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An analysis of the interaction of incentive and relative event frequency effects.

Stanley George Lipkin

University of Massachusetts Amherst

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AN ANALYSIS OF THE INTERACTION OF INCENTIVE
AND RELATIVE EVENT FREQUENCY EFFECTS

A Dissertation Presented
By
Stanley G. Lipkin

Submitted to the Graduate School of the
University of Massachusetts in
partial fulfillment of the requirements for the degree of
DOCTOR OF PHILOSOPHY

May, 1966
Psychology
AN ANALYSIS OF THE INTERACTION OF INCENTIVE AND RELATIVE EVENT FREQUENCY EFFECTS

A Dissertation

By

Stanley G. Lipkin

Approved as to style and content by:

[Signatures and names with dates]

Date

May 25, 1961
ACKNOWLEDGMENTS

I am very grateful for the guidance supplied by each of the members of my Dissertation Committee, Drs. Jerome L. Myers, Pao L. Cheng, and John W. Moore. I am especially indebted to Professor Myers, the Committee Chairman, who has been extremely generous in his personal contributions to this research.

Thanks are also due to Mrs. Ruth Crabtree and Mrs. Dorothy Thayer for technical assistance during the execution of the experiment, and to the students who made helpful suggestions in the design of the experiment.
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INTRODUCTION

It is a common assumption in reinforcement theories that variation in incentive magnitude produces corresponding changes in performance. However, this assumption is not supported by the results of research in several experimental situations in which incentive has been varied between Ss with each S experiencing a single incentive level. In a review of incentive studies with animal Ss, Pubols (1960) concluded that variation in incentive magnitude has no significant effect on time-independent measures of learning when an absolute method (i.e., between-Ss design) is used. Pubols' conclusion has more recently received support from studies employing human Ss in which incentive magnitude was varied using the absolute method. For example, Harley (1965b) observed no significant difference in mean percent correct responses between two groups rewarded with 0% and 25% for each correct response in a list of paired associates (PAs). Myers, Fort, Katz, and Suydam (1963) found only insignificant differences in $P(A_1)$ (the probability of choosing the more frequent event) between two groups which risked 10% and 1% on each trial in a probability learning task. In light of these facts, Pubols' conclusion about the lack of an incentive effect under the absolute
method may be extended to include human Ss.

While the absolute method has led to negative findings, a differential method in which each S experiences more than one incentive level has consistently produced significant incentive effects. In his review, Pubols concluded that animal learning is directly related to incentive magnitude when a differential, or within-Ss, design is used. Once more, later research has revealed parallel results with human Ss. Harley (1965a) found significantly better PA anticipation for high incentive pairs using the differential method whereas he failed to observe this incentive effect with the absolute method of incentive presentation (1965b). Similarly, Lipkin, Schnorr, Suydam, and Myers (1965) found \( P(A_1) \) directly related to the amount of risk in a noncontingent two-choice study. In a direct comparison of a between-Ss and a within-Ss design, Schnorr, Lipkin, and Myers (In press) found a marked incentive effect under the within-Ss manipulation while both risk levels of the between-Ss design led to asymptotic \( P(A_1) \) equal to that obtained under the higher within-Ss risk level. Apparently, it may be inferred as a general principle of behavior that the effects of incentive on performance depend upon the type of experimental design employed. A necessary condition for observing an effect among positive levels of incentive would seem to be that each S has experience with more than
one incentive level.

The Schnorr et al. study allowed observation of the manner in which exposure to more than one incentive level operates to elicit an incentive effect. The significance of the incentive effect in their within-Ss design was attributed to a negative contrast effect in the form of a "depression" of P(A₁) on low risk trials. There was no concurrent enhancement of performance on the high risk trials. Schnorr, et al. also collected subjective estimates of π, the proportion of occurrences of the more frequent event. These measures (π₃) also exhibited negative contrast and were significantly correlated with P(A₁) for each subject. This may indicate that the incentive effect is caused by a distortion in Ss' perception of π under the differential method; however, other explanations cannot yet be ruled out.

The finding of a negative contrast effect between levels of incentive is not novel in animal or human work. With two levels of reward "randomly alternated," Bower (1961) found that rats decreased their running speed in the small reward runway after an initial block of trials in which running speed was equal for the two alleys. As controls, he also ran constant reward groups with the large and small rewards. In the differential group there was a significant depression in the small reward runway relative to the constant
small reward group without concurrent elation on high reward trials. Spence (1956) also found that a shift to a smaller reward produced a depression in the reciprocal of running time of rats while a shift to a larger reward did not lead to elation relative to a pre-shift asymptote. Similarly, with human Ss in a three-choice situation, Swensson (1965) observed a significant depression in P(A1) for a group shifted from payoff to no payoff, but there was no corresponding upward shift for a group with the opposite change in reward.

The present experiment was designed to investigate negative contrast effects between two incentives as a function of event frequencies under each incentive level. The methodology was very similar to that used in the within-Ss portion of the Schnorr et al. study. In this procedure, Ss must predict which of two designs is on the back of each card in a deck while the amount to be risked on any prediction appears on the exposed side of that card. As in the Schnorr et al. study, random sequences of 50 high risk and 50 low risk trials per trial block were used. However, in the present design eight groups were formed on the basis of π on high and low risk trials (π_H = .60, .70; π_L = .50, .60, .70, .80). Whereas Schnorr et al. had held both π_H and π_L equal to .60, the present study pits the strength of the incentive effect against an increasingly high π_L, and thereby
affords information about the relative strengths of and incentive.

Schnorr et al. found that $P(A_1)$ approximated .50 on low risk trials which had a $\pi$ of .60, and suggested that this "chance level" performance may have resulted from Ss "ignoring" low risk trials to deal more effectively with the high risk trials. The present study provided a test of this hypothesis through a comparison of performance levels on low risk trials across the eight groups. If the hypothesis is correct, $P(A_1)$ in the present study should approximate .50 on low risk trials for all four groups at each of the two $\pi_H$ levels. However, to find no differences in $P(A_1)$ over these four $\pi_L$ levels would be surprising since $\pi$ has generally been the most potent variable in the probability learning literature. On the other hand, the Schnorr et al. findings may have been due to a distortion of $\pi_S$, which would lead to a prediction of low risk performance in the present study increasing with, but consistently undershooting, $\pi_L$.

Several other interesting comparisons and theoretical implications are suggested by the two-choice discrimination studies of Popper and Atkinson (1958), Atkinson, Bogartz, and Turner (1959), Uhl (1964), and others. The procedure of the present study is identical to that used in two-choice discrimination studies with the exception that the pre-trial
cue in those studies has no value aside from its information role. In view of the similar experimental paradigms of discrimination studies and the present within-Ss incentive manipulation, a brief description of some relevant mathematical models of discrimination learning seems appropriate. There are several models available (e.g., Burke and Estes, 1957; Estes, 1959; Atkinson, 1958) which have been able to predict the general data trends in certain relatively simple discrimination experiments.

Popper and Atkinson (1958) and Atkinson et al. (1959) have shown that \( P(A_1) \) given a highly reliable cue (i.e., one with a high or low \( \pi \) value) is an increasing function of the reliability of a second cue. Within the range of \( \pi \) values of the present study, these two discrimination studies have also shown that \( P(A_1) \) given the cue of lower reliability was displaced away from \( \pi \)-matching toward the \( \pi \) value associated with the more reliable cue. This convergence of response probabilities or generalization between cues is relevant in judging the appropriateness of mathematical models.

Provided such generalization is not observed in the present experiment, the pattern model (Estes, 1959) or an extension designed to handle the overshooting typically found with incentives (Myers and Atkinson, 1964) may be applicable to the data. In the pattern model, responses
become conditioned to total patterns of stimulation rather than to stimulus components. There is no parameter in the model to relate degree of stimulus similarity to discrimination performance. Thus, successful discrimination is predicted whenever the two relevant stimulus patterns are not identical. Since the two pre-trial cues in the present study (indicating the incentive levels) were noticeably different, the pattern model would predict asymptotic probability matching.

On the other hand, a pattern model would not be compatible with the data if generalization between the two incentives occurs. For instance, the pattern model must predict equal asymptotes for 10-chip trials with a \( \pi_H \) of .60 regardless of differences in \( \pi_L \) across groups. If observed 10-chip performance is a function of the value of \( \pi_L \) as well as \( \pi_H \), a model which deals with a stimulus overlap would be more appropriate (Burke and Estes, 1957; Bush and Mosteller, 1951). In the Burke-Estes component model, increased overlap between two hypothetical stimulus sets leads to predictions of poorer discrimination performance. Consequently, if 10-chip performance is a function of \( \pi_L \) or 1-chip performance a function of \( \pi_H \), then the component model could account for such generalization in terms of stimulus overlap or similarity.

Aside from the component model, other models have been
developed by Atkinson (1958) and Lee (1966) which can predict generalization. Atkinson's (1958) Markov model can fit the general data trends of the Popper and Atkinson (1958) and Atkinson et al. (1959) studies. According to this model, the S may either make an observing response or not at the start of each trial; and if he does not observe, he samples stimulus elements common to the two sets of discriminative stimuli. Whereas this sampling of common elements enables the Atkinson model to predict generalization between cues, Lee (1966) accounts for the same results with a conditioning-parameter model. This model also allows a prediction of generalization since reinforcement not only increases the probability of a response under identical stimulating conditions, but also under all other conditions. While both this model and the component model can predict generalization, \( P(A_i) \) must be a linear function of \( \pi_H \) and \( \pi_L \). However, the Atkinson (1958) model has the added advantage of being able to predict a nonlinear relationship. Although one or more of these models may be able to predict some of the general data trends in the present study, some inadequacy was expected because of the overshooting of \( \pi \) usually observed in incentive studies.

Another important consideration for mathematical models in the present study was the observed probability that a given response occurred on trial \( n \) given the response-event
combination on trial \( n-1 \). Analysis of such first-order conditional statistics from the Schnorr et al. study revealed gross deviations from theoretical predictions. It was found that the probability of repeating an \( A_1 \) response was greater following an incorrect than a correct prediction of \( A_1 \) on the preceding trial, i.e.,

\[
P(A_1,n | A_1,n-1 E_2,n-1) > P(A_1,n | A_1,n-1 E_1,n-1)
\]

Precisely the opposite order of these conditionals is predicted by all existing models which call for increments in response probability following reinforcement and decrements following nonreinforcement. Halpern (1965) and Moore and Halpern (1966) have obtained similar inversions in the conditional statistics from two-choice auditory discrimination experiments. This type of deviation from theory is interesting with regard to the models, and it is, perhaps, much more important in presenting a basic challenge to reinforcement theory in general. Therefore, in the present study, the observed conditional statistics and possible generalization between incentives were assessed in relation to existing models.
METHOD

Subjects. The Ss were 80 male and 80 female undergraduate volunteers from the introductory psychology course at the University of Massachusetts.

Design and Procedure. The major independent variables were incentive level (1 or 10 chips) and the proportion of the more frequent event within each incentive level (π_H and π_L). There were 8 experimental groups, of 20 Ss each, formed by crossing the 2 π_H levels and 4 π_L levels. Within each group, the 20 Ss were evenly divided among 4 combinations of 2 experimenters and the 2 sexes to form 32 cells of 5 Ss each.

After being randomly assigned to one of the 32 cells, each S was required to predict which of two geometric designs (point or line) was on the back of each card of a 100 card deck. Each of the two experimenters used a different geometric design as the more frequent event (E_1) to allow for statistical control of design preferences and experimenter differences as a single confounded source of variance. On the face of 50 cards there was a "10" and on the face of the other half of the cards there was a "1" designating the number of chips S must risk on the succeeding prediction. Each S was given an initial stake of 200 chips. The relative event frequencies (E_1:E_2) were specific to the level
of risk so that the overall $\pi$ for each $\text{S}$ was the average of his specific values of $\pi_H$ and $\pi_L$.

Each experimental session consisted of 400 trials obtained by having $\text{S}$ go through the deck four times. The sequence of incentive levels and events was randomized by thoroughly shuffling the deck in full view of $\text{S}$ before each block of trials. This procedure and the instructions were intended to discourage any search for specific patterns in the cards. The $\text{Ss}$ were run one at a time and were self-paced.

Three types of dependent measures were obtained for the data analysis. The marginal and conditional response probabilities were calculated from the sequence of each $\text{S's}$ responses. Following the fourth block of trials each $\text{S}$ was given a questionnaire to record his subjective estimates of (a) the proportion of 10-chip trials, (b) $\pi_H$, and (c) $\pi_L$. 
RESULTS AND DISCUSSION

Marginal response probabilities. Table 1 shows the F-ratios from the analysis of variance of $P(A_1)$ scores for the eight experimental groups. The magnitude of $P(A_1)$ overall eight groups varied directly and significantly as a function of both $\pi_H$: $F(1,128) = 24.43, p<.001$, and of $\pi_L$, $F(3,128) = 17.45, p<.001$. The overall incentive effect was also significant, $F(1,128) = 78.57, p<.001$, with $P(A_1)$ greater on 10-chip trials than on 1-chip trials. A more analytic description of the combined effects of $\pi_H$, $\pi_L$, and incentive is presented in Fig. 1 which shows the course of training for each group of 20 Ss. For convenience, each group is designated according to the $\pi_H: \pi_L$ combination used.

The incentive effect shown for the .60:.50 group was expected since both $\pi$ and incentive favored the observed separation of performance in this group. However, the data for the .60:.60 group is interesting because it replicated the Schnorr et al. study. As the figure shows, $P(A_1)$ was about 19% higher on the 10-chip trials after 400 trials in which $\pi_H$ and $\pi_L$ both equalled .60. The magnitude of this incentive effect and the performance levels under the two incentive conditions are quite similar to those obtained by Schnorr et al. and by Lipkin et al. (1965).
Table 1

F-Ratios from Analysis of Variance of P(A_1) Scores

<table>
<thead>
<tr>
<th>Source of Variance</th>
<th>df</th>
<th>F-Ratio</th>
<th>p-value</th>
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<td>Experimenter (E)</td>
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<td>Sex (X)</td>
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<td>.06</td>
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<tr>
<td>L H</td>
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Table 1 (cont'd.)

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<td>T S/L E X H</td>
<td>384</td>
<td>(35.77)</td>
<td></td>
</tr>
<tr>
<td>Source of Variance</td>
<td>df</td>
<td>F-Ratio</td>
<td>p-Value</td>
</tr>
<tr>
<td>--------------------</td>
<td>----</td>
<td>---------</td>
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</tr>
<tr>
<td>T I</td>
<td>3</td>
<td>2.01</td>
<td>&lt;.20</td>
</tr>
<tr>
<td>T I H</td>
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</tr>
<tr>
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<td>3.08</td>
<td>&lt;.005</td>
</tr>
<tr>
<td>TE I</td>
<td>3</td>
<td>.36</td>
<td></td>
</tr>
<tr>
<td>TX I</td>
<td>3</td>
<td>.15</td>
<td></td>
</tr>
<tr>
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<tr>
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<td>.15</td>
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</tr>
<tr>
<td>TX I H</td>
<td>3</td>
<td>1.45</td>
<td></td>
</tr>
<tr>
<td>TL E I</td>
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<td>.64</td>
<td></td>
</tr>
<tr>
<td>TL X I</td>
<td>9</td>
<td>2.13</td>
<td>&lt;.05</td>
</tr>
<tr>
<td>TX E I</td>
<td>3</td>
<td>.15</td>
<td></td>
</tr>
<tr>
<td>TL E I H</td>
<td>9</td>
<td>.58</td>
<td></td>
</tr>
<tr>
<td>TL X I H</td>
<td>9</td>
<td>1.23</td>
<td></td>
</tr>
<tr>
<td>TX E I H</td>
<td>3</td>
<td>.27</td>
<td></td>
</tr>
<tr>
<td>TL E X I</td>
<td>9</td>
<td>.69</td>
<td></td>
</tr>
<tr>
<td>TL E X I H</td>
<td>9</td>
<td>1.13</td>
<td></td>
</tr>
<tr>
<td>TIS/LEXH</td>
<td>384</td>
<td>(37.28)</td>
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Figure 1. Probability of predicting the more frequent of two events over trial blocks for all experimental groups.
The graph for the .60:.70 group demonstrates the strength of the within-Ss incentive effect. In this group, 1-chip trials had a $\pi_L$ equal to .70, and 10-chip trials had a $\pi_H$ equal to .60. Thus, in spite of the fact that event frequencies were fixed in such a way as to oppose an incentive effect, $P(A_1)$ was higher on the 10-chip trials throughout the course of training. The data from a condition even more unfavorable to observing an incentive effect are shown in the graph for the .60:.80 group. Even though $\pi_L$ exceeded $\pi_H$ by 20%, the results show that $P(A_1)$ on 10-chip trials was slightly greater than $P(A_1)$ on 1-chip trials for 200 trials, after which the incentive effect was reversed.

The lower half of Fig. 1 displays the results for the groups in which $\pi_H$ was .70 and $\pi_L$ was either .50, .60, .70, or .80. The very large effect of incentive in the .70:.50 group was to be expected since it was favored by both $\pi$ and incentive, but the fact that $P(A_1)$ on 1-chip trials was less than 50% is of interest since it illustrates the depressed performance levels on low risk trials that has previously been observed in within-Ss incentive work (Schnorr et al., Lipkin et al., 1965). In the .70:.60 group, the incentive effect persisted as would be expected since $\pi_H$ was still 10% greater than $\pi_L$. However, the incentive difference displayed by the .70:.70 group is worth note in view of the Myers et al. (1963) study which showed no difference between
10% and 1% \( P(A_1) \) levels when incentive was varied between-Ss with \( \pi \) equal to .70. Again, the strength of the within-Ss incentive effect is revealed in the data of the .70:.80 group. Here \( \pi_L \) exceeded \( \pi_H \) by 10% but an early incentive difference persisted through 300 trials. It is interesting to note the similarities between the data for the .70:.80 and .60:.80 conditions. Apparently, the incentive effect makes itself felt very early in training and then may be reversed with further training if \( \pi_L \) is greater than \( \pi_H \).

Figure 2 shows the data points for the fourth block of trials from each of the preceding figures. Aside from the groups with \( \pi_L \) of .50, \( P(A_1) \) levels on 10-chip trials were well within the range of what would be expected on the basis of between-Ss incentive work (Myers et al., 1963; Schnorr et al.). On the other hand, \( P(A_1) \) on the 1-chip trials was depressed downward relative to the between-Ss data, a finding which replicates the negative contrast effect observed by Schnorr et al. However, the increments in \( P(A_1) \) on 1-chip trials following increases in \( \pi_L \) dispels the suggestion by Schnorr et al. that the within-Ss incentive effect may be due to Ss "ignoring" low risk trials.

Figure 2 also reveals evidence of generalization in that \( P(A_1) \) for 10-chip trials was affected by the value of \( \pi_L \), and \( P(A_1) \) for 1-chip trials was affected by the value of \( \pi_H \). For instance, the \( P(A_1) \) curve for 10-chip trials
Figure 2. Probability of predicting the more frequent of two events over trials 301-400 as a function of $\pi_H$, $\pi_L$, and incentive.
with $\pi_H$ equal to .70 tended to increase as a function of $\pi_L$, and the curve for 10-chip trials with $\pi_H$ equal to .60 appeared to be a more complex function of $\pi_L$. Similarly, $P(A_1)$ for 1-chip trials appears to have been affected by the value of $\pi_H$. The two $P(A_1)$ functions for 1-chip trials in Fig. 2 are based on groups with identical levels of risk and $\pi_L$; however, $P(A_1)$ on 1-chip trials was higher for the groups with $\pi_H$ equal to .70. Whereas Fig. 2 shows only the fourth block of trials, Figs. 3 and 4 illustrate this generalization over all trials. In Fig. 3, the positive slope of the 10-chip function follows the change of $\pi_H$, but the positive slope of the 1-chip function is further evidence of generalization since $\pi_L$ is constant. This Incentive x $\pi_L$ interaction was statistically significant, $F(1,128) = 14.14$, $p<.001$, and the positive slope of the 1-chip function approaches significance, $F(1,256) = 2.91$, $p<.10$. Additional confirmation of generalization is provided by the data in Fig. 4. In this figure, the positive slope of the 1-chip function is not surprising; however, the significant nonzero slope of the 10-chip function reflects the interdependence of $P(A_1)$ for the two incentive levels, $F(3,256) = 7.68$, $p<.001$. Thus, it seems safe to conclude that $P(A_1)$ for 10-chip trials was a function of $\pi_L$, and $P(A_1)$ for 1-chip trials also tended to be directly related to $\pi_H$.

Figure 4 is of additional interest in relation to the
Figure 3. Per cent A1 choices over all trials as a function of incentive.
Figure 4. Per cent $A_1$ choices over all trials as a function of $\pi_L$ and incentive.
findings of Popper and Atkinson (1958) and Atkinson et al. (1959). The present study shows $P(A_1)$ for 10-chip trials to be an increasing function of the value of $\pi_L$ just as the two earlier discrimination studies found responses to one cue directly related to the reliability of a second cue.

In review, marginal response probabilities showed strong within-Ss incentive effects whenever $\pi_H$ was greater than or equal to $\pi_L$. Furthermore, the incentive effect occurred early in training and tended to persist even when $\pi_L$ exceeded $\pi_H$. This within-Ss incentive effect was in the form of depressed $P(A_1)$ under the lower incentive compared with that observed in between-Ss work. Generalization was also found since $P(A_1)$ for 10-chip trials was a function of $\pi_L$, and $P(A_1)$ for 1-chip trials tended to be a function of $\pi_H$.

From these findings with marginal response probabilities, it is clear that the pattern model is not compatible with the present data. The pattern model does not predict the observed interdependence of $P(A_1)$ levels for 10-chip and 1-chip trials. Since this model and the component model both predict an upper bound of $\pi_H$ for 10-chip performance and a lower bound of $\pi_L$ for 1-chip performance, they fail to account for the incentive effect which typically leads to overshooting of $\pi_H$ and sometimes to
depressions of 1-chip performance below $\pi_L$. While the Myers and Atkinson (1964) weak-strong model, a modification of the pattern model, can predict the overshooting of $\pi_H$, it cannot predict undershooting of $\pi_L$ nor can it handle the observed changes in performance under each incentive level as a function of the $\pi$ value of the second incentive. The Lee (1966) and Atkinson (1958) models which can account for generalization are not applicable to incentive data for the same reasons the pattern and component models are not.

**Subjective reports.** Figure 5 shows the subjective estimates of $\pi_H$ and $\pi_L$ for all eight groups. These estimates were obtained immediately following the fourth trial block. Analysis of variance shows subjective estimates to be directly related to $\pi_H$, $F(1,152) = 9.29, p<.005$, $\pi_L$, $F(3,152) = 6.75, p<.001$, and incentive, $F(1,151) = 33.81, p<.001$. In general, Ss tended to perform at levels above their estimates of $\pi$ as is revealed in a comparison of Figures 2 and 5. A comparison of the 1-chip curves for $\pi_S$ and $P(A_1)$ data shows that while $P(A_1)$ on 1-chip trials was shifted upward by a higher $\pi_H$, $\pi_S$ 1-chip data were about the same for the two levels of $\pi_H$. In other words, Ss gave similar estimates of $\pi_L$ after having performed quite differently on 1-chip trials as a function of $\pi_H$. This finding implies that an S's actual choice performance is not solely dependent on his perception of relative event frequencies.
Figure 5. Subjective estimates of $\pi_H$ and $\pi_L$ for all experimental groups.
The finding that the 1-chip $n_S$ function had a positive slope, supports the conclusion, mentioned above, that the within-SSs incentive effect is not merely the result of SSs "ignoring" low risk trials. The data are more compatible with the notion that the within-SSs procedure caused downward distortions in $n_S$ on low risk trials. While this lowered $n_S$ was related to lower $P(A_1)$, the relationship was not simply one-to-one as was pointed out above. As Fig. 5 shows, there was an interaction in SSs' perception of $n_H$ and $n_L$ such that $n_S$ undershoot $n_L$ and overshoot $n_H$. In order to reach a more complete description of the relation between $n_S$ and $P(A_1)$, it would have been useful to obtain subjective estimates at several points in training.

Subjects were also asked to estimate the proportion of trials worth 10 chips. As Table 2 shows, these estimates were quite accurate with a very slight underestimation for the groups with $n_H$ of .60 and an even slighter overestimation for the groups with $n_H$ equal to .70.

Conditional statistics. The first-order conditional statistics from the fourth trial block are shown in Table 3. Each of the 16 statistics for the 8 groups of SSs gives the probability of an $A_1$ response on trial $n$, given the incentive level on trial $n$ and the incentive-response-event combination on trial $n-1$. Table 3 shows these statistics subdivided in terms of the four possible incentive combinations.
Table 2

Average Subjective Estimates for Per Cent 10-Chip Trials, \( \pi_H \) and \( \pi_L \).

<table>
<thead>
<tr>
<th>( \pi_L )</th>
<th>( \pi = .60 )</th>
<th>( \pi = .70 )</th>
</tr>
</thead>
<tbody>
<tr>
<td>% 10-chip Trials</td>
<td>( \pi_H )</td>
<td>( \pi_L )</td>
</tr>
<tr>
<td>.5</td>
<td>48.0</td>
<td>64.5</td>
</tr>
<tr>
<td>.6</td>
<td>51.0</td>
<td>65.5</td>
</tr>
<tr>
<td>.7</td>
<td>48.5</td>
<td>65.0</td>
</tr>
<tr>
<td>.8</td>
<td>45.0</td>
<td>68.0</td>
</tr>
</tbody>
</table>

\( \bar{X} = 48.1 \) \( \bar{X} = 50.5 \)
from trial \( n-1 \) to trial \( n \).

The rank order of the conditionals is of prime interest in order to assess the appropriateness of several mathematical models. According to these models, the first statistic in each set of four must be the largest and the fourth must be the smallest while the ranks of the second and third depend on the values of the parameters of the models (i.e.,

\[
P(A_1,n \mid A_1,n-1,E_1,n-1) > P(A_1,n \mid A_1,n-1,E_2,n-1) \text{ and } P(A_1,n \mid A_2,n-1,E_1,n-1) > P(A_1,n \mid A_2,n-1,E_2,n-1)\]

). A brief presentation of the predicted conditional statistics from the component model (Burke and Estes, 1957), the pattern model (Estes, 1959), the conditioning-parameter model (Lee, 1966), and the observing response model (Atkinson, 1958) is given below to demonstrate the rank order of these expressions.

In the component model, there is a hypothetical set of stimulus elements for each incentive level and an intersection of the sets comprising common elements and background cues. On each trial \( S \) samples stimulus elements from both the unique set associated with the relevant incentive and from the common set. The probability of either the \( A_1 \) or \( A_2 \) response is determined by the proportion of sampled stimulus elements conditioned to it. The model calls for an increment in \( P(A_1) \) following an \( E_1 \) and a decrement following an \( E_2 \) according to two specific linear operators. For simplicity, the expressions given below are for the case in
Table 3

Asymptotic First-Order Conditional Statistics

| Trials 301 - 400 | \( P(A_{1,n}|S_j,n S_{k,n-1} A_{1,n-1} E_{m,n-1}) \) |
|------------------|----------------------------------|
|                  | .6:.5   | .6:.6   | .6:.7   | .6:.8   | .7:.5   | .7:.6   | .7:.7   | .7:.8   |
| \( S_j,n \)      | \( S_{k,n-1} \) | \( A_{1,n-1} \) | \( E_{m,n-1} \) | .6:.5   | .6:.6   | .6:.7   | .6:.8   | .7:.5   | .7:.6   | .7:.7   | .7:.8   |
| 10               | 10      | 1       | 1       | .7273   | .8018   | .7925   | .7156   | .8625   | .8270   | .9053   | .8834   |
| 10               | 10      | 1       | 2       | .7500   | .7972   | .8065   | .8120   | .8319   | .8992   | .8720   | .8783   |
| 10               | 10      | 2       | 1       | .3333   | .6026   | .4865   | .5750   | .5846   | .5660   | .7925   | .8780   |
| 10               | 10      | 2       | 2       | .4730   | .6000   | .5102   | .5000   | .4286   | .7727   | .5556   | .8696   |
| 1                | 10      | 1       | 1       | .3559   | .5111   | .6073   | .7837   | .3617   | .6399   | .7713   | .8328   |
| 1                | 10      | 1       | 2       | .4758   | .5409   | .6111   | .7643   | .4046   | .6889   | .7787   | .9247   |
| 1                | 10      | 2       | 1       | .4545   | .5926   | .7629   | .8837   | .6418   | .6667   | .8000   | .9149   |
| 1                | 10      | 2       | 2       | .6835   | .6471   | .6400   | .7879   | .6786   | .8929   | .7273   | 1.0000  |
| 1                | 1       | 1       | 1       | .5691   | .7716   | .8217   | .7654   | .8425   | .8815   | .8797   | .8684   |
| 1                | 1       | 1       | 2       | .6964   | .7383   | .6277   | .5672   | .7905   | .8933   | .8790   | .8864   |
| 1                | 1       | 2       | 1       | .7049   | .7095   | .7440   | .6484   | .8385   | .8851   | .8090   | .7959   |
| 1                | 1       | 2       | 2       | .6043   | .6737   | .5645   | .8571   | .8803   | .8429   | .6452   | .6000   |
| 1                | 1       | 1       | 1       | .5556   | .6335   | .7444   | .8419   | .5413   | .8084   | .8226   | .8625   |
| 1                | 1       | 1       | 2       | .6667   | .6780   | .7576   | .8261   | .6814   | .8934   | .7545   | .9012   |
| 1                | 1       | 2       | 1       | .3404   | .4016   | .4919   | .5556   | .3582   | .4713   | .6750   | .7167   |
| 1                | 1       | 2       | 2       | .3623   | .4231   | .4419   | .6111   | .4118   | .5185   | .4848   | 1.0000  |
which the incentive level is the same on trials \( n \) and \( n-1 \).

\[
P(A_1,n|A_1,n-1E_1,n-1) = (1-\theta)a + \theta
\]

\[
P(A_1,n|A_1,n-1E_2,n-1) = (1-\theta)a
\]

\[
P(A_1,n|A_2,n-1E_1,n-1) = (1-\theta)b + \theta
\]

\[
P(A_1,n|A_2,n-1E_2,n-1) = (1-\theta)b
\]

where  
\[
a = \frac{\alpha_{2,n-1}}{\alpha_{1,n-1}}, \quad b = \frac{\alpha_{1,n-1} - \alpha_{2,n-1}}{1-\alpha_{1,n-1}}
\]

\( a_i \) = \( i \)th raw moment of the distribution of response probabilities over Ss.

Clearly, the first statistic must be larger than the second, and the third must be larger than the fourth regardless of the parameter values.

In a pattern model with each of \( N \) stimulus patterns conditioned to \( A_1 \) or \( A_2 \), it is assumed that one pattern is sampled on a trial. The conditioning state of the sampled pattern determines the response; and if the response is correct, the pattern's conditioning state remains unchanged. However, if the response is incorrect, there is a probability \( c \) that the pattern becomes conditioned to the event presented on that trial. Again, the following theoretical expressions are for two consecutive trials with the same incentive level.

\[
P(A_1,n|A_1,n-1E_1,n-1) = \frac{N-1}{N}\pi + \frac{1}{N}
\]

\[
P(A_1,n|A_1,n-1E_2,n-1) = \frac{N-1}{N}\pi + \frac{1-c}{N}
\]
\[ P(A_1,n | A_2,n-1 E_1,n-1) = \frac{N-1}{N} \pi + \frac{c}{N} \]

\[ P(A_1,n | A_2,n-1 E_2,n-1) = \frac{N-1}{N} \pi \]

Inspection shows that the first statistic must be the largest and the fourth, the smallest while the ranks of the second and third depend on the value of c.

Lee's conditioning-parameter model is very similar to the pattern model with the exception that there is an additional parameter to allow for generalization. However, this generalization does not enter into the case of two consecutive trials with identical incentives. For this case, the theoretical expressions for the conditional statistics are of the same form as those from the pattern model.

The derivation of theoretical expressions of conditional statistics from the Atkinson observing response model is much more complex. The complications result from the assumptions that there are 16 possible conditioning states, an observing response preceding the actual A_1 or A_2 choice response, and the usual conditioning parameters. While it is beyond the scope of this study to present the complete proof that the rank order of the conditionals is identical to that of the preceding models, a proof that \( P(A_1,n | A_1,n-1 E_1,n-1) > P(A_1,n | A_1,n-1 E_2,n-1) \) is given in the Appendix.

While the preceding models concur on the rank order of the predicted conditional statistics, the observed first-order
conditional statistics shown in Table 3 did not show this rank ordering. For instance, when the incentive level was 10 chips on trials \( n \) and \( n-1 \), \( P(A_1,n\mid A_1,n-1E_2,n-1) \) exceeded \( P(A_1,n\mid A_1,n-1E_1,n-1) \) in 4 groups, and \( P(A_1,n\mid A_2,n-1E_2,n-1) \) exceeded \( P(A_1,n\mid A_2,n-1E_1,n-1) \) in 3 groups. When the incentive level was 10 chips on trial \( n-1 \) and 1 chip on trial \( n \), \( P(A_1,n\mid A_1,n-1E_1,n-1) \) was the smallest of the 4 conditional statistics in 6 groups and third in rank in the other two groups. In other words, \( P(A_1) \) increased less following a direct reinforcement than after any other response-event combination. Further inspection of Table 3 reveals a variety of inversions in the ranks of the conditionals when the incentive level was 1 chip on trial \( n-1 \) and 10 chips on trial \( n \). Finally, the conditionals based on two consecutive 1-chip trials show that \( P(A_1,n\mid A_1,n-1E_2,n-1) > P(A_1,n\mid A_1,n-1E_1,n-1) \) and \( P(A_1,n\mid A_2,n-1E_2,n-1) > P(A_1,n\mid A_2,n-1E_1,n-1) \) each in 6 of the 8 groups. In short, under very few conditions were the conditional statistics in the present study found to be in a rank order predicted by available mathematical models.

The fact that in several cases the increase in \( P(A_1) \) following an error was greater than that following a correct \( A_1 \) prediction apparently contradicts reinforcement theory. However, this contradiction may be the result of rather narrow interpretations of the probability learning situation by the above models. These models share the assumption that...
reinforcement depends only upon the immediately preceding event. For instance, Burke and Estes (1957) call for the application of a linear operator if an $E_1$ occurs regardless of what has happened on earlier trials. Similarly, in the Estes (1959) and Atkinson (1958) models, a Markovian interpretation leads to predictions of one-step dependencies in conditioning states. As Anderson (1964) points out, such theories neglect the effects of certain rather potent variables such as memory beyond one trial, patterns arising in the sequencing of events, and transfer from pre-experimental decision situations. Despite these neglected variables, Myers and Atkinson (1964) show that the conditional statistics from the Myers et al. (1963) between-Ss incentive study are in the predicted order.

The methodology of the present study was similar to that of the Myers et al. (1963) study with the obvious exception of the present within-Ss manipulation of incentive magnitude. This within-Ss procedure used pre-trial cues thereby introducing additional structure into the sequence of events. Moore and Halpern (1966) also reported inversions in the conditional statistics from a probability discrimination study which by its nature necessitated the use of pre-trial cues. It is plausible that the structure added to the sequence of events by pre-trial cues causes Ss to make choices on the basis of small blocks of trials and that this, in turn, may tend to reinforce negative recency.
SUMMARY

Each of 160 Ss received 400 two-choice trials which consisted of a random sequence of 200 10-chip and 200 1-chip risks. The Ss were divided into 8 groups formed by crossing two levels of $\pi_H$ (the probability of the more frequent event—$E_1$—under high risk) with 4 levels of $\pi_L$ (the probability of $E_1$ under low risk). A strong overall incentive effect was found, and the probability of choosing the more frequent event ($P(A_1)$) continued to be greater on 10-chip trials than on 1-chip trials even for groups in which $\pi_L$ exceeded $\pi_H$. Observed $P(A_1)$ under each incentive level was related to $\pi$ under the second incentive level. Subjective estimates of $\pi_H$ and $\pi_L$ also showed an incentive effect with overestimation of $\pi_H$ and underestimation of $\pi_L$. These findings and observed first-order conditional statistics were discussed in relation to mathematical models.
REFERENCES


APPENDIX A

INSTRUCTIONS

We are studying how people make decisions. In this experiment, we will be using this deck of cards. On the back of each card there is drawn one of two geometric designs—a point or a line. Your task is to try to predict which design is on the back of each card. Your only clue will be your experience in the situation.

For each card you will make your prediction, and then I will turn the card over, showing you the design. If you are correct, you win the amount written on the top of the card—either 1 chip or 10 chips. If you are wrong, you lose this same amount.

Before you is a stake of 200 chips with which to begin the game. You will keep score with these chips. If you make a correct prediction, take the appropriate number of chips from the bank. If you make an incorrect prediction, put the appropriate number of chips into the bank.

In this game you can win often and add considerably to the chips you already have. But it is impossible for you to win on every prediction since there is no pattern in this deck of cards. So your task is to be correct as often as you can over the long run. I have no control over the sequence of events, and I will be recording it as we go along.
Before we start, I will shuffle the cards to ensure that they are in a strictly random order.

Are there any questions? Now let's begin with your prediction for the first card. Remember, the two choices are point and line. Try to win as many chips as you can during the session.
APPENDIX B

The following is a proof that for two consecutive trials with the same incentive level, \( P(A_1,n|A_1,n-1 E_1,n-1) > P(A_1,n|A_1,n-1 E_2,n-1) \) in the Atkinson (1958) observing response model. This proof does not involve a full description of the model; the interested reader should refer to the original paper.

Definition of symbols:

\[ u_i = \text{the probability of being in state } i \text{ on trial } n-1. \text{ The trial subscript will be omitted for convenience.} \]

\[ T_i = \text{trial of type } i \text{ where } i \text{ designates our incentive level.} \]

\[ \beta = \text{the probability of a } T_i \text{ trial.} \]

\[ \pi_1 = P(E_1|T_i) \]

\[ 1 - \pi_1 = P(E_2|T_i) \]

\[ w_1 = \frac{N_1}{N_1 + N_c} \]

\[ N_1 = \text{number of stimulus elements in the unique } S_1 \text{ set.} \]

\[ N_c = \text{number of stimulus elements in the common set.} \]

\[ \omega = \text{the conditioning parameter.} \]
APPENDIX B (cont'd.)

Proof: \( P(A_1, n | T_1, n \ T_1, n-1 \ A_1, n-1 \ E_1, n-1) > P(A_1, n | T_1, n \ T_1, n-1 \ A_1, n-1 \ E_2, n-1) \)

\[ P(A_1 | T_1 \ T_1 \ A_1 \ E_1) = \frac{P(A_1 \ T_1 \ T_1 \ A_1 \ E_1)}{P(T_1 \ T_1 \ A_1 \ E_1)} \]

\[ = \pi \beta^2 \left\{ \sum_{i=1}^{4} u_i + u_9 + u_{10} + (u_{11} + u_{12})w_1\sqrt{\theta} + (1-\theta)w_1 - \sqrt{\theta} + (u_{13} + u_{14})(1-w_1)\sqrt{\theta} + (1-\theta)(1-w_1) \right\} \]

\[ = \pi \beta^2 \left\{ \sum_{i=1}^{4} u_i + u_9 + u_{10} + w_1(u_{11} + u_{12}) + (1-w_1)(u_{13} + u_{14}) \right\} \]

\[ = \frac{\sum_{i=1}^{4} u_i + u_9 + u_{10} + (u_{11} + u_{12})w_1\sqrt{\theta} + (1-\theta)w_1 - \sqrt{\theta} + (u_{13} + u_{14})(1-w_1)\sqrt{\theta} + (1-\theta)(1-w_1) \right\} \]

\[ = \frac{\sum_{i=1}^{4} u_i + u_9 + u_{10} + w_1(u_{11} + u_{12}) + (1-w_1)(u_{13} + u_{14})}{\sum_{i=1}^{4} u_i + u_9 + u_{10} + w_1(u_{11} + u_{12}) + (1-w_1)(u_{13} + u_{14})} \]
APPENDIX B (cont'd.)

\[
P(A_1 | T_1 T_1 A_1 E_2) = \frac{P(A_1 T_1 T_1 A_1 E_2)}{P(T_1 T_1 A_1 E_2)}
\]

\[
= (1 - \pi_1) \beta^2 \left\{ \frac{(u_1 + u_2)\Theta(1-w_1)+(1-\Theta)\sum (u_3 + u_4 + u_9 + u_{10})(1-\Theta) + w_1^2(u_{11} + u_{12})(1-\Theta)+(u_{13} + u_{14})(1-w_1)^2(1-\Theta)}{(1-\pi_1) \beta^2 \left( \sum_{i=1}^{4} u_1 + u_9 + u_{10} + w_1 (u_{11} + u_{12}) + (1-w_1) (u_{13} + u_{14}) \right) \right\}
\]

\[
= \frac{\sum_{i=1}^{4} u_i - \Theta w_1(u_1 + u_2) - \Theta (u_3 + u_4) + (u_9 + u_{10}) (1-\Theta) + w_1^2 (u_{11} + u_{12}) (1-\Theta) + (u_{13} + u_{14}) (1-w_1)^2(1-\Theta)}{\sum_{i=1}^{4} u_i + u_9 + u_{10} + w_1 (u_{11} + u_{12}) + (1-w_1) (u_{13} + u_{14})}
\]

Note that the denominators of the two terms are equal.
APPENDIX B (cont'd.)

If $P(A_1|T_1 T_1 A_1 E_1) > P(A_1|T_1 T_1 A_1 E_2)$:

$$\sum_{i=1}^{4} \left[ u_i + u_9 + u_{10} + (u_{11} + u_{12}) \right] w_1 \sqrt{\theta} + (1-\theta) w_1 \sqrt{\theta} + (u_{13} + u_{14}) \left( (1-w_1) \sqrt{\theta} + (1-\theta) (1-w_1) \right) >$$

$$\sum_{i=1}^{4} \left[ u_i - \theta w_1 (u_1 + u_2) - \theta (u_3 + u_4) + (u_9 + u_{10}) (1-\theta) + w_1^2 (u_{11} + u_{12}) (1-\theta) + (u_{13} + u_{14}) (1-w_1)^2 (1-\theta) \right]$$

let $U = \sum_{i=1}^{4} u_i + u_9 + u_{10}$

$$U + (u_{11} + u_{12}) \left( w_1 \sqrt{\theta} + (1-\theta) w_1 \sqrt{\theta} \right) + (u_{13} + u_{14}) \left( (1-w_1) \sqrt{\theta} + (1-\theta) (1-w_1) \right) >$$

$$U - \theta w_1 (u_1 + u_2) - \theta (u_3 + u_4 + u_9 + u_{10}) + w_1^2 (u_{11} + u_{12}) (1-\theta) + (u_{13} + u_{14}) (1-w_1)^2 (1-\theta)$$

For simplicity we may subtract U from each side of the inequality.

$$(u_{11} + u_{12}) \left( w_1 \sqrt{\theta} + (1-\theta) w_1 \sqrt{\theta} \right) + (u_{13} + u_{14}) \left( (1-w_1) \sqrt{\theta} + (1-\theta) (1-w_1) \right) >$$

$$-\theta w_1 (u_1 + u_2) - \theta (u_3 + u_4 + u_9 + u_{10}) + w_1^2 (u_{11} + u_{12}) (1-\theta) + (u_{13} + u_{14}) (1-w_1)^2 (1-\theta)$$
APPENDIX B (cont'd.)

The inequality will be true if the first and second terms on the left exceed the third and fourth terms on the right, respectively, since two remaining terms on the right are negative.

The proof reduces to determining whether
\[
(u_{11} + u_{12}) w_1 \sqrt{\theta} + (1-\theta) w_1 \gamma > w_1^2 (u_{11} + u_{12}) (1-\theta)
\]
and
\[
(u_{13} + u_{14}) (1-w_1) \sqrt{\theta} + (1-\theta) (1-w_1) \gamma > (u_{13} + u_{14}) (1-w_1)^2 (1-\theta)
\]

Expanding the first inequality:
\[
(u_{11} + u_{12}) w_1 \theta + (u_{11} + u_{12}) w_1^2 (1-\theta) > (u_{11} + u_{12}) w_1^2 (1-\theta)
\]
Inspection shows this inequality to be true since \((u_{11} + u_{12}) w_1^2 \theta > 0\)

Expanding the second inequality:
\[
(u_{13} + u_{14}) (1-w_1) \theta + (u_{13} + u_{14})(1-\theta)(1-w_1)^2 > (u_{13} + u_{14})(1-w_1)^2(1-\theta)
\]
Inspection shows this inequality true since \((u_{13} + u_{14}) (1-w_1) \theta > 0\)

It necessarily follows that:
\[
P(A_1, n \ T_1, n \ T_1, n-1 \ A_1, n-1 \ E_1, n-1) > P(A_1, n \ T_1, n \ T_1, n-1 \ A_1, n-1 \ E_2, n-1)
\]
## APPENDIX C

Preasymptotic First-Order Conditional Statistics

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### APPENDIX C (cont'd.)

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**APPENDIX D**

**F-ratios from Analysis of Variance of Subjective Reports (\( \pi_s \))**

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APPENDIX E

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(.60-.50 Group)

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