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SEQUENTIAL DETERMINANTS OF INFORMATION PROCESSING IN SERIAL AND DISCRETE CHOICE REACTION TIME[1]

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It is shown that the measure of average stimulus information (H) is confounded with the probability of nonrepetition of the stimuli in most of the experimental conditions whose results have been taken as evidence in support of the linear relationship between choice reaction time (RT) and H. The results of a serial and a discrete experiment, so designed as to unconfound these two variables, lead to a rejection of the information hypothesis. The RT for repetitions is found to be faster than for nonrepetitions, and both are decreasing linear functions of their respective conditional probabilities. Some of the discussion focuses on the manner in which the slope and the intercepts of these linear functions are affected by changes in the number of alternatives, stimulus-response compatibility, and response-to-stimulus interval. It is also shown that the present approach not only accounts for data which had previously been described by the information hypothesis, but for results which departed from the hypothesis as well. Errors are discussed in a manner which supplements the main argument. Finally, it is shown that the molar results of RT experiments can be systematically accounted for in terms of the characteristics of easily distinguishable differentially sensitive partitions in these data.

Some of the results of a serial choice reaction time (RT) experiment were recently summarized in a brief report in which it was argued that the experiment constituted a critical test of the Information Hypothesis (Kornblum, 1968). The present paper has the following aims: (*a*) to present the results of that experiment and the arguments leading up to it in more complete detail, (*b*) to report the results of a discrete choice RT experiment whose findings extend the scope of the original conclusion, (*c*) to discuss some of the problems and implications that follow from these experiments, and (*d*) to provide a bridge between the approach that these findings suggest and the information theoretic and other molar approaches to RT problems.

The initial suggestion that choice RT might be linearly related to stimulus information was made by Miller (1951, pp. 205–206); Hick (1952), Hyman (1953), and Crossman (1953) verified this conjecture, as did many other investigators under a variety of experimental conditions. Welford's (1960) and Smith's (1968) reviews make it abundantly clear that implicit in all these studies is the "information hypothesis" whose essential elements are expressed in the following formulation:

> All other things being equal, equiinformation conditions give rise to equal overall mean RTs; that is, RT is a function of average stimulus information (H).

The more generally accepted form of the hypothesis is far more restricted in that the precise functional relationship is specified as: $RT = a + bH$. The latter implies the minimal statement that all other things being equal, there is a one-to-one correspondence between transmitted information and the overall mean RT.

The information hypothesis was one of the most influential propositions in the area of choice RT. One of its most valu-

[1] Some of the results in this paper were presented at the Donders Centennial Symposium on Reaction Time held in August 1968 in the Netherlands, under the auspices of the Institute for Perceptual Research in Eindhoven. Portions of this paper were also presented at the APA Symposium on "Inhibition/Facilitation Aspects of Information Processing" held in San Francisco in September 1968.

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able outcomes was the identification of a set of variables which subsequently had to be specified and explicitly included among "all other things being equal" (e.g., stimulus-response (S-R) compatibility, S-R mapping, stimulus discriminability, training, error rate, etc.). These variables all dealt with different aspects of the experimental conditions and the qualifications that had to be placed on them. However, the probability structure of the stimulus sequences has remained unrestricted except for the unheeded caveats and observations in Hyman's (1953) classic paper, and Bertelson's (1961) more recent study. Hence, in spite of the relatively large experimental and theoretical effort that the hypothesis stimulated, and the precautionary statements that are found in Bricker (1956) and Leonard (1961), the hypothesis itself has not been subjected to a direct, critical examination.

In an investigation on the effects of sequential redundancy in a two-choice task, Bertelson (1961) found that the overall mean RT in two equiinformation conditions differed depending on the manner in which the redundancy was presented. If we define the overall probability of nonrepetitions as:

$$p_{nr} = \sum_{j} p(j) \sum_{i \neq j} p(i \mid j) \qquad [1]$$

where:

$p(j)$ = the absolute probability of occurrence of Stimulus j

$p(i \mid j)$ = the probability of Stimulus i occurring on Trial n, given that Stimulus j has occurred on Trial $n-1$,

and the overall probability of repetitions as $p_r = 1 - p_{nr}$, Bertelson's results indicated that the condition with the higher probability of repetitions ($p_r = .75$) had a faster mean RT than the condition with a higher probability of nonrepetitions ($p_{nr} = .75$); the signals themselves were equiprobable in both sequences, and the average stimulus information was the same for both sequences ($H = .81$). Bertelson accounted for his results in terms of a facilitating process associated with repeated events. His findings also confirmed Hyman's original observations that "(for more than two alternatives) whenever a stimulus was immediately followed by itself in the series, S [the subject] seemed to respond unusually fast to it [pp. 194–195]."

Considering the above findings it would be tempting to conjecture that the overall mean RT ($\overline{RT}$) is solely dependent on the proportion of repetitions and nonrepetitions in the data. One possible set of conditions under which this might be the case is the following: Suppose that the mean RT for repetitions ($\overline{RT}_r$) and the mean RT for nonrepetitions ($\overline{RT}_{nr}$) are ordered such that $\overline{RT}_r < \overline{RT}_{nr}$. Suppose, further, that the values of $\overline{RT}_r$ and $\overline{RT}_{nr}$ are constant for different values of p_{nr}. Since

$$\overline{RT} = p_{nr}\overline{RT}_{nr} + p_r\overline{RT}_r \qquad [2]$$

it follows that $\overline{RT}$ would be a monotone increasing function of p_{nr}. This simple conjecture has to be firmly rejected on the basis of experimental evidence (Kornblum, 1967) which indicates that while p_{nr} may be an important determinant of $\overline{RT}$, the number of alternatives has an effect which is over and above the effects attributable to p_{nr}; these same data also indicate that $\overline{RT}_r$ and $\overline{RT}_{nr}$ are not constant. However, the results do verify the ordering $\overline{RT}_r < \overline{RT}_{nr}$ within sets of conditions having the same number of alternatives. Even though $\overline{RT}_r$ and $\overline{RT}_{nr}$ are not constant they may, nevertheless behave lawfully; since they also constitute a partition of $\overline{RT}$ it may still be instructive to examine them in some detail. Such an examination may have additional importance when the extent of the confounding between H and p_{nr} is made apparent.

THREE METHODS OF VARYING H

Three of the most widely used ways of varying H consist of either varying the number of alternatives, the absolute probabilities of the stimuli, or the one-step sequential dependencies. One possible test of the information hypothesis might, therefore, be to compare the regression of $\overline{RT}$ on

H when H is varied in these three different ways *one at a time*. (It is important to note that whereas Hyman did vary H in these three ways, the data from the three experiments to which he fitted his regression lines all include different numbers of alternatives as a variable.) However, it can be shown that these three ways of varying H lead to H and p_{nr} being confounded. Hence, a test of the information hypothesis would have to take this confounding into account as well.

1. Varying the number of alternatives (K), where the alternatives occur equiprobably and independently.

The value of p_{nr} in a sequence with K independent, equiprobable events is given by $p_{nr} = (K - 1)/K$. The value of H in such a sequence is given by $H = \log_2 K$. Since $K = 1/1 - p_{nr}$ it is obvious that as K increases not only does H increase but p_{nr} increases as well. Hence, this method leads to H, K, and p_{nr} being necessarily confounded.

2. Varying the absolute probabilities of occurrence of the stimuli where K is fixed and the probabilities of occurrence are independent.

For independent stimuli $p(i|j) = p(i)$ for all i and j, and the probability of repetitions is given by $p_r = \sum_i p^2(i)$. When the K alternatives occur equiprobably, the quantity p_r is a minimum which means that p_{nr} is a maximum. Since H is also a maximum in the equiprobable case it follows that the introduction of nonequiprobable signals will result in a simultaneous decrease in the values of both H and p_{nr}. Even though this particular method allows one to find different sets of $p(i)$ values which will generate equal values of H, with different values of p_{nr}, and vice versa, H and p_{nr} will in general be confounded with this procedure, particularly with small values of K.

3. Varying the one-step sequential dependencies where K is fixed, and the stimuli occur equiprobably.

One particular set of transition probabilities that will generate stimulus sequences in this third way is illustrated in the following matrix where $0 < p^* \leq 1$:

$$\begin{array}{cc} & \text{Stimulus on Trial } n \\ & \begin{array}{cccc} \quad 1 \quad & \quad 2 \quad & \cdots & \quad K \quad \end{array} \\ \begin{array}{c} \\ \text{Stimulus} \\ \text{on Trial} \\ n-1 \end{array} \begin{array}{c} 1 \\ 2 \\ \cdot \\ \cdot \\ K \end{array} & \begin{bmatrix} 1-p^* & \frac{p^*}{K-1} & \cdots & \frac{p^*}{K-1} \\ \frac{p^*}{K-1} & 1-p^* & \cdots & \frac{p^*}{K-1} \\ \cdots & \cdots & \cdots & \cdots \\ \cdots & \cdots & \cdots & \cdots \\ \frac{p^*}{K-1} & \frac{p^*}{K-1} & \cdots & 1-p^* \end{bmatrix} \end{array}$$

The cell entries in the matrix represent the probability $p(i|j)$ of presenting Stimulus i on Trial n, given that Stimulus j has been presented on Trial $n - 1$. The equality among all the diagonal elements (i.e., repetitions), and all the off-diagonal elements (i.e., nonrepetitions), makes this a special case of doubly stochastic matrices in which $p_{nr} = p^*$.

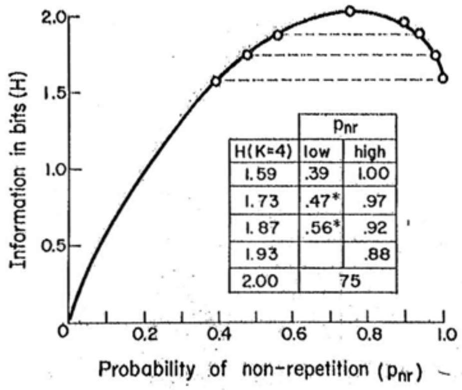
The value of H for sequences with one-step sequential dependencies is given by:

$$H = -\sum_j p(j) \sum_i p(i|j) \log p(i|j) \quad [3]$$

In the case of sequences that are generated within the constraints of the matrix above, H is a monotone increasing function of p^*, or p_{nr}, in the range of $0 < p^* \leq (K - 1)/K$ (corresponding to $0 < H \leq \log K$); however, for the range of values $(K - 1)/K \leq p^* \leq 1$ [corresponding to $\log K \geq H \geq \log(K - 1)$], H is a monotone decreasing function of p_{nr}. Hence in the range $\log(K - 1) \leq H \leq \log K$ there exist pairs of high and low values of p_{nr} which will generate pairs of sequences with identical values of H which in principle provides the necessary conditions to unconfound H and p_{nr}. This is the property that was exploited in designing the experiment that is described below.

A Critical Experiment

Figure 1 illustrates H as a function of p_{nr} for four-choice sequences that are generated by transition matrices of the type discussed in Section 3 above. The range within which the effects of H can, in principle, be



	p_{nr}	
H(K=4)	low	high
1.59	.39	1.00
1.73	.47*	.97
1.87	.56*	.92
1.93		.88
2.00	75	

FIG. 1. The curve represents the average stimulus information $[H = -\sum_j p_j \sum_i p(i|j) \log p(i|j)]$ as a function of the probability of nonrepetitions $[p_{nr} = \sum_j p_j \sum_{i \neq j} p(i|j)]$ for four-choice, equiprobable sequences. (The table insert represents the values of p_{nr} and H that were used in the serial experiments. The dotted lines connect equiinformation conditions. In the discrete experiment the p_{nr} values that were used were slightly different from those indicated in the asterisked cells (*); the values were $p_{nr} = .45$ and $.55$ with $H = 1.71$ and 1.84, respectively.)

pitted against the effects of p_{nr} include the upper 20% of H ($1.59 \leq H \leq 2.00$), and the upper 60% of p_{nr} ($.39 \leq p_{nr} \leq 1.00$).

Clearly, if the values of $\overline{RT}_r$ and $\overline{RT}_{nr}$ were identical, then the confounding between H and p_{nr} would be an interesting formal observation, at best. However, this observation could shed some light on the information hypothesis in view of the evidence indicating that in general a systematic difference is found between the RT for repetitions and nonrepetitions (i.e., $\overline{RT}_r < \overline{RT}_{nr}$). Such a difference is, therefore, a necessary experimental condition for unconfounding the effects of H and p_{nr}. Bertelson has shown that this difference is affected by at least two factors. First, the magnitude of this difference decreases as the length of the delay between a response and the next signal increases (Bertelson, 1961; Bertelson & Renkin, 1966). Second, the magnitude of the difference increases as the S-R compatibility of the task decreases; the latter seems primarily due to a much larger increase in the RT for nonrepetitions associated with a reduction in S-R compatibility (Bertelson, 1963). In terms of these two factors, therefore, a differential effect may be expected with a serial task in which the S-R compatibility is high, or with a discrete task in which the S-R compatibility is relatively lower. The expectation in the latter case is predicated on the following assumption: Given a short R-S interval, the greater the difference between $\overline{RT}_r$ and $\overline{RT}_{nr}$, the longer the R-S interval for which this difference will be sustained with the same sign. This is essentially the same assumption that Bertelson recently made in describing some of his own data (Bertelson & Renkin, 1966).

While for the purpose of unconfounding the effects of H and p_{nr}, either a serial or a discrete task might be sufficient, results which are based on serial data alone would be less than fully convincing, since most of the evidence supporting the information hypothesis is based on discrete tasks. The conditions illustrated in Figure 1 were, therefore, run with both a serial and a discrete task. In the former, the R-S interval was 140 milliseconds (msec.); in the latter the R-S interval was approximately 3 seconds.

Method

Stimuli-serial experiment. Four neon lights (signalite-neptune, New Jersey; RT2-32-1A) were mounted on a Masonite board behind ⅜ inch long lucite rods, ½ inch in diameter, whose surfaces had been uniformly buffed on all sides. The board was painted gray and placed vertically approximately 5.5 feet in front of a seated S with the lights approximately at eye level and forming a horizontal array of four circles. The two middle lights were slightly lower than the two outer ones and were separated by ½ inch; the lights of the left and right pairs were separated by ¼ inch. A thin white vertical line between the two middle lights served as a fixation line. In addition, the bottom edge of each light was underscored by a thin white line which made the alternatives clear at all times.

Stimuli-discrete experiment. The stimuli consisted of the digits 1 through 4, displayed on a standard IEE, CRT read-out tube, model BA-000-P31 (Industrial Electronic Engineers, Van Nuys, California). The tube had a diameter of 1.135 inch and the digits were ⅝ inch high. The tube was mounted on a Masonite board which was placed vertically approximately 5.5 feet in front of a seated S with the tube at eye level.

Responses and trial structure. The responses were made by depressing the appropriate one of four keys with the middle or index finger of the left or right hand. Between 5 and 6 ounces of pressure and $\frac{1}{8}$ inch travel were required to depress a key. The *S*s were instructed to keep their fingers resting on the keys during a sequence. The stimuli stayed on until a response had been made, and then went off simultaneously with a key press.

In the serial experiment, the extreme left light was responded to with the left middle finger, the next light with the left index finger, and so on. There was 140 msec. interval between a response and the presentation of the next stimulus.

In the discrete experiment, the digit 1 was responded to with the left middle finger, the digit 2 with the left index finger, and so on. A warning tone of 1000 cycles per second and 150 msec. duration was presented 750 msec. after the response. Following the warning tone, the next stimulus was presented after a randomly chosen interval between 2 and 2.5 seconds.

Procedure and Design

Eight different sets of sequences corresponding to the eight conditions illustrated in Figure 1 were used in each experiment. In the serial experiment a sequence consisted of approximately 300 trials; in the discrete experiment, this number was approximately 150 trials. Each *S* received all eight conditions on each day of the experiment in an order determined by a Latin-square technique as modified by Bradley (1958). With this modification, which requires an even number of conditions, each condition is preceded by a different condition in every row.

The first day in each experiment was used as a training day on which all *S*s received the full eight sequences in the same order (p_{nr}=.75, .39, 1.00, .47, .97, .56, .92, .88). In the serial experiment, the next 2 days were considered experimental days and different Latin squares were used for each day. In the discrete experiment, training was followed by 4 experimental days with different Latin squares for each day. The same two Latin squares were used in the last 2 days of the discrete experiment as were used in the serial experiment.

Throughout both experiments *S*s wore earphones through which they heard white noise at 73–75 decibels which masked any distracting noises that may have occurred. The *S*s were instructed to respond as quickly and as accurately as possible. At the end of each sequence they were told how well they had done. They were encouraged to ask for a rest any time they felt the need for it. The *S*s were told that the stimuli were always equiprobable, and before starting the next sequence they were also told what the probability of repetition and nonrepetition for that sequence would be; hence, *S*s were always completely informed with regard to the probabilistic properties of the sequence.

Results and Discussion

Three types of responses may be distinguished in such tasks: (*a*) "correct responses"—that is, correct responses preceded by correct responses, (*b*) "posterror responses"—that is, correct responses preceded by errors, and (*c*) "errors"—incorrect responses. The major emphasis will be placed on "correct responses"; results regarding the other two types of responses will be brought in as supplementary or ancillary information.

Figure 2 illustrates the overall mean *RT* for correct responses and the overall error rate for both experiments, as a function of stimulus information.[3] It can be seen that the *RT* for the high p_{nr} sequences in both experiments is longer than for the low p_{nr} sequences matched for stimulus information. The error rate in both experiments is between the 3% and 4% level. Except for a slightly higher error rate in the high p_{nr} sequences of the discrete experiment, the errors do not display any overall systematic differences between conditions; any decrement in transmitted information (H_t), or "rate of gain of information" (Hick, 1952), as a result of errors in performance is, therefore, approximately uniform over all eight conditions of both experiments.[4] The slope of *RT* with respect to *H* for the low p_{nr} points (which arbitrarily includes the condition with 2.00 bits) is 107.7 msec./bit for the serial experiment, and 96.1 msec./bit for the discrete experiment; the slope of the high p_{nr} points is 10.9 msec./bit for the serial experiment, and 44.5 msec./bit for the discrete experiment. The difference between the slopes of the high and low p_{nr}

[3] All the results are based on data from both experimental days of the serial experiment, and the last 2 experimental days of the discrete experiment.

[4] Equal decrements in transmitted information (H_t) for the eight experimental conditions obtain only if all the errors are weighted equally, as they would be in the ordinary stimulus-response transition matrix as illustrated in Hick (1952). However, when the sequential properties of the sequence are taken into account, sharp distinctions in error patterns emerge which bring such simple procedures into serious question (cf. later sections of this paper). It would perhaps have been more accurate to estimate H_t by explicitly treating the stimuli and responses, and the transitions between them as Markov sequences; this was not done. However, it is doubtful that such an analysis would have altered our conclusions.

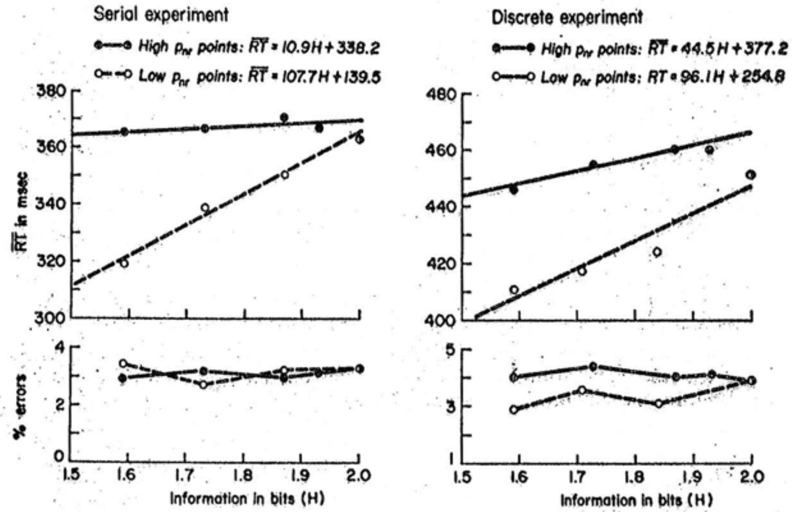


FIG. 2. Overall mean RT of correct responses, and overall error level, in both experiments, plotted against the average stimulus information in each condition.

points within each experiment is statistically significant (serial: $t = 7.3$, $df = 165$, $p < .01$; discrete: $t = 2.2$, $df = 148$, $p < .05$). The linear trend of the low p_{nr} points is statistically significant for both experiments (serial: $t = 12.3$, $df = 84$, $p < .01$; discrete: $t = 6.2$, $df = 84$, $p < .01$) whereas the linear trend for the high p_{nr} points is significant in the discrete case only (serial: $t = 1.09$, $df = 84$; discrete: $t = 2.5$, $df = 84$, $p < .05$).

If, by merely varying the probability of nonrepetitions, equiinformation conditions have indeed been constructed which retain the property of "all other things being equal," then these data must be interpreted as being at variance with the information hypothesis. An alternative interpretation might be entertained in which the principle of "all things being equal" is viewed as having been violated by the use of values of $p_{nr} > (K - 1)/K$. Such an interpretation would render the hypothesis immune to at least one procedure to unconfound H and p_{nr}, and at the same time reduce to a questionable level the remaining value that the hypothesis may have.

RT FOR REPETITIONS AND NONREPETITIONS

In the introduction it was argued that the questions raised by the confounding between H and p_{nr}, while possibly engaging as formal curiosities, are substantively inconsequential in the absence of a systematic experimental difference in the RT for repetitions and nonrepetitions. As the next step, therefore, the correct responses for repetitions and nonrepetitions were examined separately. As had previously been found in the two-choice case (Bertelson, 1961; Kornblum, 1967) the RT for repetitions were found to be inversely related to the overall probability of repetitions and nonrepetitions, respectively, with repetitions being faster than nonrepetitions.

Figure 3 illustrates the RT for repetitions and nonrepetitions as a function of their respective conditional probabilities in both experiments. While the data seem more variable in the discrete than in the serial case, a number of observations do stand out quite clearly: (*a*) the RT for repetitions and nonrepetitions are decreasing linear function of their respective conditional probabilities (linear trend-serial: repetitions $t = 15.4$, $df = 42$, $p < .01$, nonrepetitions $t = 10.0$, $df = 49$, $p < .01$; linear trend-discrete: repetitions $t = 8.48$, $df = 42$, $p < .01$, nonrepetitions $t = 3.9$, $df = 49$, $p < .01$); (*b*) the deviations from linearity are not significant in either the serial or the discrete case, however, the F

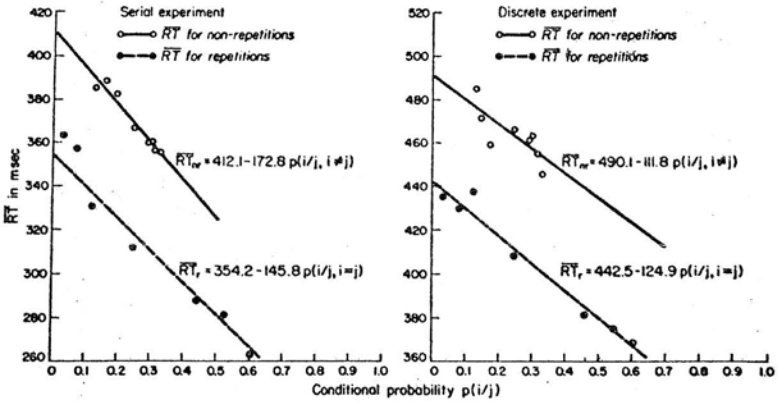


FIG. 3. Mean RT of correct responses for repetitions and nonrepetitions as a function of the conditional probability of repetitions $p(i|j, i = j)$ and nonrepetitions $p(i|j, i \neq j)$, respectively, in both experiments. (The equations represent the best fitting least-squares lines to the points.)

ratio is highest for the discrete nonrepetition data (deviations from linear trend-serial: repetitions $F = 1.28$, $df = 6/42$, $p > .25$, nonrepetitions $F = .63$, $df = 7/49$; discrete: repetitions $F = .3$, $df = 6/42$; nonrepetitions $F = 1.78$, $df = 7/49$, $.20 < p < .10$); (*c*) the difference in the slope for repetitions and nonrepetitions is not statistically significant in either experiment (serial: slope for repetitions $m = -145.8$, slope for nonrepetitions $m = -172.8$, $t = 1.37$, $df = 76$, $.2 < p < .1$; discrete: slope for repetitions $m = -124.9$, slope for nonrepetitions $m = -111.8$, $t = .41$, $df = 74$); (*d*) the difference in the intercepts for repetitions and nonrepetitions is statistically significant in both experiments (serial: intercept for repetitions $b_r = 354.2$, intercept for nonrepetitions $b_{nr} = 412.1$, $t = 20.6$, $df = 9$, $p < .01$; discrete: intercept for repetitions $b_r = 442.5$, intercept for nonrepetitions $b_{nr} = 490.1$, $t = 4.3$, $df = 9$, $p < .01$.

On the basis of these data the RT for repetitions and nonrepetitions may be expressed as:

$$\overline{RT}_r = -mp(i|j, i = j) + b_r \quad [6a]$$

and

$$\overline{RT}_{nr} = -mp(i|j, i \neq j) + b_{nr} \quad [6b]$$

where:

m = the slope of RT for repetitions and nonrepetitions as a function of their respective conditional probabilities.

b_r, b_{nr} = the intercept of the RT for repetitions and nonrepetitions, respectively.

Substituting Expressions [illegible], 6a, and 6b in Equation 2 and simplifying the notation by denoting $p(i|j, i \neq j)$ by p we obtain the original tautologous expression in the following form:

$$\begin{aligned}\overline{RT} &= (K-1)p(-mp + b_{nr}) \\ &+ [1 - (K-1)p] \\ &\times [-m(1 - (K-1)p) + b_r] \quad [7]\end{aligned}$$

which reduces to:

$$\begin{aligned}\overline{RT} &= p^2[-m(K^2 - K)] \\ &+ p[(K-1)(b_{nr} - b_r + 2m)] \\ &- m + b_r \quad [8]\end{aligned}$$

Equation 8 expresses the overall mean RT as a parabolic function of $p(i|j, i \neq j)$. The only assumptions that have been made are that the $\overline{RT}$s for repetitions and nonrepetitions are linear in $p(i|j)$, and that their slopes are equal. The linearity assumption implies a decision to interpret the relatively higher F ratio for the nonlinear component of the nonrepetitions in the discrete case as due to random experimental variability. The assumption of equality for the slopes is not necessary but appears to be justified by the data and simplifies the expression. Since the slope

(m) and the intercepts (b_r and b_{nr}) of the RT for repetitions and nonrepetitions are the parameters in the equation, the effect of K on these parameters would have to be clarified before generalizing the expression to other values of K. Strictly, therefore, the parameters in Equation 8 should have a K subscript pending such a clarification. A detailed treatment of these questions is beyond the scope of the present paper; however, the effect of K will be discussed briefly in a later section.

Figure 4 illustrates the overall mean RT for correct and error responses in both experiments as a function of $p(i\,|\,j, i \neq j)$. The parabola was generated by substituting the appropriate parameter values for each experiment in Equation 8 and fits the data quite well—as indeed, it should; for it will be recalled that Equation 8 is simply a computational formula derived from Equation 2 by substitutions. One could well ask why the fit is not better than it is, and a different reason would have to be provided in each experiment. In the serial case it was assumed that the slopes for repetitions and nonrepetitions were equal, and a mean slope was used in the computation; however, even though the slopes were not *statistically* different, they were nevertheless *numerically* different—hence, the slight departures of the data from the parabola. In the discrete experiment the variability of the data is probably a reflection of the variability of the nonrepetitions themselves, as indicated by their relatively higher F ratio. There are two additional features of interest to these data. First, the overall mean RT for errors is faster than for correct responses, by what appears to be a constant amount, in both experiments. This would indicate that few, if any, errors occur at random. That is, the RT as well as the occurrence of errors seem to be determined by the same sequence properties and processing microstructures as determine the correct responses;[5] this point will be discussed in more detail in a later section. Second, the condition in which $p(i\,|\,j, i \neq j) = 0$ represents a sequence in which the same stimulus is repeatedly presented on every trial which, of course, corresponds to a simple reaction time task. As can be seen from Equation 8, the mean RT in that case is completely determined by m and b_r. It is encouraging to note that the extrapolated $\overline{RT}$ value for this condition in the serial experiment (195 msec.) corresponds closely to the value of simple RT for visual stimuli (Teichner, 1954; Woodworth, 1938); it is equally disconcerting, however, that the mean RT for the same condition in the discrete task is far too high (324

[5] This contention receives additional support from the posterror responses whose RT is longer than for correct responses by a constant amount; this finding confirms previous more general reports by Burns (1965) and Rabbitt (1966).

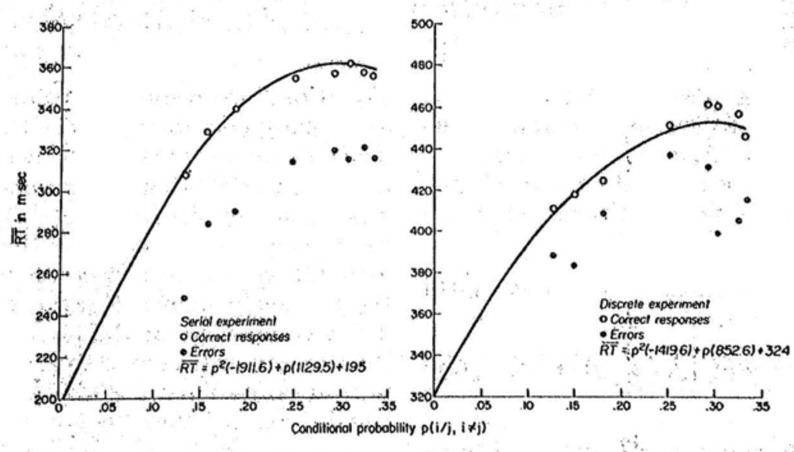


FIG. 4. Overall mean RT for correct (○) and error (●) responses for both experiments, as a function of $p(i\,|\,j, i \neq j)$. (The parabolas were generated by substituting the appropriate parameter values in Equation 8.)

msec.) to allow for a simple interpretation of the intercept.

Finally, setting $d\overline{RT}/dp = 0$ in Equation 8 yields the following value of p for the maximum $\overline{RT}$:

$$p = \frac{1}{K}\left[1 + \frac{b_{nr} - b_r}{2m}\right] \qquad [9]$$

Thus, if the $\overline{RT}$s for repetitions and nonrepetitions are equal (i.e., if $b_{nr} = b_r$), the fraction within the brackets vanishes and the maximum $\overline{RT}$ will be obtained at $p = 1/K$; this, of course, is in precise accord with the information hypothesis. However, if $b_{nr} \neq b_r$, then the maximum $\overline{RT}$ will be at variance with predictions from the information hypothesis. At the maximum, therefore, the term within the parentheses in Equation 9 may be interpreted as a correction term for the information-hypotheses formulation of the problem. The correction term for other values of the function is more difficult to calculate, even for values of $p < 1/K$, since it involves the comparison of a logarithmic with a second degree equation. The conclusion regarding the information hypothesis, however, is fairly clear cut without these comparisons: it appears reasonable to affirm that since transmitted information has been shown to be neither a necessary nor a sufficient condition for choice RT the information hypothesis must be rejected. In the remaining sections of this paper further consequences and conjectures based on these findings will be examined.

AREAS OF CONTINUITY BETWEEN THE SEQUENTIAL AND THE INFORMATION APPROACH

If this discussion is to be extended beyond merely calling attention to the confounding between H and $p(i|j, i \neq j)$, cognizance should be taken of the remarkable success that the information-based approach has had in dealing with choice RT problems. It would, in addition, be desirable to explore the relationships and points of contact between the approach that is being suggested here and that body of data which owes its discovery to the information-based approach, so as not to lose the wealth of empirical findings that it represents. An attempt will now be made toward these ends by briefly examining three problems: (*a*) the effect of increasing the number of alternatives, (*b*) Hyman's "nonadditive combination of components within conditions," and (*c*) the effect of S-R compatibility.

1. *The Effect of Increasing the Number of Alternatives* (K)[6]

The general question that is being raised at this point concerns those experiments where H, K, and $p(i|j, i \neq j)$ were simultaneously varied, and whose data are fitted by a linear regression of $\overline{RT}$ on H. Can the goodness of those fits now be accounted for in terms of the characteristics of the RT for repetitions and nonrepetitions and their associated probabilities?

As a first step, it is important to note that in Hyman's study, which is the most representative, the average range of uncertainty per number of alternatives was .75 bits (.53 bits with $K = 2$ up to 1.19 bits with $K = 6$). Within such a narrow range it is most unlikely that an experimental distinction could be made between the linear relationship which the information hypothesis requires, and the parabolic relationship described by Equation 8. If, therefore, it is assumed that $\overline{RT}$ is linear over a 1-bit range, then the slope of $\overline{RT}$ with respect to H can be calculated from Equation 8 for the eight-choice and the four-choice case where $p(i|j, i \neq j) < 1/K$. Table 1 lists the different values of $p(i|j)$ that would generate eight- and four-choice sequences with 2 and 3 bits, and 1 and 2 bits, respectively (cf. Equation 3). Substituting these values in Equation 8, the

[6] In addition to the confounding between H, p_{nr}, and K which has already been discussed, Brebner and Gordon (1962) have pointed out that the effects of increasing the number of alternatives are also confounded with the effects of decreasing the probability of occurrence of these alternatives. This, however, will not be discussed in this paper; suffice it to say that meticulous care must be exercised to obtain unconfounded results in this area.

TABLE 1

VALUES OF $p(i|j)$ FOR GENERATING 1 AND 2 BIT, AND 2 AND 3 BIT SEQUENCES WITH FOUR AND EIGHT ALTERNATIVES, RESPECTIVELY

K	H	$p(i\|j)$	
		$i = j$	$i \neq j$
4	1	.80	.0667
	2	.25	.25
8	2	.625	.0536
	3	.125	.125

Note.—See Equation 3.

following values of $\overline{RT}$ are obtained:

$$\text{for } (K = 8, H = 3\overline{RT}) = -.875m + .875(\Delta b_k + 2m) - m + b_{rk} \quad [10a]$$

$$\text{for } (K = 8, H = 2)\overline{RT} = -.161m + .375(\Delta b_k + 2m) - m + b_{rk} \quad [10b]$$

$$\text{for } (K = 4, H = 2)\overline{RT} = -.75m + .75(\Delta b_k + 2m) - m + b_{rk} \quad [10c]$$

$$\text{for } (K = 4, H = 1)\overline{RT} = -.053m + .20(\Delta b_k + 2m) - m + b_{rk} \quad [10d]$$

Subtracting the lower from the higher $\overline{RT}$ values within each value of K we obtain the following slopes of $\overline{RT}$ with respect of H:

$$\text{for } K = 8, \Delta RT = .286m + .50\Delta b_k \quad [11a]$$

$$\text{for } K = 4, \Delta RT = .403m + .55\Delta b_k \quad [11b]$$

where:

$\Delta b_k = b_{nrk} - b_{rk}$ (i.e., the difference in the intercept between repetitions and nonrepetitions for $K = 4$, and 8).

m = the slope of RT for repetitions and nonrepetitions (cf. Equations 6a and 6b).

ΔRT = slope of $\overline{RT}$ with respect of H.

Equations 11a and 11b make it clear that the slope of $\overline{RT}$ with respect to H is determined by the values that m and Δb assume with different numbers of alternatives. Let us first consider m. There is no obvious theoretical, or readily available experimental evidence that would indicate that merely changing the number of alternatives also changes the slope of repetitions and nonrepetitions. If anything, these slopes appear to be fairly constant over different values of K. On the other hand, there is strongly suggestive evidence to indicate that Δb does increase with increasing values of K. First, after having noted that the RT for repetitions was faster than for nonrepetitions, Hyman states that,

> An examination of the data showed that this phenomenon was quite marked for the situation with four or more alternatives and steadily declined until it disappeared or became slightly negative for the case with just two alternatives [Hyman, 1953, p. 195].

This statement is corroborated by his own data (Hyman, 1953, Fig. 2). Second, in replotting some of our previous repetition and nonrepetition data (Kornblum, 1967) so as to display the values of Δb, the evidence points quite clearly to an increase in Δb as a function of K (see Figure 5). This increase, furthermore, appears to be primarily attributable to a far greater increase in the RT for nonrepetitions than for repetitions, although both increase with increasing values of K. Thus, varying K

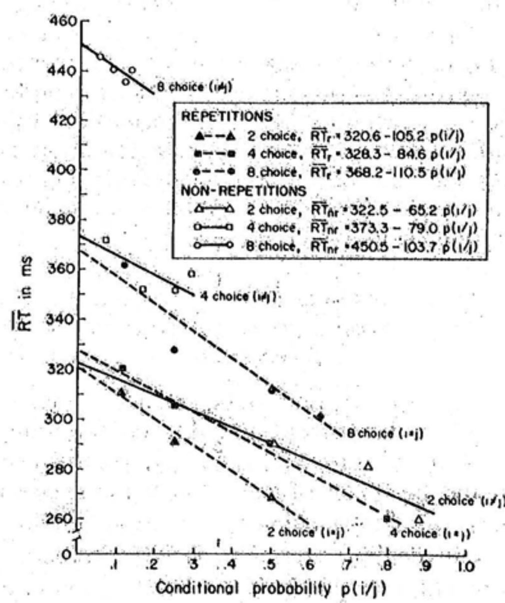


FIG. 5. RT for repetitions and nonrepetitions as a function of $p(i|j)$ for two-, four-, and eight-choice sequences with equiprobable stimuli; these data were previously reported in a different form (Kornblum, 1967, Fig. 1).

would appear to have a differential effect on the RT for repetitions and nonrepetitions. In particular, increasing K leads to a greater increase in b_{nr} than in b_r, which in turn leads to Δb being an increasing function of K. From Equation 11, it therefore follows that as long as the increase in Δb exceeds a minimal value, the slope of $\overline{RT}$ with respect to H will be an increasing function of K.

When Hyman's original data are reconsidered in this light, and the overall mean RT for two, four, and eight alternatives are examined separately (see Figure 6), it can be seen that the slope of $\overline{RT}$ with respect to H does indeed tend to increase with K among his four Ss. Similar findings are also apparent from the results of other studies (e.g., Kornblum, 1967).

We would, therefore, conclude that the high correlations that Hyman and others have reported for $\overline{RT}$ data that spanned different values of K is a fortuitous consequence of:

(*a*) Δb being an increasing function of K, jointly with

(*b*) the highly restricted range in H which each value of K represented in those experiments.

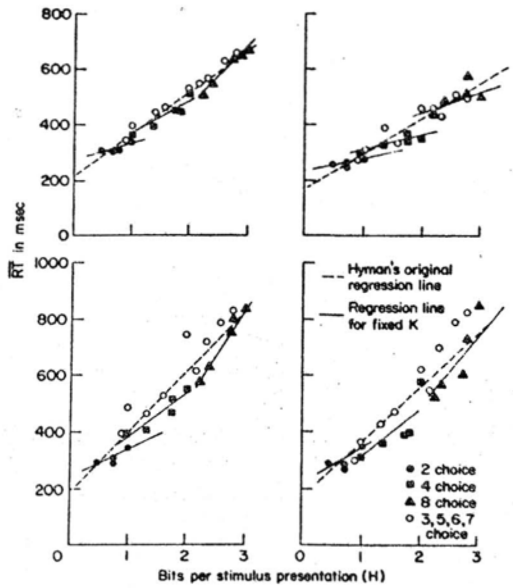


FIG. 6. This is a slightly modified version of a figure which was originally published by Hyman (1953, Fig. 1). (The dotted line is Hyman's original regression line. The solid lines are the best-fitting straight lines, estimated by eye, for his data with two, four, and eight alternatives.)

2. *Hyman's "Nonadditive Combination of Components within Conditions"*

A consideration of this second problem follows quite naturally from the previous discussion. It will be recalled that Hyman encountered a dilemma when he attempted to predict the mean RTs for the components within a condition (e.g., the RT to the rare or the frequent stimulus within a condition) from the regression line which had been fitted to the overall means of all the conditions. When the predicted values were compared with the actual observations, he found that the observed mean RTs for the low information components were slower than the predicted means, and the RTs for the high information components were faster than the predicted means.

If our conclusions in Section 1 above are at all reasonable, they may well lead to a resolution of this dilemma. Since the conditions means which were fitted by the overall regression line include sets of means for which the numbers of alternatives differ, and we concluded that the slope of the regression line for *fixed* numbers of alternatives increases with K (the lines themselves also being upwardly displaced by small increases in b_r), it follows that the slope of the *overall* regression line will be steeper than the slope of some, if not all, of the regression lines which are separately fitted to the means of conditions having the same value of K. This is quite apparent in Figure 6. Now, as long as the range in H which is represented by these different values of K is kept within a sufficiently narrow band, the discrepancy between the slopes of the overall and the individual regression lines will remain hidden. It is only when components within a condition are selected, which may be 3 to 4 bits apart, that this difference is sufficiently amplified to be noted.

Further corroborating evidence for this argument can be found in the results reported by Fitts, Peterson, and Wolpe

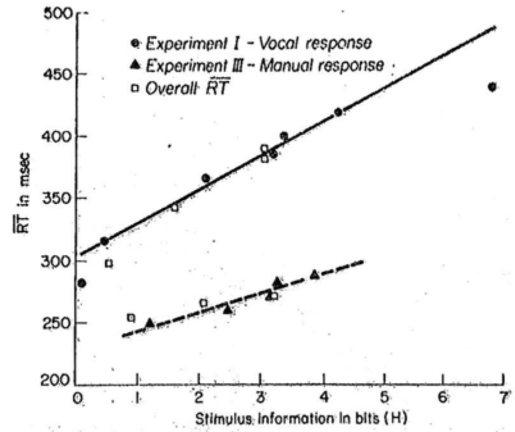


FIG. 7. This is a slightly modified version of a figure which was originally published by Fitts and Posner (1967, Fig. 31) from data reported by Fitts, Peterson, and Wolpe (1963). (The overall mean RT points—□—were added for purpose of the present paper—see text for explanation.)

(1963) who used two nine-choice tasks to study the effects of redundancy on RT. In one task (Experiment I) the stimuli consisted of visually presented numerals with vocal responses being made to them; in the other task (Experiment III) the stimuli consisted of semicircularly arranged neon lights with a touch-sensitive circuit beneath each light. Different levels of redundancy were obtained in each task by having one of the stimuli occurring very frequently, while the others occurred proportionately less frequently and equiprobably. Figure 7 represents these data as plotted in Fitts and Posner (1967) in addition to the *overall* means for the four conditions of Experiment I, and for the three conditions of Experiment III. Except for the two extreme components of Experiment I, the same regression line fits both the component and the overall means quite well. This, of course, is perfectly reasonable for, in terms of the argument that has been presented, as long as the number of alternatives is kept constant, the line which is fitted to the conditions means and the line which is fitted to the component means within conditions are one and the same line.

The remaining puzzle in this dilemma is the fact that Hyman seems to report that the direction of the difference between the predicted and the observed component means is "without exception" one in which the low information component is underestimated, and the high information component is overestimated by the overall regression line. This would follow as long as the slope of the overall regression line is steeper than the slope of the individual regression lines. However, as can be seen from Figure 6, for three out of the four *S*s the slope for the eight-choice conditions appears to be steeper than the slope for the overall regression line. The fact that the high information components in the eight-choice conditions are also overestimated remains to be explained.

3. *The Effect of S-R Compatibility*

Slightly different considerations must be brought to bear on this problem depending on whether one is dealing with conditions for which K is fixed, or whether K is allowed to vary.

Varying S-R compatibility with constant K. Bertelson's findings regarding the effect of S-R compatibility on the RT for repetitions and nonrepetitions have already been referred to briefly in a previous section (Bertelson, 1963). In comparing the results of a more compatible and a less compatible mapping for a two-choice and a four-choice task, with equiprobable and independent stimuli, Bertelson reports that in both cases the less compatible mapping had the effect of increasing the difference between the RT for repetitions and nonrepetitions. This increase, furthermore, was primarily attributable to a much larger rise in the RT for nonrepetitions than for repetitions in both cases. In terms of the present formulation, these findings indicate that an increase in the S-R incompatibility of a task leads to an increase in Δb and a slight rise in b_r. (Whether or not the slope of the repetitions and nonrepetitions is affected by changes in compatibility cannot be ascertained from these data alone; however, this aspect of the problem will be dealt with below.)

At first glance, these findings would appear to contradict our own results from the serial and discrete experiments in which, it will be recalled, a more and a less compatible task were used, respectively. Our results indicated that Δb was smaller and the slope shallower for the less compatible task. However, it must also be recalled that the less compatible task was deliberately chosen for the discrete experiment in order to have reasonable assurances that a difference between the RT for repetitions and nonrepetitions would be obtained at an R-S interval which was over 20 times what was used in the serial experiment. Had both tasks been run with the same R-S interval, then according to the assumption on which our choice of tasks was predicated, Δb for the less compatible task would indeed have been larger than for the more compatible one. Our results, therefore, in no way contradict the findings cited above.

Given that an increase in the S-R compatibility of a task leads to an increase in Δb, and a slight rise in b_r, two straightforward consequences follow:

According to Equation 8 we would expect the mean overall RT for fixed K to be an increasing function of the incompatibility of the task. The clearest evidence in this regard can be found in a study done by Peterson (1965) in which 24 different S-R mappings of four stimuli and four responses were used as RT tasks. The stimuli were four neon lights placed at the corners of a 12-inch tilted square; the response was made by lifting the right index finger from a home position at the center of the square, and touching the appropriate target, which consisted of a circular area 1/2 inch in diameter around each light. Peterson reports a .905 Spearman rank-order correlation between the ranked difficulty of the mappings, which we take to be an index of S-R incompatibility, and the overall mean RT. Additional confirmation is also evident in Figure 7 when the overall mean RT of the two experiments is compared at 3.17 bits, which represents the condition in which the stimuli are equiprobable and independent. The $\overline{RT}$ in Experiment III is over 100 msec. faster than in Experiment I. Fitts et al. (1963) note that ". . . the spatial ensemble used in Experiment III is one of the most highly compatible that has yet been studied. . . [p. 431]."

The second consequence follows from Equation 11, according to which one would expect the slope of $\overline{RT}$ with respect to H to increase as the incompatibility of the task, that is, Δb, increased. Here, again, from Figure 7 it is quite evident that the slope of the less compatible task (Experiment I) is distinctly steeper than the slope of the more compatible task (Experiment III).

Varying both S-R compatibility and K. The most critical evidence with regard to this question is Leonard's (1959) finding of no measurable increase in the overall mean RT for two-, four-, and eight-choice sequences in which the stimuli consisted of tactile vibrations to the finger tips and the responses consisted of depressing the key under the finger that had been so stimulated. Since the overall mean RT has been shown to be a weighted sum of the RT for repetitions and nonrepetitions (Equations 1 and 8) and, according to Leonard's results, varying the proportions of these two components does not bring about any measurable changes in the overall mean RT, it must be concluded that for the highly compatible task that he used the slope of the RT for repetitions and nonrepetitions is zero, and either (*a*) Δb decreases with K, which is most unlikely, or (*b*) Δb is zero for the range of K that he used. The studies in which S-R compatibility was varied over a range of K were summarized by Fitts and Posner (1967, Figure 32), who show that increasing the incompatibility of the task leads to an increase of the slope of $\overline{RT}$ with respect to H. Since this slope in the extreme case is zero (Leonard, 1959), it must be concluded that incompatibility and K combine multiplicatively so as to bring about a change in Δb as well as in the slope of the repetitions and nonrepetitions.[7]

[7] Our conclusions regarding the effects of K and S-R compatibility are essentially the same as those reached by Sternberg (1968) via a completely different argument.

ERROR RESPONSES

The Probability of Errors and "States of Readiness"

One of the most persistent themes in the *RT* literature can be found in the recurrent attempts to account for the changes in *RT*, both simple and choice, in terms of concomitant changes in *S*'s "readiness" or "preparedness." These efforts are easily and variously identifiable by noting the pivotal role of such terms as "set," "context," "plans," "expectancy," "anticipation," "attention," or "preparedness" in the main arguments. Falmagne's (1965) model represents the most recent as well as the most precise statement of this approach. However, the attempts to bring conceptual clarity and specificity to these concepts have, on the whole, been marked by the perpetuation of the original ambiguities and obscurities. The present paper is, of course, not the appropriate place to resolve such a major issue. However, in spite of these inherent ambiguities, concepts like "readiness" or "preparedness" seem sufficiently useful in their common-sense meaning that one is tempted to try to use them in reaching for an understanding of at least the molar aspects of *RT* data.

TABLE 2

PERCENTAGE OF REPETITION AND NONREPETITION TRIALS ON WHICH ERRORS WERE MADE

Condition: $p(i\mid j, i=j)$	% erroneous repetition trials	
	Serial	Discrete
.03	7.0	4.0
.08	2.3	2.1
.12	1.7	2.8
.25	1.0	2.8
.44	.2	1.8
.53	.04	1.3
.61	.2	.6
Condition: $p(i\mid j, i\neq j)$	**% erroneous nonrepetition trials**	
	Serial	Discrete
.13	8.5	6.6
.16[a]	5.8	6.5
.19[a]	5.6	4.2
.25	3.9	4.2
.29	3.2	4.3
.31	3.0	4.2
.32	3.0	4.4
.33	2.9	4.0

Note.—"Condition" identifies the particular sequences under consideration and refers to the conditional probability of repetitions and nonrepetitions for those sequences. In the former this corresponds to $1-p_{nr}$, in the latter it corresponds to $p_{nr}/3$.

[a] The $p(i\mid j)$ values should be modified slightly for the discrete experiment (cf. caption for Figure 1).

The common-sense meaning of "ready" or "prepared" would seem to be preserved in the notion that whatever events one is more ready or prepared for, one's performance with respect to those events would be faster and more accurate than in the case of events that one is less prepared for. Furthermore, the meaning of "prepared" or "ready" would not seem to be violated if it were supposed that one's "preparedness" is greater for more probable events than it is for less probable events.

If the data for correct responses are examined in this light, the decrease in the *RT* for repetitions and nonrepetitions as a function of the appropriate values of $p(i\mid j)$ provides initial support for an interpretation of the data in terms of readiness where the events are simply repetitions or nonrepetitions. However, the error responses provide a far richer source of information in this regard.

The overall error rate in both experiments is approximately constant between the 3% and 4% level across all the conditions that were used. This gross way of looking at errors is, therefore, rather uninteresting. Table 2 presents the percentage of repetition and nonrepetition trials on which errors occurred for both experiments. Two aspects of these data are worth noting: (*a*) nonrepetition trials lead to a proportionately higher error rate than repetition trials, and (*b*) the higher the conditional probability $[p(i\mid j)]$ for repetitions or nonrepetitions, the lower the probability of an error on that type of trial. This last effect is much more marked and systematic in the serial than in the discrete case.

Since the meaning of "readiness," as presently used, includes the concept of increased readiness for more probable events, this last effect could be interpreted as a decrease in the probability of an error with

TABLE 3

PERCENTAGE OF ERRORS IN A PARTICULAR SEQUENCE THAT WERE INCORRECT NONREPETITION RESPONSES

Condition: $p(i\mid j, i\neq j)$	Serial	Discrete
.13	25.6	54.1
.16	38.8	59.1
.19	44.1	72.7
.25	61.5	79.8
.29	71.4	72.6
.31	77.3	66.0
.32	81.0	71.0
.33	84.2	80.8

Note.—Based on all errors. The percentage of incorrect repetitions is given by the complement of the values shown.

an increase in one's readiness for events of the type on which the errors occur. If this is the case, then this increased state of readiness should not only be reflected in the level of errors for different types of trials but an indication should also be found in the very type of response that is made when an error occurs. That is, whatever the type of event for which one is the most prepared, be it repetitions or nonrepetitions, the error response itself should reflect that differential state of readiness. Table 3 shows the percentage of erroneous nonrepetitions based on all the errors that were made in particular sequences with different values of conditional probability for nonrepetitions. Since very few errors occurred on repetition trials, and when they did occur they were necessarily nonrepetition types of responses, Table 4 is presented in which only the errors that occurred on nonrepetition trials are considered. It is evident from both tables that as the conditional probability of nonrepetition increases so does the proportion of erroneous nonrepetition responses. As was true in Table 2, this effect is more systematic in the serial than in the discrete case.

The data in Tables 2, 3, and 4 indicate that as readiness for the more probable event increases, be it a repetition or a nonrepetition, not only does the likelihood of an error on that type of trial decrease, but when an error does occur, then the error response itself is of the type for which readiness is greatest. Similar findings have been reported previously for an eight- and a four-choice task (Kornblum, 1967). When error responses are examined in terms of RT, it has already been noted in Figure 4 that the pattern of RT for errors is similar to that for correct responses. This observation receives further support from a breakdown of the RT for errors into repetitions and nonrepetitions; as was the case for the correct responses, the RT for repetitions and nonrepetitions error responses appears to be a decreasing linear function of $p(i\mid j)$ with a steeper slope and a lower intercept than was found for correct responses; the scant and variable nature of the data do not warrant a more accurate quantitative assessment at this time. However, it seems evident that a process such as readiness or preparedness does play a central role in determining performance in such tasks.

Errors as Confusions

In the previous section stimuli and error responses were both treated as belonging to the class of repetitions or nonrepetitions. These two classes were treated as if they were homogeneous sets, and no attempt was made to find any further distinguishing features within those sets. However, it will be recalled that the keyboard task in the serial and the discrete experiments used the middle and index fingers of the left and the right hand for the execution of the responses. If the two index fingers are con-

TABLE 4

THE PERCENTAGE OF THOSE ERRORS MADE ON NONREPETITION TRIALS WHICH CONSISTED OF ERRONEOUS NONREPETITION RESPONSES

Condition: $p(i\mid j, i\neq j)$	Serial	Discrete
.13	22.8	47.6
.16	37.9	49.3
.19	42.7	62.5
.25	58.1	75.3
.29	69.3	67.8
.31	75.7	64.6
.32	79.9	70.0
.33	84.2	80.8

Note.—The percentage of incorrect repetitions is given by the complement of the values shown.

TABLE 5

CLASSIFICATION OF STIMULI AND ERROR RESPONSES

Error response classification	Serial experiment					Discrete experiment				
	Stimulus classification					Stimulus classification				
	Repetitions	Nonrepetitions			Totals	Repetitions	Nonrepetitions			Totals
	HF	$H\bar{F}$	$\bar{H}F$	$\bar{H}\bar{F}$		HF	$H\bar{F}$	$\bar{H}F$	$\bar{H}\bar{F}$	
$H\bar{F}$	3.8	24.6[a]	20.6	23.4	73.0	6.4	12.4[a]	14.2	15.4	48.4
$\bar{H}F$	.6	5.9	10.4[a]	4.4	21.3	4.5	12.9	13.9[a]	14.5	45.8
$\bar{H}\bar{F}$	.1	.03	.7	5.0[a]	5.8	.1	1.0	.8	3.9[a]	5.8
Totals	4.5	30.5	31.7	32.8	100.0	11.0	26.3	28.9	33.8	100.0

Note.—The stimuli on the trials on which errors occurred are classified in terms of a comparison between the response assignment of the stimulus on that trial with the response assignment of the stimulus on the preceding trial. The error responses are classified in terms of a comparison between the response that was executed and the response that was in fact called for by the stimulus on that particular trial. An HF row does not appear in this Table because these are, by definition, correct responses. The cell entries represent the proportion of all the errors that fall in these cross-classifications (see text).

[a] Repetitions type error responses.

sidered as equivalent, and the two middle fingers are considered as equivalent, then the sequential aspects of a stimulus on any one trial may be specified by noting exactly how the response assignment for the stimulus on any one trial changes, if at all, from the response assignment for the stimulus on the preceding trial. Similarly, an error response may be specified by noting the exact discrepancy between the response assignment of a stimulus on any one trial and the response that was executed on that very same trial. An example of the way in which stimuli and error responses would be classified will clarify the method. If the response assignment for the stimulus on Trial n is the left index finger, and the assignment for the stimulus on Trial $n+1$ is the left index finger again, then clearly the stimulus on Trial $n+1$ is a repetition and would be characterized as HF, indicating that neither the hand nor the finger assignments had been changed between trials; consider the same response assignment on Trial n, if now the stimulus on Trial $n+1$ calls for a response with the left middle finger, then this trial would be characterized as $H\bar{F}$ (where the bar is used to indicate change, or negation), indicating that on Trial $n+1$ the hand assignment had remained the same, but the finger assignment had been changed to a nonequivalent finger. The other two types of nonrepetitions would be characterized as $\bar{H}F$ and $\bar{H}\bar{F}$. Whereas stimuli on any one trial are classified in terms of the stimulus on the preceding trial, the error responses are classified with reference to events within the same trial. Thus, given that a response was made with the left index finger and that the stimulus on that trial called for the right index finger, the error would be characterized as $\bar{H}F$, indicating that the response had been executed with a finger equivalent to the one that had been called for, but with the other hand. The other two types of error responses would be characterized as $H\bar{F}$ and $\bar{H}\bar{F}$; HF responses, of course, are correct.

This method of characterizing stimuli and responses leads to the cross-classification that is illustrated in Table 5. The data in the table represent the percentage of all the errors that fall within these particular cross-classifications. Since they are based on all the errors that were made in each experiment, they are crude at best, and overlook some important differences that have already been noted between conditions. The results are nevertheless instructive even though some of them recapitulate earlier findings; for example, in both experiments repetition trials (HF) lead to considerably fewer errors than nonrepetition trials. It can also be seen that the three types of nonrepetition trials all lead to approximately the same error level within each experiment. One of the most striking features of these data, however, is the relatively rare occurrence of an error response in which both the finger and the hand are wrong. Even though such double confusions are rare they occur most frequently on those trials in which the stimulus itself calls for a change in both finger and hand; that is, these errors are erroneous repetitions of the previous response. The other two types of error responses display a different pattern in each experiment. In the serial experiment, the majority of the errors (73%) consist of finger confusions ($H\bar{F}$); hand confusions ($\bar{H}F$) accounting for only 21% of all the errors. However, here as before, hand confusions occur most frequently on those trials where the stimulus itself calls for a change of hands; these errors are again erroneous repetitions of the previous response. The data of the discrete experiment stand in sharp contrast to the serial data; in the discrete experiment, hand and finger confusions occur with equal likelihood and appear not to be related to the classification of the stimulus, except in the $\overline{HF}$ case.

The nature of these confusions and their disproportionalities within and between the two experiments are indicative of the stimulus having been processed prior to the execution of an error response. Had the responses been made without reference to the stimulus, the proportion of errors in each row of Table 4 would have been one-third of the total—this is clearly not the case. If hands and fingers are considered as attributes of the stimulus and/or response space, then these data also indicate that errors are the result of a failure to distinguish the separate points on either one or on the other attribute (but rarely both). These data unfortunately do not allow any hard conclusions to be drawn regarding the relative magnitude of "stimulus" or "response" effects. Even though it is doubtful that either could be characterized independently of the other, the most fertile source of clues to this problem may well be found in a further analysis of the posterror responses, classified according to the method which has been described in this section. The richness of the posterror responses lies in the fact that they are correct responses preceded by a trial on which the stimulus and the response differ from each other. The sequential aspects of the posterror trial may, therefore, be characterized either in terms of the preceding stimulus or in terms of the preceding response; in this way the two effects may be unconfounded. A further disappointment in these data is that they do not allow any hard conclusions to be drawn regarding the manner in which errors are attributable to properties of the stimulus space, the response space, or the correspondence and mapping between them; this is unfortunate, because this question is probably at the heart of the S-R compatibility problem.

THE HETEROGENEITY AND PARTITIONING OF REPETITIONS AND NONREPETITIONS

An attempt has been made in this paper to identify some of the variables of choice *RT* tasks in terms of which performance could be described with some degree of precision and parsimony. The dichotomous classification of repetitions and nonrepetitions seems to be promising in this regard and also sets the stage for the next effort in which the question must be posed: What models, mechanisms, or processes could generate the orderliness that has been found? The success of this later effort will be partially determined by the success with which the essential properties of the data have been identified in this initial effort. If the effects which we have described are themselves the consequence of further distinctions and regularities within the class of repetitions and nonrepetitions then the models or mechanisms would probably gain in scope and generality by addressing themselves to these more fundamental structures.

In the preceding section on "errors as confusions" it has been shown that orderly distinctions can be drawn between different types of error responses. Systematic differences in the *RT* for correct responses have also been reported within repetitions and nonrepetitions as a function of their rank, or ordinal position (Bertelson, 1961; Falmagne, 1965; Hyman, 1953). The rank of a repetition refers to the first, second or *n*th consecutive occurrence of a repetition. Nonrepetitions may be ordered in two different ways: (*a*) they may be ranked without regard to the stimulus itself, in which case the critical event is simply the occurrence of a nonrepetition, be it the first, second, or *n*th consecutive occurrence of a nonrepetition, or (*b*) the ranking may be done on the basis of the number of different stimuli that intervene between two occurrences of the same stimulus. The experimental results indicate that the *RT* for repetitions decreases with rank, while the *RT* for nonrepetitions increases with rank (in Sense *b* above). These findings are confirmed by a cursory examination of the serial data in the present study.

It seems perfectly obvious that as the first-order conditional probabilities are changed [i.e., $p(i|j, i \neq j)$] so are the probabilities of repetitions and nonrepetitions of different rank. Hence, as was true of the overall mean *RT*, the mean *RT* for repetitions and nonrepetitions may themselves be partitioned into subsets according to rank, and expressed as the sum of the *RT* with particular ranks, weighted by the probability of those ranks; that is,

$$\overline{RT}_r = \sum_s p_{s,r}\overline{RT}_{s,r}$$

and

$$\overline{RT}_{nr} = \sum_s p_{s,nr}\overline{RT}_{s,nr}$$

where

p_s = probability of repetitions (r) or nonrepetitions (nr) with Rank s

$\overline{RT}_s$ = mean RT for repetitions or nonrepetitions with Rank s

The analysis and examination of the *RT*s for repetitions and nonrepetitions with different ranks constitutes the core of Falmagne's (1965) model. As such, it represents one of the more promising and fruitful approaches for discovering the basic microstructure of the psychological processes that determine performance in such tasks.

CONCLUSIONS

A descriptive-reductionist analysis of choice *RT* has been undertaken in which the molar aspects of performance are described in terms of a few simple and systematic properties of the constituents of such performance. The information hypothesis has been shown to be untenable in the form in which it has commonly been held. This, of course, in no way precludes the possibility of finding a more appropriate space for which *H* could be a suitable measure of the determinants of performance. The main thrust of this paper, however, is descriptive and is intended to provide a data base for future theoretical endeavors. The argument has essentially proceeded by decomposing a mean into a weighted sum of differentially sensitive partitions; the gross phenomenon, thus dissected, seems far richer and more tractable than it originally appeared.

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