obtained using the cascade classifier to results produced by a "conventional" maximum likelihood classifier using all of the multitemporal features simultaneously. To perform the latter classifications, equal a priori probabilities were assumed. The results were:

Fayette County: 80.8 percent correct

Grant County: 64.1 percent correct

It is curious that neither of these results is any better than the cascade classifier results achieved. It is possible that these slightly poorer results represent the price paid for having to estimate 8-dimensional statistics as opposed to 4-dimensional statistics in the face of limited training data.

Discussion and Conclusions

The approach we have adopted for classifying data in a non-stationary environment was based on application of classical statistical decision theory in a straightforward manner. However, we used the conditions of the problem to approximate some of the statistical quantities involved. This step simplified the interdependencies of the data involved and led to a "cascade classifier"

model. In the time-varying environment, this model is seen to:

(1) Successfully incorporate the temporal information in the classification process, resulting in improved classification accuracy;

(2) Reduce the dimensionality of the probability functions used and thereby make less stringent demands with respect to the size of the training set required;

(3) Facilitate distribution of the computational load over time.

Each time a set of observations becomes available, discriminant functions are calculated which can be used, if desired, to make a classification. However, the values of the discriminant functions are also passed along and contribute to a new set of discriminant functions calculated when the next set of observations is obtained. Although we have demonstrated the use of the cascade model only for the case of two stages, extension to an arbitrary number of stages presents no difficulty.

The prospective user of this approach should be aware that a casual implementation of the likelihood computers may result in computational difficulties of two sorts: loss of precision and very large computation times as compared with, say, a conventional Gaussian maximum likelihood classifier. Both of these difficulties can be overcome or at least substantially reduced by appropriate measures (scaling, ignoring zero terms, etc.) in carrying out the likelihood computations.

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Acknowledgements

The author wishes to thank Mr. Carlos A. Pomalaza for programming the cascade classifier model and testing it with the remote sensing data. This research was supported in part by NASA Contract NAS9-14970.

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of the Fayette County, Illinois, data.

(a) June 29, 1973 data

Group	No. of Samples	Percent Correct	No. of Samples Classified into			
			CORN	OTHERS	SOYBEAN	WOODS
CORN	186	65.1	121	36	24	5
OTHERS	100	40.0	33	40	22	5
SOYBEAN	227	82.4	10	30	187	0
WOODS	44	72.7	0	4	8	32
TOTAL	557		164	110	241	42

OVERALL PERFORMANCE = 68.2 percent correct

(b) July 17, 1973

Group	No. of Samples	Percent Correct	No. of Samples Classified Into			
			CORN	OTHERS	SOYBEAN	WOODS
CORN	186	89.2	166	16	1	3
OTHERS	100	45.0	38	45	15	2
SOYBEAN	227	73.6	24	36	167	0
WOODS	44	56.8	4	9	6	25
TOTAL	557		232	106	189	30

OVERALL PERFORMANCE = 72.4 percent correct

(c) Multitemporal results (cascade classifier)

Group	No. of Samples	Percent Correct	No. of Samples Classified Into			
			CORN	OTHER	SOYBEAN	WOODS
CORN	186	90.3	168	11	4	3
OTHERS	100	48.0	29	48	20	3
SOYBEAN	227	94.3	3	10	214	0
WOODS	44	84.1	0	5	2	37
TOTAL	557		200	74	240	43

OVERALL PERFORMANCE = 83.8 percent correct

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Table 2. Test results for classification of the
Grant County, Kansas, data.

(a) May 9, 1974

Group	No. of Samples	Percent Correct	No. of Samples Classified Into				
			ALFALFA	CORN	FALLOW	PASTURE	WHEAT
ALFALFA	58	84.5	49	0	0	0	9
CORN	428	57.0	0	244	183	1	0
FALLOW	526	54.4	0	196	286	36	8
PASTURE	1513	52.6	127	148	220	796	227
WHEAT	913	82.5	97	17	0	49	767
TOTAL	3455		273	605	689	882	1006

Overall Performance = 62.0 percent correct

(b) July 20, 1974

Group	No. of Samples	Percent Correct	No. of Samples Classified Into				
			ALFALFA	CORN	FALLOW	PASTURE	WHEAT
ALFALFA	58	5.2	3	3	0	10	42
CORN	428	53.0	15	227	105	15	66
FALLOW	526	62.9	0	113	331	5	77
PASTURE	1513	42.4	64	329	213	641	266
WHEAT	913	76.2	22	108	33	58	709
TOTAL	3455		104	780	682	729	1160

Overall Performance = 55.3 percent correct

(c) Multitemporal results (cascade classifier)

Group	No. of Samples	Percent Correct	Number of samples classified Into				
			ALFALFA	CORN	FALLOW	PASTURE	WHEAT
ALFALFA	58	41.4	24	0	0	2	32
CORN	428	59.6	5	255	165	1	2
FALLOW	526	76.4	0	107	402	2	15
PASTURE	1513	46.3	101	205	224	701	282
WHEAT	930	88.3	77	19	0	13	821
TOTAL	3455		207	586	791	719	1152

Table 3. Cascade classifier results for adjusted transition probabilities (Grant County data).

Group	No. of Samples	Percent Correct	Number of samples classified Into				
			ALFALFA	CORN	FALLOW	PASTURE	WHEAT
ALFALFA	58	94.8	55	0	0	0	3
CORN	428	70.3	5	301	122	0	0
FALLOW	526	68.1	0	139	358	7	22
PASTURE	1513	48.1	105	211	195	727	275
WHEAT	930	89.1	82	9	0	10	829
<hr/>			<hr/>				
TOTAL	3455		247	660	675	744	1129

Overall Performance = 65.7 percent correct

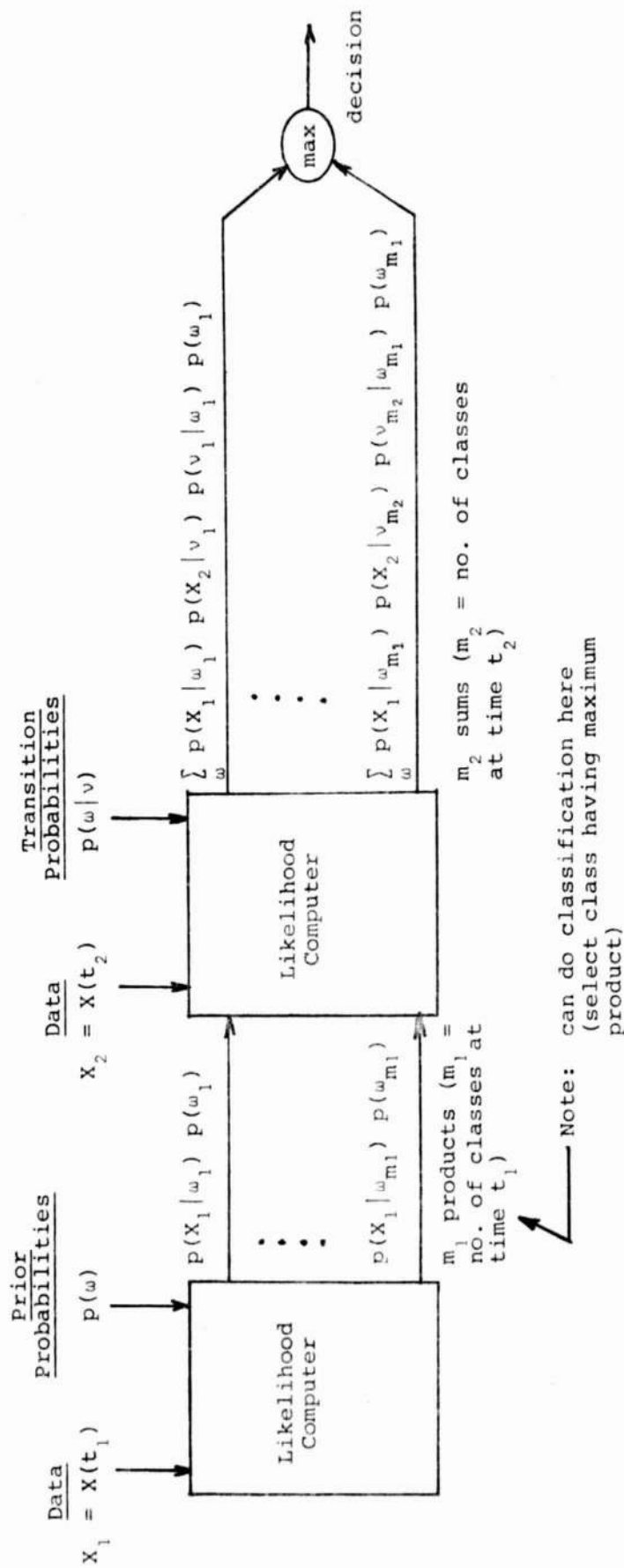


Figure 1. The cascade classifier model.

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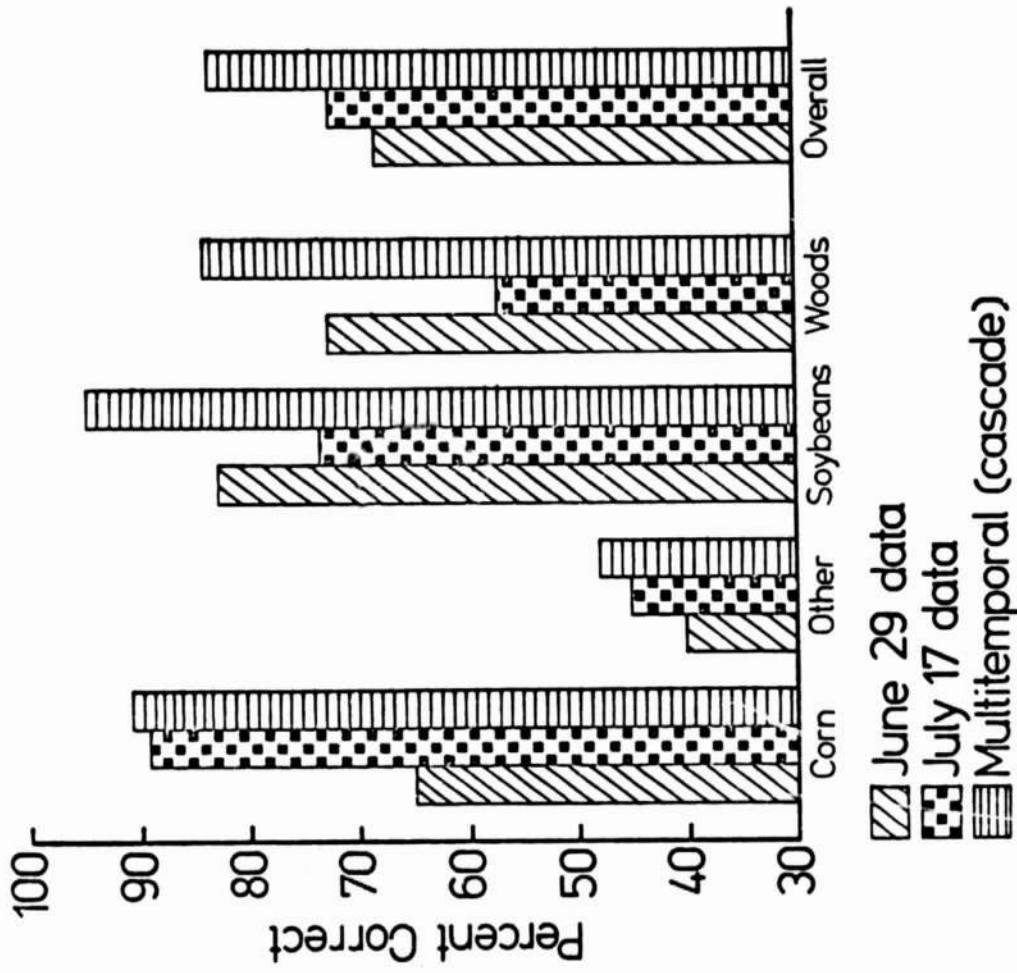


Figure 2. Test results for Fayette County data.

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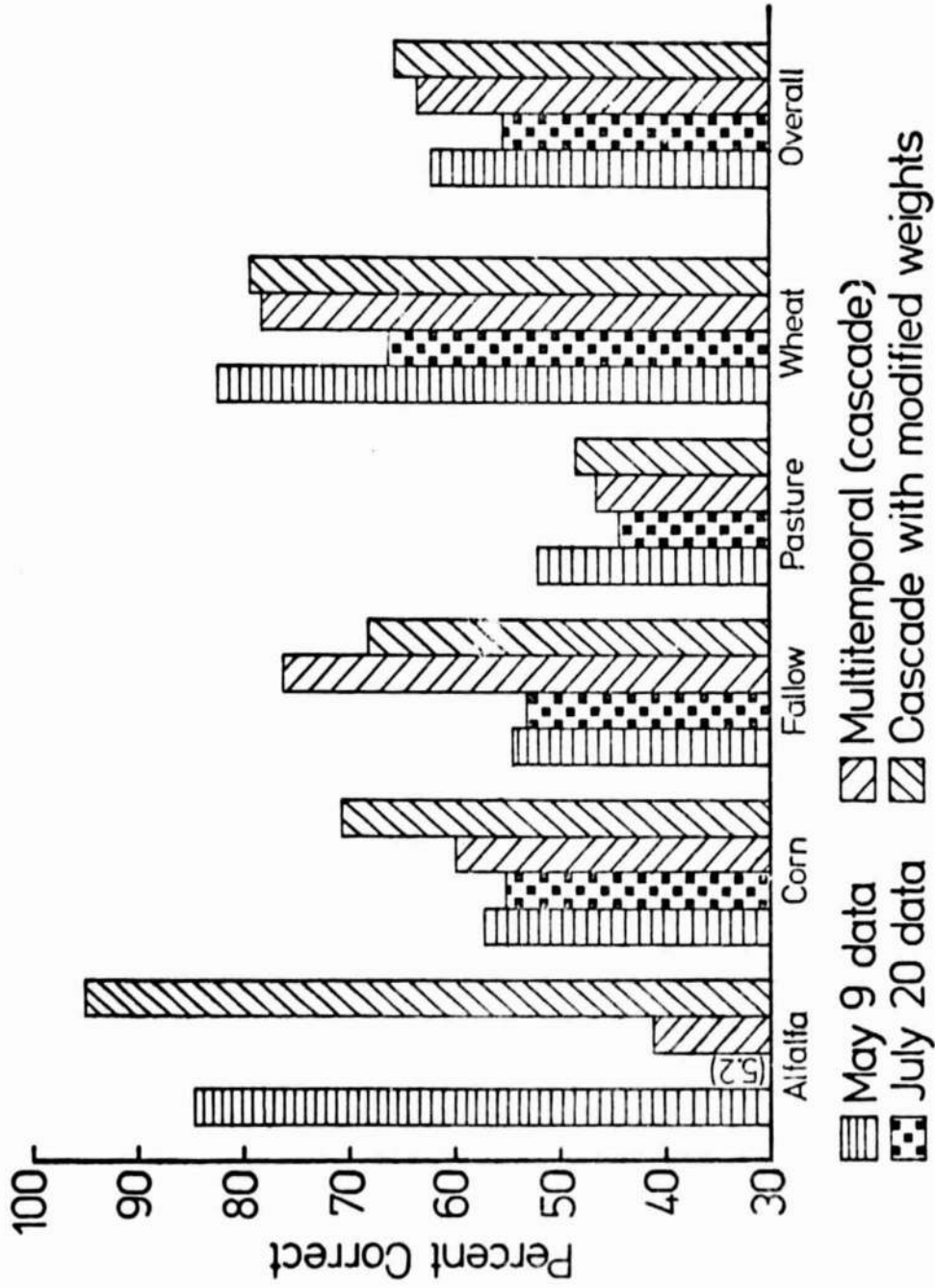


Figure 3. Test results for Grant County data.