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ABSTRACT

VARIABILITY OF VISUAL LATENCY IN BINOCULAR AND MONOCULAR VIEWING CONDITIONS: JUDGMENTS OF TEMPORAL ORDER AND SIMPLE REACTION TIME

By

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Temporal order judgment (TOJ) and simple reaction time (RT) were used to estimate visual latency variability in monocular and binocular viewing conditions. Stimulus luminance was used to manipulate latency variability parametrically. Several models of probability summation were examined and it was concluded that there should be no difference in TOJ in the two viewing conditions due to probability summation. A "baseline for independence" was established which, if exceeded, would imply that the monocular latencies were dependent. Pilot data suggested that binocular viewing increases sensitivity to temporal order. Experiment 1 replicated this result. Latency variability inferred from TOJ and RT variability measured in Experiment 2 were both lower for binocular viewing and higher luminance. A large non-sensory source of variability was attributed to RT. The results of both experiments were best accounted for by assuming that binocular latency was most nearly equal to the shorter of the two monocular latencies.

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LIST OF SYMBOLS

SYMBOL	MEANING
a	In the two-component model of arrival latency used in the low- and high-threshold versions of the binocular summation models, "a" designates the first component, that is, the time from the onset of the stimulus until its arrival at the combination point. This symbol is also used as a coefficient in an arbitrary power function.
b	This may designate the second component latency in the two-component model of arrival latency used in the low- and high-threshold versions of the binocular summation models. Here it is the time between the arrival at the combination point and the arrival at the decision center. This symbol is also used as the arbitrary exponent in a power function.
B	A subscript which specifies that some other symbol refers to the binocular viewing condition.
c	This may used as a subscript on p or P to indicate the probability of a correct response.
d	A particular value of the stimulus onset asynchrony
d'	The sensitivity measure in signal detection theory
E	The expected value operator
f	A function, usually but not always a probability density function
F	A cumulative distribution function
f_o	The standard normal probability density function
F_o	The standard normal cumulative distribution function

SYMBOL	MEANING
F_o^{-1}	The inverse standard normal cumulative distribution function
F(d)	A temporal order judgment psychometric function
g	A probability density function in the derivation of the low- and high-threshold binocular summation models
G	The decision function in the general model for temporal order judgment
l	A latency in the left monocular system.
L	A subscript specifying the left monocular viewing condition
m	The mean of the motor component in the two-component serial model of simple reaction time or a subscript which specifies that component.
M	A subscript specifying the monocular viewing condition
p,P	A probability
p_{lin}	The "linearized" proportion of "S1 first" judgments under the assumption that F(d) is a normal cumulative distribution function
PSS	The point-of-subjective-simultaneity
r	An arrival latency in the right monocular system
R	A subscript specifying the right monocular viewing condition
RT	Reaction time
$\overline{RT}$	A mean reaction time
s	An arbitrary sensory state
S	A stimulus, either S1 or S2
SOA	Stimulus onset asynchrony
TOJ	Temporal order judgment

SYMBOL	MEANING
v	A function
VAR	The variance operator
x	An arrival time. Arrival time is equal to the arrival latency for the signals evoked by S1, but is equal to the arrival latency plus the stimulus onset asynchrony for signals evoked by S2.
Y	The two-component arrival latency in the high- and low-threshold summation model derivations.
1	A subscript specifying a S1 presentation
2	A subscript specifying a S2 presentation
α	The criterion in signal detection theory
β	In the probability summation models, this is the confidence measure associated with the judgment based on a particular sensory state
θ	The threshold
λ	A rate parameter for the poisson process in the neural counting and neural timing models
μ	A mean of a distribution. In describing temporal order judgment psychometric functions, it is equivalent to the point-of-subjective-simultaneity.
μ_o	A constant in the low- and high-threshold summation models
σ	A standard deviation. In describing psychometric functions, it is the standard deviation of the underlying distribution whose cumulant is assumed to be the psychometric function.
σ_m^2	The variance of the motor latency component in the two-component serial reaction time model
σ_o^2	A constant in the low- and high-threshold summation models.
ν	A reaction time standard deviation
ω	The biasing factor in the average arrival time model. It is positive and less than one.

SYMBOL	MEANING
ℓ	The likelihhod ratio in signal detection theory

INTRODUCTION

Organization of the Paper

In this paper evidence is presented which suggests that binocular viewing produces a reduction in the variability of visual latency -- the time from the presentation of a light stimulus until it is detected -- over monocular viewing. This paper is divided into eight major sections plus eight appendices. The first section, the Introduction, provides a brief review of experiments which may be classified as binocular summation experiments and puts forth the temporal order judgment experiment as a method of assessing the variability of visual latency: An increase in sensitivity to temporal order is thought to be due to a decrease in visual latency variability. Before any experimental results can be considered, it is necessary to establish a baseline binocular performance which, if exceeded, would imply that some binocular summation in temporal order judgment had occurred. In many types of tasks the mere fact that two eyes receive the stimulus can statistically increase the observer's chance of making a correct judgment even without the operation of specialized binocular processing mechanism or without the interaction of the monocular signals. Some simple probabilistic considerations in these cases usually leads to an approximation of such an inflated baseline performance. If the binocular performance does exceed this probabilistically defined baseline, then binocular summation is said to have occurred. The second section, Probability Summation, treats these probabilistic considerations in detail. The implications of signal detection theory in this situation are considered in the third section, The Likelihood Ratio Observer. The formulae used in this section are derived in Appendix A. In the fourth and fifth sections, specific models for the visual judgment of temporal order are developed. In A General Model for Temporal Order Judgment a frequently used model of temporal order judgment is examined for the monocular viewing case and adapted to provide the prerequisites for analogous models of binocular temporal order judgment. Within the framework provided by this fourth section, several models for the binocular judgment of temporal order are developed in the fifth section, The Models, and are derived in Appendices B through E.

The last three major sections of this paper deal with three experiments. Some data is presented in a Pilot Experiment which does suggest that binocular viewing does sharpen temporal order judgment. Experiment 1 replicates the result of the pilot study under more rigorously controlled circumstances. Experiment 2 sought to provide converging evidence for the results of Experiment 1 and the pilot experiment. The interpretation that increased sensitivity to temporal order results from decreased variability in a sensory latency is dependent upon the assumptions of the general model of temporal order judgment presented in section four. Experiment 2 employs an unrelated measure which includes a perceptual latency, namely simple reaction time, to see if reduced latency variability is again implied by the results.

As mentioned already, Appendices A through E are devoted to the derivation of the various formulae employed throughout this paper. Appendices F and G present the raw temporal order judgment data which were obtained in the pilot experiment and Experiment 1. Finally, Appendix H details the analyses of variance which were employed in Experiments 1 and 2.

Binocular Temporal Summation

The degree to which binocular viewing "enhances" perception relative to monocular viewing and over what dimensions such enhancement occurs has long been an issue in visual research. In addition to the obvious role of two eyes in stereopsis, recent investigations have revealed that binocular performance is superior on detection tasks including detection at absolute threshold (Collier, 1954; Matin, 1962; Thorn and Boynton, 1974), detection of incremental and decremental flashes (Cohn and Lasley, 1976; Westendorf and Blake, 1974), acuity tasks (Kahneman, Norman and Kubovy, 1967), recognition of forms (Eriksen and Greenspon, 1968; Eriksen, Greenspon, Lappin and Carlson, 1966), detection of equal-energy flashes (Westendorf, Blake and Fox, 1972), and flicker fusion thresholds (Baker, 1952; Peckham and Hart, 1960; Sherrington, 1906). Binocular enhancement of brightness (DeSilva and Bartley, 1930; Fry and Bartley, 1933; Levelt, 1968; Stevens, 1967; deWeert and Levelt, 1974) and of the magnitude of the visual evoked response (e.g., Harter, Seiple and Salmon, 1973; White and Bonelli, 1970) have also been reported. Blake and Fox (1973) provide a good review of such experiments which may be catalogued under the rubric of binocular summation. Binocular masking level differences have also been noted in the enhancement of sinusoidal grating detection (Henning and Hertz, 1973, 1977) where the improved detection seems to be the result of a binocular reduction of "noise" and the resulting increase in signal-to-noise ratios.

In this paper, research into binocular interaction along another physico-perceptual dimension, time, is described. It is known that the latency with which a visual stimulus evokes a sensation shows considerable variability as estimated by reaction time studies (e.g., Vicars and Lit, 1975), by temporal order judgment studies (e.g., Gibbon and Rutschmann, 1969; Zacks, 1973), and by physiological studies using single cell recordings from cat retinal ganglion cells (Levick, 1973; Levick and Zacks, 1970). Given such variability, a binocularly presented stimulus must arouse signals in the two monocular subsystems which very seldom have equal perceptual latencies. The simple observation that such binocular observation seems to produce only a single phenomenal onset, rather than two temporally separated onsets, suggests that some binocular temporal combination mechanism is in operation. All previous experiments have tacitly assumed that there is such binocular temporal summation, that occurrences in each eye do combine to a single cyclopean occurrence.

One simple explanation for such binocular temporal summation might be that although the monocular latencies are variable, they are perfectly correlated so that there is never any temporal disparity for the binocular system. Such a rigid explanation is not in order because there is evidence that binocular combination of visual events for many phenomena does not even require that the two retinal images be presented simultaneously. Thus, Langland (1929) showed that a sense of depth could be realized when a scene is alternated from left to right eyes at a sufficient rate. An alternation would result in stereoscopic vision if the two presentations were separated by inter-stimulus intervals of from 18 to 25 ms. Efron (1957) examined this effect in greater detail and found a maximal interval of 6 to 13 ms for fusion with the exact value decreasing slightly with increased flash intensity. More recently, Ogle (1963) has found that disparity cues contribute to the perception of depth even when a delay of greater than 50 ms is introduced between the stimuli provided to separate eyes. Matin (1962) using 2 ms flashes found that for detection threshold not only did binocular summation (i.e., greater performance than predicted by probabilistic considerations alone) exist for delays as long as 100 ms. between stimulation of the separate eyes, but also that the amount of summation as a function of the delay interval was not a monotonic function. Summation for delays of about 40 to 90 ms was greater than at longer and shorter separations. A similar nonmonotonicity was noted in the binocular identification of forms (Eriksen et al., 1966; Eriksen and Greenspon, 1968) with a peak performance at an inter-stimulus interval of about 10 ms. A summation exceeding probabilistic expectations existed for delays up to 50 ms for flashes of duration 25 ms. Battersby and Defabaugh (1969) found that the increment threshold for detection of a 5 ms flash presented to one eye was significantly decreased when paired with a 5 ms subthreshold flash presented to the corresponding area of the other eye for inter-stimulus intervals of less than 50 ms. Smith and Schiller (1966) have shown that both forward and backward masking is obtained when the test stimulus is presented to one eye and a patterned masking stimulus to the other eye when the stimuli are similar. However, when the stimuli are dissimilar, such as when positive and negative flashes are delivered each to separate eyes (Westendorf and Fox, 1974) or when differently shaped stimuli are delivered to each eye (Westendorf and Fox, 1975), they do not produce detection above the level expected on probabilistic grounds alone. Dichoptic stimulation on noncorresponding areas of the two retinae also eliminates summation or reduces it to levels expected on the basis of probability summation

(Battersby and Defabaugh, 1969; Eriksen and Greenspon, 1968; Westendorf and Fox, 1977). Since such experimentally produced magnitudes of inter-stimulus intervals can still result in binocular summation, it is not surprising that the viewing of simultaneously presented binocular flashes are always summated since their arrivals at a central binocular mechanism due to sensory variance would not be expected to exceed the large values cited here.

Temporal Order Judgment and the Variability of Perceptual Latency

The method used in the experiments described in this paper to estimate the variability of visual latency was through temporal order judgments where the observer was asked to determine which of two identical flashes came first. Zacks (1973) argued that temporal order judgment data can be analyzed to provide an estimate of an upper limit on the variability with which a particular flash evokes a sensation, although all information is lost about the absolute magnitude of the latencies because only the latency differences influence temporal order judgment (Sternberg and Knoll, 1973). Zacks' estimates of latency variabilities were not grossly different from latency variabilities determined by single cell recordings from cat retinal ganglion cells (Levick, 1973).

In both the physiological and the psychophysical experiments the latency variability increased considerably for the weaker flashes. Zacks accounts for this dependency of the variability on luminance in terms of the impulse responses of the visual system. First, he notes that the detection response is presumably highly correlated with the consecutive inter-spike intervals of a train of action potentials somewhere in the nervous system. Second, increases in stimulus strength tend to increase spike density above the irregular maintained activity (Kuffler, Fitzhugh and Barlow, 1957). When the luminance of a flash is low, "there are fewer samples per unit time from which to estimate the running average spike frequency ... If there is noise in the system, it is clear that the time at which an increase in activity would be detected would be more subject to fluctuations due to momentary noise activity when the stimulus strength, and hence the increase in spike density, is smaller" (p. 834). Babkoff and Sutton (1963) had shown at an earlier time that the apparent intensity, the loudness, of auditory stimuli influences temporal order judgment. Roufs (1963), Efron (1963), Matteson (1970) and Rutschmann (1973) had examined the effects of varying the intensities of the two stimuli differentially on the point-of-subjective-simultaneity, that is, the stimulus onset asynchrony for which the accuracy of temporal order judgment was at the chance level. Silverstein (1976) later showed that higher sensitivity to temporal order resulted when stimulus intenity was increased in an experiment where the stimuli were the rotary displacement of the observer in an Ames Man-

Carrying Rotation Device and the displacement of a visual stimulus on a cathode-ray tube. In the design of the present experiments, it was explicitly assumed that increased sensitivity to temporal order results from reduced variability in the latency to detect the response evoked by the stimulus and that increasing stimulus intensity is one way to parametrically produce decreases in this variability.

There is, however, some standing controversy in the literature of temporal order judgments as to whether performance on the task can be meaningfully manipulated at all. In particular, the dependency of the level of performance in temporal order judgment on stimulus intensity conflicts with the earlier results of Hirsh and Sherrick (1961). They claimed to show that any stimuli, including visual flashes, must be separated by "about 20 msec" for order to be judged correctly and that this value is independent of the sense modality or modalities used, provided only that the stimuli can be separately identified and named by the observer. This value has been supported by a number of other papers (Hirsh, 1959; Roufs, 1963; Rutschmann and Link, 1964; Sherrick, 1969, 1970) and has even been purported to hold for the binocular viewing condition (Robinson, 1967). Sherrick (1969), however, does concede that flash duration, retinal locus, form quality of images, and specific viewing strategies may reduce this "limen for order" by significant amounts but that the 20 ms value still reflects a fundamental "threshold" for the central timing mechanism. Almost all subsequently reported studies of temporal order have either implicitly or explicitly questioned the validity of this generalization. Rutschmann (1966) contradicted Hirsh and Sherrick's claim by showing that there was a significant dependence of order judgments upon the retinal locus of stimulation, a finding also supported by Oatley, Robertson and Scanlan (1969) and others who showed that there is a slight increase in the variability from extrafoveal as compared to near foveal presentation of small (6') light spots. Westheimer and McKee (1977) found decreasing "thresholds" for temporal order as stimulus separation decreased down to an optimal separation of 2' to 6' where the threshold was found to be 3 to 4 ms in onset asynchrony. Yund and Efron (1974) demonstrated that it is possible to differentiate light flashes whose color components differ in onset by as little as 2 ms. In the tactile modality position of stimulation can also affect apparent simultaneity and in the direction, but not necessarily to the full extent that might be predicted from afferent conduction times (Halliday and Mingay, 1964). The slope of psychometric functions may be such that the interval between the 50 and 75 per cent correct judgments is 2 to 3 times the 20 ms found by Hirsh and Sherrick (see Kristofferson, 1967). Oatley, Robertson and Scanlan (1969) reported that 36 to 98 ms stimulus onset asynchronies were required to increase from 50 to 75 per cent correct. There seems, then, to be sufficient grounds to assume that the onset asynchrony which leads to 75 per cent correct performance is not fixed but varies with conditions.

Although acknowledging this variability many investigators (e.g., Sherrick, 1969; Silverstein, 1976; Sternberg and Knoll, 1973) still assume that the 20 ms "limen for order" represents a very real attribute of the central mechanism. Sternberg and Knoll (1973) suggest that deviations of 75 per cent limens from the 20 ms range should be considered as special cases that are due to external influences, such as peripheral interactions between the stimuli (Babkoff and Sutton, 1963) which may generate extra cues, such as apparent movement (Thor, 1968), and so make the order judgment more accurate, or which may degrade the temporal information in the inputs and so be detrimental to the judgment of temporal order (Robinson, 1967). Westheimer and McKee (1977) examined temporal order judgments both for binocular and dichoptic viewing of closely adjacent stimuli. For stimuli closer than about 12', the binocular threshold for temporal order was appreciably lower than the corresponding dichoptic threshold. They demonstrated that this lower threshold in the binocular case could not have been due entirely to fewer eye vergence errors and so must have been due, at least in part, to some retinal interaction which increased sensitivity to order when both stimuli are viewed by one eye. Yet, even at the 12' separation where binocular and dichoptic thresholds were comparable and where no peripheral interactions were inferred, the threshold for temporal order was only about 10 ms -- still well below the 20 ms level specified by

Hirsh and Sherrick. Further, Westheimer and McKee noted that apparent movement was pronounced with their stimuli, "but that this is not necessary for achieving even the lowest threshold for detection of temporal order. For example, no clear movement supervenes when the two lines, whose relative delay constitutes the stimulus parameter, form a cross; yet for short lines forming a cross the threshold is as low as when two lines are parallel and side by side" (p. 891). This controversy as to the continuity of temporal perception is hardly resolved and is reminiscent in many respects of the controversy which has surrounded the intensive detection threshold in earlier times (Corso, 1963; Green and Swets, 1966). A threshold-less position will be taken by this author throughout the remainder of this paper.

In addition to temporal order judgment, simple reaction time also provides an estimate of a sensory latency and will be investigated in a second experiment. Gibbon and Rutschmann (1965) showed temporal order judgment and simple reaction time produce estimates for the variability in perceptual latency which are very similar and which are influenced by similar parameters. Therefore, if binocular viewing increases sensitivity to temporal order and if the interpretation of temporal order which is given here is correct, then one would expect to observe decreases in reaction time variance with binocular viewing. Several models will be discussed which will relate the sensory arrival times in the temporal order judgment and reaction time experiments.

PROBABILITY SUMMATION

Definition

In any binocular summation experiment which purports to examine mechanisms for the combination of the monocular inputs one must carefully decide on a baseline binocular performance which, if exceeded, can be attributed to a true, neural interaction between the two inputs. In some tasks superior binocular performance is expected due to simple probabilistic considerations alone. An explanation in terms of such considerations is generally referred to as the *probability summation* model. This distinction is a subtle one because probability summation also requires some neural combination of the inputs. The basic idea underlying probability summation is that some increased binocular sensitivity is expected in most tasks because the simultaneous presentation of inputs to both eyes provides the observer with two opportunities to make a detection or otherwise produce a correct response. Binocular summation can be said to occur only if the binocular performance exceeds the level predicted by probability summation.

Two-State Probability Summation Models

The simplest and earliest formulation of the probability summation hypothesis (Pirenne, 1943) can be reformulated as a simple two-state detection model. Assume that the inputs through the left and right systems are statistically independent and result in one of two mutually exclusive states: s_1, the correct perception state in which the observer always responds correctly in the monocular case, and s_2, the incorrect perception state in which the observer would respond incorrectly. There are then four possible states for the binocular viewing situation. The probability of a correct response, p_c, on a given trial is then given by the sum of the probability of making a correct response given that the observer is in that state weighted by the probability that the observer is in that state. That is,

$$\begin{aligned} p_c = {} & p_1 p_R p_L + p_2(1 - p_R)p_L \\ & + p_3 p_R(1 - p_L) + p_4(1 - p_R)(1 - p_L) \end{aligned} \tag{1}$$

where p_R and p_L are the probabilites of being in s_1 for the right and left systems, respectively, and p_i $(i = 1, 2, 3, 4)$ is the conditional probability that the observer makes the correct decision given that the two states have occurred in the monocular subsystems.

The probability summation model for simple detection experiments asserts that binocular stimulating conditions are logically the same as though one eye received two successive stimulations and the two corresponding states are used by a rule that states that a correct perception (i.e., s_1) on either the first or the second presentation or on both presentations will lead to the correct response with a probability of 1.0. Two incorrect perceptions, on the other hand, lead to the correct response with probability 0.0. This is equivalent to setting $p_1 = p_2 = p_3 = 1.0$ and $p_4 = 0.0$ in equation 1 which then reduces to

$$p_c = p_R + p_L - p_R p_L \tag{2}$$

which predicts a higher proportion of correct responses on the detection task.

A similar two-state model can be applied to the judgment of temporal order by identifying the correct response state with a "Stimulus 1 (S1) preceded stimulus 2 (S2)" state and the incorrect response state with the alternative "S2 preceded S1" state. Then, given that S1 did, indeed, come first, equation 1 does express a two-state probability summation model for temporal order judgment. However, there is a difference in the manner of assignment of the values to p_1, p_2, p_3 and p_4. The observer should again respond with certainty ($p_1 = 1.0$) that "S1 preceded S2" when both monocular systems are in s_1 and with certainty ($p_4 = 0.0$) that "S2 came first" when both subsystems are in s_2. The values of p_2 and p_3, on the other hand, must certainly differ from 1.0. In the detection experiment, the occurrence of a single s_1 was assumed to be sufficient to lead to a correct detection for the binocular system. In temporal order judgment, if one subsystem reports that "S1 came first" and the other reports that "S2 came first," then no decision can be made with certainty. If the observer makes an unbiased guess in this situation, then p_2 and p_3 equal the a priori probability of 0.5 and equation 1 becomes

$$p_c = 0.5p_L + 0.5p_R \tag{3}$$

which is simply the arithmetic average of the two monocular probabilities of correct response. This can hardly be considered to be an "enhancement" of performance through probability summation.

Two-state models for the performance on the binocular detection task like the one described by equation 2 are not generally used to estimate the probability summation baseline for the evaluation of experimental results because they tend to overpredict performance from two independent chances to perceive by giving double weight to the guessing component (Blake and Fox, 1973). Eriksen (1966) has hypothesized that the probability of a correct response on a single monocular observation is composed of at least two parts: (a) trials on which the observer correctly perceived and responded correctly, and (b) trials on which the observer did not perceive correctly but did guess correctly. These two parts make up each monocular probability. When the two monocular probabilities are combined by equation 2 to predict the binocular probability, the guessing component is included in each. But during actual binocular viewing, the observer would guess only if he failed to perceive with both left and right eyes. Eriksen's model will not be elaborated here, but it does tend to provide a less inflated baseline for binocular detection tasks.

Several authors (Braddick, 1972; Guth, 1971) have objected to the use of all such two-state models of probability summation in detection experiments because their application is appropriate only within a very restricted and probably incorrect theory of visual intensive threshold. (The problem is covered much more extensively in Corso, 1963, and in Green and Swets , 1966). The use of models with guessing states to estimate probability summation levels for binocular temporal order judgment also implies a classical threshold. Hirsh and Sherrick's (1961) dictum discussed in the introduction to this paper may be interpreted as such an attempt to apply a classical threshold concept to order perception: their 20 ms value was interpreted as reflecting properties of a central time organizing mechanism. Variability in the organism would be assumed to produce "momentary thresholds" for stimulus onset asynchronies less than 20 ms and to reduce the effects of some asynchronies greater than 20 ms to subthreshold levels thus producing a smoothly graded psychometric function -- the 75 per cent correct level has no significance without such assumptions. Perceptual moment theory (Stroud, 1949) also postulates a classical threshold for temporal order. According to this theory, psychological time is quantized into approximately equal periods (perceptual moments) during which only the average change of stimulus parameters is known. If the order of two events is to be discerned, their responses must lie in different perceptual moments; judgments about the order of stimuli whose responses lie in the same moment are assumed to be governed by chance alone. Variability in temporal order judgment (in addition to sensory variability in some versions) for asynchronies less than the duration of a moment occurs in this model because the signals are assumed to arrive in random phase with the sampling frequency which defines a moment. Thus even for very close intervals the first signal will occasionally lie at the end of one moment and the second at the beginning of the next. Perceptual moment theory, naturally, has important implications for the nature of models of temporal order judgment (Allan, 1975; Kristofferson, 1967; Oatley, Robertson and Scanlan, 1969; Sternberg and Knoll, 1973). It will, however, not be considered further here. In this paper, the perception of order will be treated as if it were continuous and without a threshold. To the extent that this assumption is correct, models including guessing states are not applicable to the judgment of temporal order. At least in the two-state model described here, there is no reason to expect guessing in either of the two states.

Multistate and Integration Models of Probability Summation

Eriksen (1966) has also described a multistate model for detection in which it is assumed that observers can distinguish several different subjective states of confidence about their judgments on stimulus presentation. The multistate model expresses the joint occurrence of the various monocular perceptual states as a linear combination of the proportion of times such a particular state occurs weighted with the performance measure associated with each of the states. So, for the monocular experiment, performance can be expressed by

$$p_R \left\{ \text{"S1 came first"} \right\} = \sum_{i=1}^{n} \beta_i p_{R,i} \tag{4}$$

$$p_L \left\{ \text{"S1 came first"} \right\} = \sum_{j=1}^{n} \beta_j p_{L,j} \tag{5}$$

where $p_{R,i}$ and $p_{L,j}$ are the probabilities of the occurrence of the ith state for the right and the jth state for the left systems, respectively, for the given stimulus presentation and β_i is the confidence measure which is equal to the conditional probability that the ith perceptual state was produced by a S1 first presentation, or equivalently, the conditional probability that given that the observer is in the ith state he responds that S1 came first. Then for the binocular performance, Eriksen's conceptualization would produce

$$p_B\left\{\text{"S1 came first"}\right\} = \sum_{i=1}^{n}\sum_{j=1}^{n} f(\beta_i,\beta_j)\,p_{R,i}\,p_{L,j} \quad (6)$$

where the function $f(\beta_i,\beta_j)$ relates the monocular confidence measures to a binocular confidence measure.

Eriksen estimated the measure β_i by having observers rate states of confidence and by comparing the ratings to their performance. He found that the accuracy of performance covaries with the confidence rating of the observer; the best performances as indexed by per cent correct are associated with the most confident states and as performance declines, confidence decreases. There is then strong support for a multistate notion for threshold.

However, in order to implement equation 6, one must decide on an appropriate function $f(\beta_i,\beta_j)$ to predict the binocular confidence measure. There is, however, no *a priori*, model-free, probabilistic basis for the form of $f(\beta_i,\beta_j)$. When one postulates a specific $f(\beta_i,\beta_j)$, one is actually constructing a small theory about how binocular events interact. This is hardly in the spirit of probability summation! Some combination rules may lead to increased binocular sensitivity and others may lead to no change. Examples of rules which lead to no expected difference between monocular and binocular performances include: (1) if the observer's binocular confidence measure is equal to a weighted or unweighted average of the two monocular confidences, or (2) if decisions are made independently at the level of the monocular channels and the binocular system guesses if these decisions do not agree. Both of these rules lead to the same prediction made by equation 3. Eriksen (1966) made the assumption that the observer's reports will be based upon the most sensitive of the two states; more sensitive performance with the binocular observation would be expected due to the increased likelihood of entering a more sensitve state on one of the two chances. Eriksen found close agreement between the binocular performance predicted by this model and the observed increase in detection when the two inputs were forced to be independent by temporal or spatial separation.

To summarize, the prediction of the multistate model depends heavily on the particular formulation of the binocular confidence measure. To be sure, most any observed binocular effect could be modeled by the assignment of an appropriate $f(\beta_i, \beta_j)$. For this reason, no multistate model will be offered to estimate the baseline performance expected on probabilistic grounds alone. In this paper, baseline performance will be simply that the lack of neural interaction of any type would lead to a binocular performance which is not significantly different than the performance on either monocular condition.

More recent work has extended into formulations of probability summation models based on an infinite number of perceptual states. Blake and Fox (1973) term such models *integration models*. Both integration models and multistate models are similar in that neither assumes a theoretical, dichotomous sensory threshold below which responses are governed by pure guessing. In both formulations some *a priori* function $f(\beta_i,\beta_j)$ is necessary to describe the combination of monocular confidences. In fact, the general integration model can be simply treated as a limiting case of the multistate model as the number of possible states increases beyond bounds and the probability of occurrence of any particular state approaches zero.

A Baseline for Independence

One important choice of $f(\beta_i,\beta_j)$, the function which assigns a binocular confidence from the two monocular confidence measures, is that function which maximizes the binocular sensitivity; Some $f(\beta_i,\beta_j)$ in equation 6 must lead to a maximal value for p_B. Such a choice would provide a baseline for evaluating binocular versus monocular performances which may be called a "baseline for independence." Performance in excess of this level would imply that the states were not independent. Dependence would also imply neural interaction. However, performance below this baseline leads to no such inferences -- the states may or may not be independent and there may or may not be neural interaction; Neural interaction does not imply dependence.

This "baseline for independence" differs from a level defined by probability summation. The baseline determined by probability summation would typically be less than the

the baseline for independence. The probability summation level yields no information about dependence. If there is binocular summation, the performance should exceed the baseline established by probability summation. If there is binocular summation and, further, the states are independent, then the binocular performance should be bounded from below by the probability summation level and from above by the "baseline for independence." Binocular performance below the level predicted by probability summation must be interpreted either (1) to mean that a neural interaction which produces decreased sensitivity in the binocular case is operating, or (2) to mean that an improper model was used to establish the probability summation level. An improper model for probability summation could arise if, for example, a model for the baseline for independence is confused with a model for probability summation.

THE LIKELIHOOD RATIO OBSERVER

Guth (1971) and Braddick (1972) argue strongly that probability summation levels for comparison of monocular detection performance to binocular performance should be based on the theory of signal detectability (Green and Swets, 1966). The probability summation baseline has been determined in this way in most recent binocular summation experiments (e.g., Craig, Colquhoun and Corcoran, 1976; Loveless, Brebner and Hamilton, 1970; Westendorf, Blake and Fox, 1972; Westendorf and Fox, 1974, 1975, 1977). However, it must be emphasized that the hypothesis of a likelihood ratio observer is only a hypothesis of how observers handle sensory information and not an atheoretical statement of binocular performance expected on "purely probabilistic grounds alone." Signal detection theory does supply one solution to the dilemma confronting the formulation of infinite (or multiple) state models of how sensory information is combined on multiple presentations by assuming that the observer makes his decision according to the likelihood ratio, $\ell(s)$, when he is in the perceptual state s; When $\ell(s)$ is greater than some criterion value, then the observer says that S1 came first, otherwise he says that S2 came first. The likelihood ratio for the left monocular system,

$\ell_L(s_L)$, can be defined as

$$\ell_L(s_L) = \frac{P_L(s_L | \text{"S1 came first"})}{P_L(s_L | \text{"S2 came first"})} \ . \tag{7}$$

In words, the likelihood ratio is the probability that the state s_L arose from a "S1 first" stimulus situation in the left system, divided by the probability that s_L arose from a "S2 first" presentation. Mutatis mutandis, an analogous equation could be demonstrated for the likelihood ratio of the right monocular system, $\ell_R(s_R)$.

In the binocular case, a decision would similarly be based on a likelihood ratio, $\ell_B(s_L, s_R)$, which considers the states s_L and s_R aroused in the left and right systems. This ratio is defined by

$$\ell_B(s_L, s_R) = \frac{P_B(s_L, s_R \mid \text{"S1 came first"})}{P_B(s_L, s_R \mid \text{"S2 came first"})} \ . \tag{8}$$

This model does predict an enhanced sensitivity for binocular performance. Specifically, signal detection theory predicts that the overall sensitivity (d') resulting from two independent observations as (Blake and Fox, 1973)

$$d'_{bin} = \sqrt{d'^2_{left} + d'^2_{right}} \tag{9}$$

which is a sensitivity in all cases greater than either monocular sensitivity.

Measures of Sensitivity to Temporal Order

It would be desirable to express this increased binocular sensitivity to temporal order directly as the psychometric function, F(d). This function relates the proportion of times that the observer responded that S1 preceded S2 as a function of the stimulus onset asynchrony, d. This is the form in which the temporal order data will be presented throughout this paper.

Positive values of d correspond to trials where S1 preceded S2 and negative values to trials where S2 preceded S1. F(d) has the form of a nondecreasing function on the d axis bounded by 0.0 for large negative asynchronies and by 1.0 for large positive asynchronies. It is usually characterized by two parameters, the point-of-subjective-simultaneity and the threshold for temporal order, as shown in figure 1. The point-of-subjective-simultaneity is the stimulus onset asynchrony for which F(d) is equal to the *a priori* probability of 50 per cent. At this asynchrony the temporal order is least discriminable and is usually taken to reflect the difference in latencies between responses to S1 and S2. It is, however, also potentially affected by certain response biases. The point-of-subjective-simultaneity serves to "locate" the psychometric function on the d axis.

The threshold for temporal order is an (inverse) measure of sensitivity. It is often defined as the stimulus onset asynchrony which leads to 75 per cent "S1 came first"

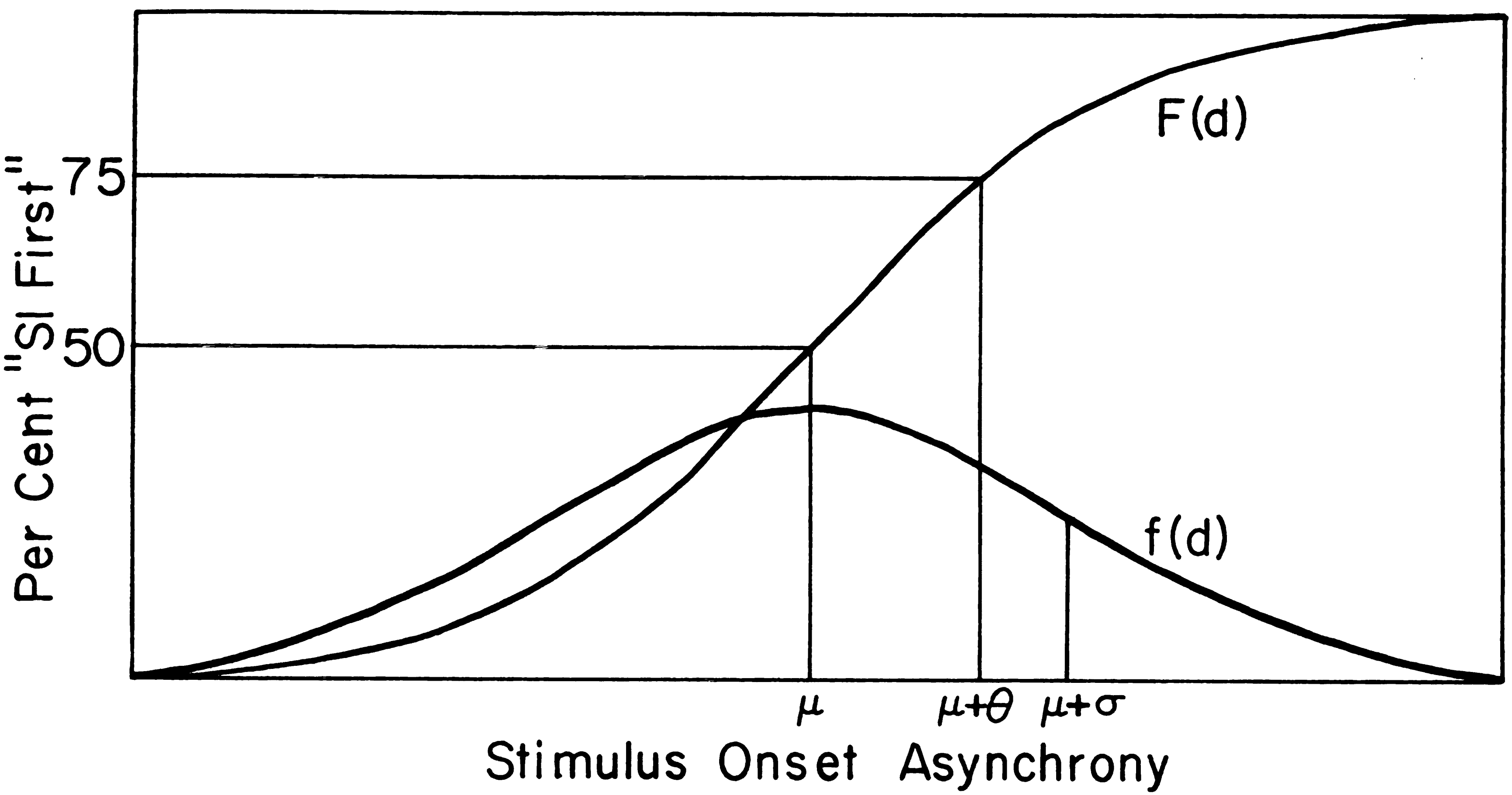

Figure 1. The relationship between measures of sensitivity to temporal order

performance. This definition assumes that the point-of-subjective-simultaneity is equal to the point of physical simultaneity ($d = 0$). A more generally acceptable definition for this threshold is the difference between the asynchrony for 75 per cent performance and the point-of-subjective-simultaneity. The threshold for temporal order is a measure of the "flatness" of F(d); the "steeper" F(d) is, the smaller is the threshold.

In this paper a different but equivalent pair of parameters will be used to describe F(d). Since F(d) is a sigmoidal function bounded by 0.0 and 1.0, it may be reasonable to approximate it with a cumulative normal distribution function characterized by a mean, μ, and a variance σ^2, as shown in figure 1. (Even if F(d) is not exactly normal, such an approximation tends to be very close for real data). The mean, μ, is equivalent to the point-of-subjective simultaneity; both are the stimulus onset asynchrony for which F(d) equals 50 per cent. The standard deviation, σ, is proportional to the threshold for temporal order. The threshold is equal to 0.675 times σ.

Because of the assumption of normality, the parameters μ and σ completely characterize F(d). Thus if the likelihood ratio model can be used to estimate μ_B and σ_B for the binocular F(d) as a function of μ_L, μ_R, σ_L, and σ_R for the monocular cases, then we will have succeeded in expressing the binocular sensitivity given by equation 9 in terms of the psychmotric function, F(d).

This general solution relating equation 9 to the parameters of the psychometric function is sufficiently complex that it is not presented here (see Appendix B). A simpler special case that arises when the performance by the two monocular systems are identical results in the prediction that,

$$\mu_B = \mu_L = \mu_R \tag{10}$$

where μ_B, μ_L and μ_R are the points-of-subjective-simultaneity for the binocular, left-eye and right-eye psychometric functions, and that

$$\sigma_B{}^2 = 0.5\ \sigma_L{}^2 = 0.5\ \sigma_R{}^2 \tag{11}$$

where σ_B, σ_L and σ_R are the standard deviations of the normal distributions whose cumulants are assumed to be the psychometric functions for the binocular, left-eye and right-eye viewing situations. In other words, the point-of-subjective-simultaneity for the binocular function is expected to be the same as those for the monocular functions, but the variance underlying the binocular psychometric function is expected to be only half of that of the monocular functions. Thus, a considerable increase in sensitivity to temporal order is predicted by signal detection theory.

The Criterion

Throughout this entire section only the case where increased binocular sensitivity is attributed to increases in d', and not to changes in α, the criterion, was considered. Braddick (1972) suggests that α many play an important role in binocular summation detection tasks: A binocular improvement may be accompanied by an increased criterion to reduce false alarms while maintaining a high hit rate. In the temporal order model presented here, an overall criterion change should effect only the point-of-subjective-simultaneity and not the "slope" of the psychometric function. The probability of large criterion differences between monocular and binocular decision criteria can be minimized by randomizing monocular and binocular viewing conditions from trial-to-trial so that the observer has no knowledge of the condition before the stimuli are presented.

However, other models suppose that the criterion shifts from trial-to-trial in the binocular viewing condition. Such a state of affairs could result in an apparent increased binocular sensitivity. In a detection task, one input could, for example, serve to "alert" the decision mechanism to the second input (Loveless, Brebner and Hamilton, 1970; Nickerson, 1973) and so improve performance by altering the decision criterion for the response to the second input. Or, one channel may affect "prior entry effects" in the other (see Sternberg and Knoll, 1973). It is not clear how such "alerting" should affect the decision criterion, but since this situation destroys any independence between the inputs, it will not be treated in this paper.

A GENERAL MODEL FOR TEMPORAL ORDER JUDGMENT

Variability in the Sensory Channels

Before continuing with the discussion of models for the binocular judgment of temporal order, a general model for the monocular judgment of temporal order will described here to add some theoretical substance to the models which follow. This model has been called the independent channel model and is discussed at length by Sternberg and Knoll (1973). The two stimuli, S1 and S2, in the temporal order judgment experiment are assumed to generate signals in each of two independent "channels" which relay the signals to the appropriate brain regions where they converge, in some sense, and where the binary decision, "which came first?" is made.

Although a particular stimulus onset and subsequent retinal impingement may be nearly perfectly specified in time, the onset of the response signal which results will not be. Rather, this physiological onset is only probabilistically related to the time of impingement presumably because of ongoing noise in the visual system and will, therefore, display variance. Further, it is possible that higher order response signals will be initiated with even greater variance at each synaptic level. Variability in the onset of the physiological visual response has been noted particularly at the level of retinal ganglion cell axons in the cat retina by Levick and Zacks (1970) and Levick (1973). In these cells there is a vigorous irregular ongoing discharge (Kuffler, Fitzhugh and Barlow, 1957) upon which responses are superimposed. In detecting the presence of a light pulse, it is assumed that the human visual system is faced with a statistical decision -- detecting when an increased spike frequency has occurred in a noisy sensory channel. So, in addition to variability in the transduction of photic energy and conduction of neural activity, we must also include a variable latency at the decision locus to detect the onset of the signal. These combined latencies -- a transduction-conduction latency and an additional detection latency -- shall be referred to as the <u>arrival latency</u> of the signal and will be denoted by r and l for the signals arising from the right and left systems, respectively.

Arrival latency is a random variable from the stimulus onset time. Let $x_{L,1}$ and $x_{L,2}$ represent the arrival times at the decision locus of the signals resulting from S1 and S2, respectively, for the left monocular system. These are represented schematically in Figure 2 as samples from the probability density functions $f_{L,1}(l_1)$ and $f_{L,2}(l_2)$, which give the point probability that S1 and S2 gave rise to signals in the left system with arrival latencies l_1 and l_2, where arrival time is equal to the arrival latency plus the

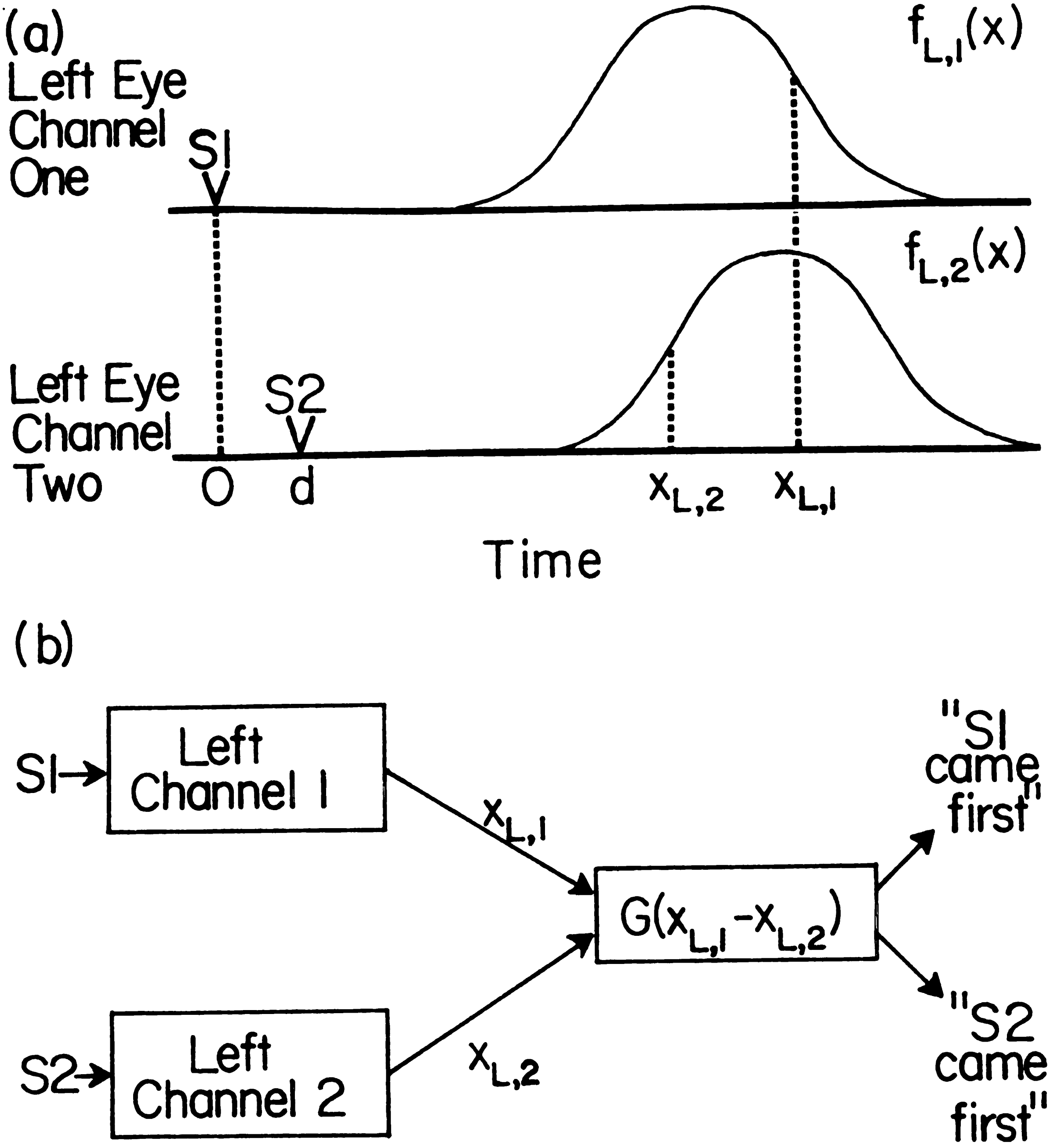

Figure 2. Two schematic representations of the monocular temporal order judgment showing (a) the statistical nature of the arrival times, and (b) the general organization of the model

time of the stimulus onset. Temporal order may then be inferred at the decision locus from the relative arrival times of the two signals. In particular, the arrival time difference, $x_{L,1} - x_{L,2}$, contains all information necessary for the decision and so is conceived as the only input to the decision mechanism. In the independent channel model, the arrival time difference is analogous to state in the previous models.

The Decision Mechanism and the Characterization of the Psychometric Function

The decision process for the left system can be characterized as a function, $G_L(x_{L,1} - x_{L,2})$, of the arrival time difference which specifies the probability that the observer will respond that S1 preceded S2. Sternberg and Knoll (1973) have described several decision rules which may operate. The one that will be employed throughout this paper is what they called the *deterministic decision rule* because it postulates that neglible variability in the judgment of temporal order arises from the decision mechanism itself. Rather all variance observed in the judgments is a result of the variable latency of signals in the sensory channels. Even if there is some variability in the decision process, the use of a model which employs a deterministic decision rule to infer sensory variability from experimental data will yield an approximate *upper bound* on the sensory variability. Such a measure is useful even if a probabilistic decision rule is more appropriate. A probabilistic decision rule could result, for example, if there were a threshold for temporal order or if there are perceptual moments (see Sternberg and Knoll, 1973).

The deterministic decision rule, specifically, is that the observer will say that S1 came first if and only if $x_{L,1}$ arrived at the decision locus before $x_{L,2}$. Otherwise, the observer will say that S2 came first -- he never guesses. Mathematically,

$$G_L(x_{L,1} - x_{L,2}) = \left[\begin{array}{ll} 1.0 & \text{if } x_{L,1} - x_{L,2} < 0.0 \\ & \\ 0.0 & \text{if } x_{L,1} - x_{L,2} > 0.0 \end{array}\right. \tag{12}$$

What has been presented thus far is sufficient to provide a simple mathematical interpretation of the psychometric function, $F_L(d)$, which represents the probability that the observer will say that S1 came first as a function of d, the stimulus onset asynchrony. Given that $F_L(d)$ can be approximated by an ogive, Sternberg and Knoll show that the independent channel model for temporal order judgment relates $F_L(d)$ to the arrival latencies according to

$$F_L(d) = P\left\{ l_1 - l_2 \leq d \right\} \tag{13}$$

where the l_1 and l_2 are the arrival latencies for the left-eye of signals evoked by S1 and S2, respectively. If $f_{L,1}(l_1)$ and $f_{L,2}(l_2)$ are assumed to be normal with means $\mu_{L,1}$ and $\mu_{L,2}$ and variances $\sigma_{L,1}^2$ and $\sigma_{L,2}^2$, respectively, then the parameters describing $F_L(d)$ are given by

$$\mu_L = \mu_{L,1} - \mu_{L,2} \tag{14}$$

and

$$\sigma_L^2 = \sigma_{L,1}^2 + \sigma_{L,2}^2 \tag{15}$$

For the right monocular system, equations similar to equations 12 through 15 can be demonstrated. If the arrival latencies, r_1 and r_2, for the signals aroused by S1 and S2 in the right monocular system are assumed to be distributed normally with means $\mu_{R,1}$ and $\mu_{R,2}$ and variances $\sigma_{R,1}^2$ and $\sigma_{R,2}^2$, respectively, then $F_R(d)$, the psychometric function describing temporal order judgment in the right monocular system can be characterized by the parameters

$$\mu_R = \mu_{R,1} - \mu_{R,2} \tag{16}$$

and

$$\sigma_R^2 = \sigma_{R,1}^2 + \sigma_{R,2}^2 \tag{17}$$

Up to this stage a mathematical system has been described which explains the monocular temporal order judgment in terms of the independent arrivals of signals evoked by S1 and S2 in either, but not both, eyes. This basic framework can be extended to the binocular viewing situation as diagrammed in Figure 3. Now, the binocular decision is to be based on four independent arrival times evoked by S1 and S2 in the two eyes. It is the operation of this four-input, binocular decision mechanism that the remainder of this research addresses.

The Assumption of Independent Channels

Models based on non-independent inputs will not be treated in this paper even though there is some evidence which suggest non-independence. Some of this evidence is reviewed below.

Experiments with even heteromodal stimuli indicate that preceding signals may delay a response as in studies of "psychological refractory period" (Smith, 1967) and that following signals facilitate it as in studies of "intrasensory facilitation" (Bernstein, Rose and Ashe, 1970). Such effects could serve to sharpen temporal discrimination beyond the level produced by the central mechanism alone. Nickerson (1973) favors a cuing model to explain intrasensory facilitation in which one modality serves to alert the other to the presence of a possible signal. When the alerting modality "decides" that a signal has been present-

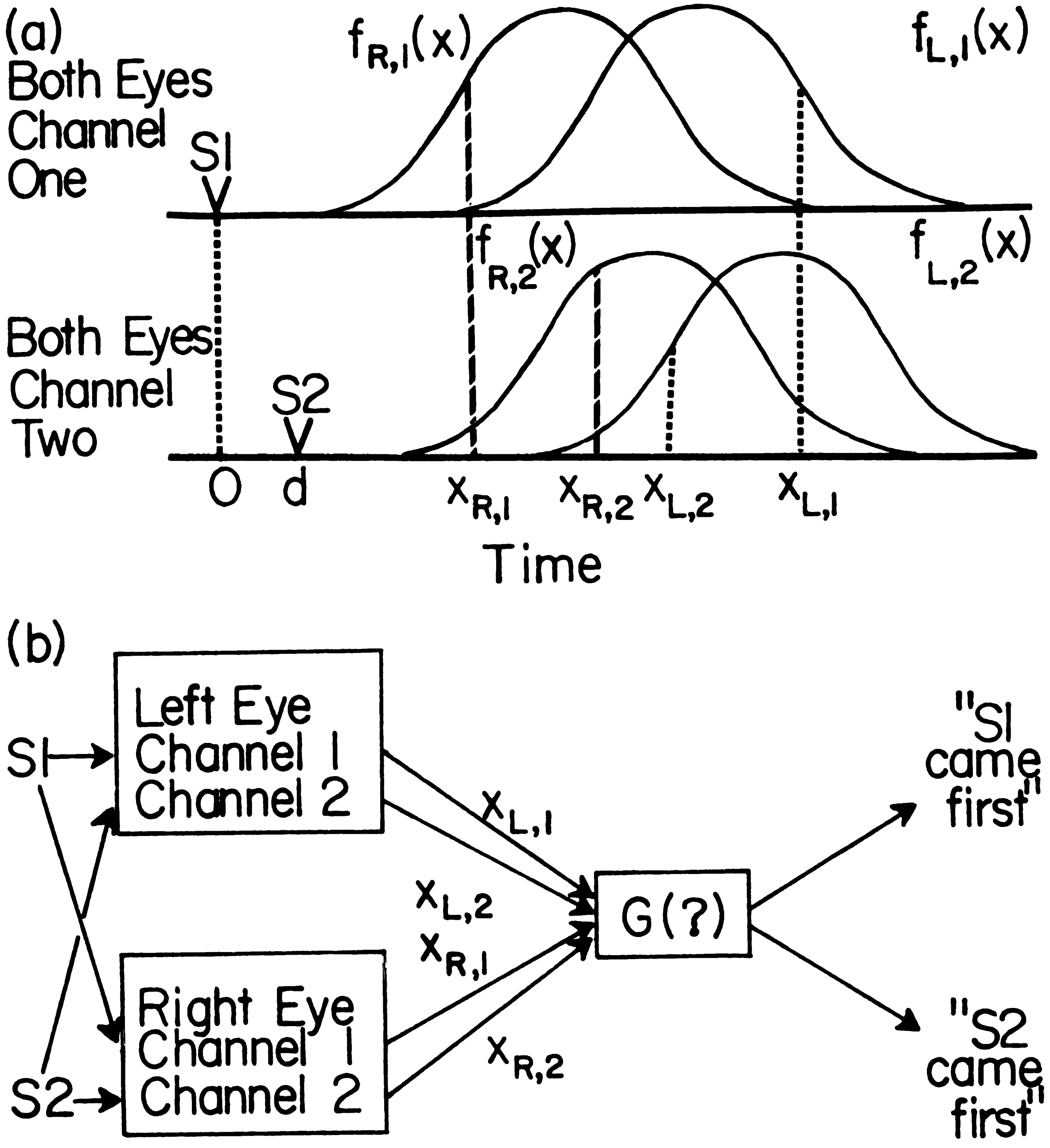

Figure 3. Two schematic representations of the basis of the binocular temporal order judgment models showing (a) the statistical nature of the arrival times, and (b) the general organization of the models

ed, this increases the probability of a detection in the cued modality by causing a reduction in the latter's criterion. On the other hand, when the alerting modality decides that a signal has not been given, then the criterion of the cued modality is increased so that the detection response is less likely. The sensitivity (d') is assumed to be unaltered in the cued modality and there is no correlation among inputs. Craig, Colquhoun and Corcoran (1976) put forth support for a model for dual mode detection which does assume truly correlated variables. This model resembles equation 2 with an additional subtractive term which contains the intermodality correlation. It differs from equation 2 in the extent to which it predicts the coincidence rather than conflict between the two modalities. Eijkman and Vendrik (1965) concluded that a high correlation between the internal noise in the auditory and visual systems exists for duration discrimination but not for intensity discrimination. Robinson (1967) and Westheimer and McKee (1977) both found that dichoptic temporal order judgments were substantially different than when one eye received both stimuli, although they disagreed on the direction of the difference. Both concluded that some interaction at the periphery was responsible for the difference. Even if the stimuli are viewed monocularly at a spatial separation which should minimize such interaction, Zacks (1973) notes the potential for correlated responses if the state of the entire retina were to vary over time.

In spite of the preceding discussion, the assumption of independent inputs to the level of the decision mechanism will be retained in order to simply the formulation of a workable model unless the independent channel models fail drastically to account for the data. These independent channel models could fail if, for example, the observed increase in binocular over monocular sensitivity to temporal order exceeded the "baseline for independence" discussed in the section on probability summation.

THE MODELS

The Average Arrival Time Model

In binocular temporal order judgment, one decision rule that an observer might apply to the four independent arrival times is simply to average the two arrival times referenced to each stimulus, S1 and S2, and then use these average arrivals in precisely the same way that the monocular arrivals were used according to the independent channel model: If the average arrival of the two signals arising from S1 is smaller (sooner) than the average arrival time of the S2 signals, then the observer would say that S1 came first. Otherwise he would say that S2 came first. This decision rule leads to what will be called the Average Arrival Time Model. This model leads to a characterization of the binocular psychometric function with mean (or point-of-subjective-simultaneity), μ_B, and variance, σ_B^2, given by (see Appendix B)

$$\mu_B = \omega\mu_R + (1 - \omega)\,\mu_L \qquad (18)$$

and

$$\sigma_B^2 = \omega^2\sigma_R^2 + (1 - \omega)^2\sigma_L^2 \qquad (19)$$

That equation 19 does represent a binocular increase in sensitivity to temporal order can be shown by considering the special case where $\omega = 0.5$ and $\sigma_R^2 = \sigma_L^2$. In this case equations 18 and 19 reduce to

$$\mu_B = 0.5\,\mu_R + 0.5\,\mu_L \tag{20}$$

and

$$\sigma_B^2 = 0.5\,\sigma_L^2 = 0.5\,\sigma_R^2 \tag{21}$$

That is, the variance observed in the binocular psychometric function is half the variance observed in the monocular cases. Notice that this is the same level of performance predicted by the signal detection theory reasoning (see equation 11). Thus, by averaging arrival times, an observer can perform like a likelihood ratio observer at least for the case when performances in each monocular system are very similar.

The claim is made that this level of performance, in fact, represents an upper limit for the binocular effect given that the inputs to the decision mechanism are independent. If this is true, then performance in excess of equation 21 would suggest that the inputs to the decision mechanism are not independent. Performance less than that predicted by equation 21, however, does not guarantee independence. The acceptance of equation 21 as a "baseline for independence" hinges on the notion that the decision rule of the averaged arrival time model represents the best statistic for estimating the difference in arrival time means and that knowledge of this difference leads to greater accuracy in temporal order judgment.

The Most Discrepant Arrival Time Model

Another decision rule for binocular temporal order judgment, the *Most Discrepant Arrival Time Model*, is based on the same sort of reasoning that Eriksen (1966) used in his multistate model of successive observation. Eriksen hypothesized that an observer's report will be based solely on the more confident of the two independent perceptual states. The abstract notion of state in Eriksen's framework is analogous to the arrival time difference in the independent channel model of Sternberg and Knoll (1973). It is also reasonable to identify the magnitude of the arrival time difference as the analog of confidence. That is, when the arrival times are separated by a relatively large interval, the arrival time difference is large and the observer would make his decision with more confidence than when the arrivals occur in close temporal contiguity where the percept should approach one of apparent simultaneity.

The most discrepant arrival time model supposes that although information is available through both monocular systems, information from only one eye is used on any given trial to make a judgment. The eye used may vary from trial-to-trial, the selection being the eye with the most discrepant arrival times and, hence, the most confident state. Once an eye was selected, the arrivals would be used in the same way as they were in the monocular temporal order judgment. If we consider the simple case where $\sigma_L = \sigma_R$, then it can be shown (see Appendix C) that the parameters, μ_B and σ_B, of the psychometric function describing the binocular temporal order judgment are given by equations 20 and 21. That is to say, the predicted binocular performance of the most discrepant arrival time model is identical to the predictions of both the average arrival time model and the signal detection model. This surprising consequence suggests that these models are equivalent in some mathematical sense, at least for the case when the responses from both eyes are similar.

The Low-Threshold Binocular Summation Model

Of the models presented thus far, none have explicitly hypothesized a direct combination of signals at some neural locus to yield a truly binocular system. Rather, mechanisms of indirect combination have been described in which information was abstracted from the signals and used by a relatively complex information processing system. If the binocular combination of signals produced by stimuli presented to corresponding retinal locations in the two eyes can be characterized as converging upon a common set of neural units and if the onset of the binocular system is dependent upon the relative arrival times of the two signals, then a model for the binocular temporal order judgment can be postulated which is very different from those presented up to this point.

A low-threshold binocular summation model can be simply expressed as follows (see Figure 4): Suppose that the cell or set of units which receives inputs from both monocular systems will respond to its inputs at a threshold which is low enough that a single monocular input in isolation is sufficient to excite it. Suppose also that the onset of the binocularly tuned system's signal will be related, at least to a first approximation, to the arrival of this suprathreshold monocular signal. Then the onset of the binocular response when both eyes observe the stimulus will correspond to the first of the two arrival times. A statistical consequence of this arrangement (see Appendix D) is that this binocular latency will show reduced variability. Manipulation of the mathematical structure implied by such a model (see Appendix D) leads to the characterization of binocular temporal order judgment with the parameters

$$\mu_B = \mu_L - 0.5642(\sigma_{L,1} - \sigma_{L,2}) + \mu_o \tag{22}$$

and

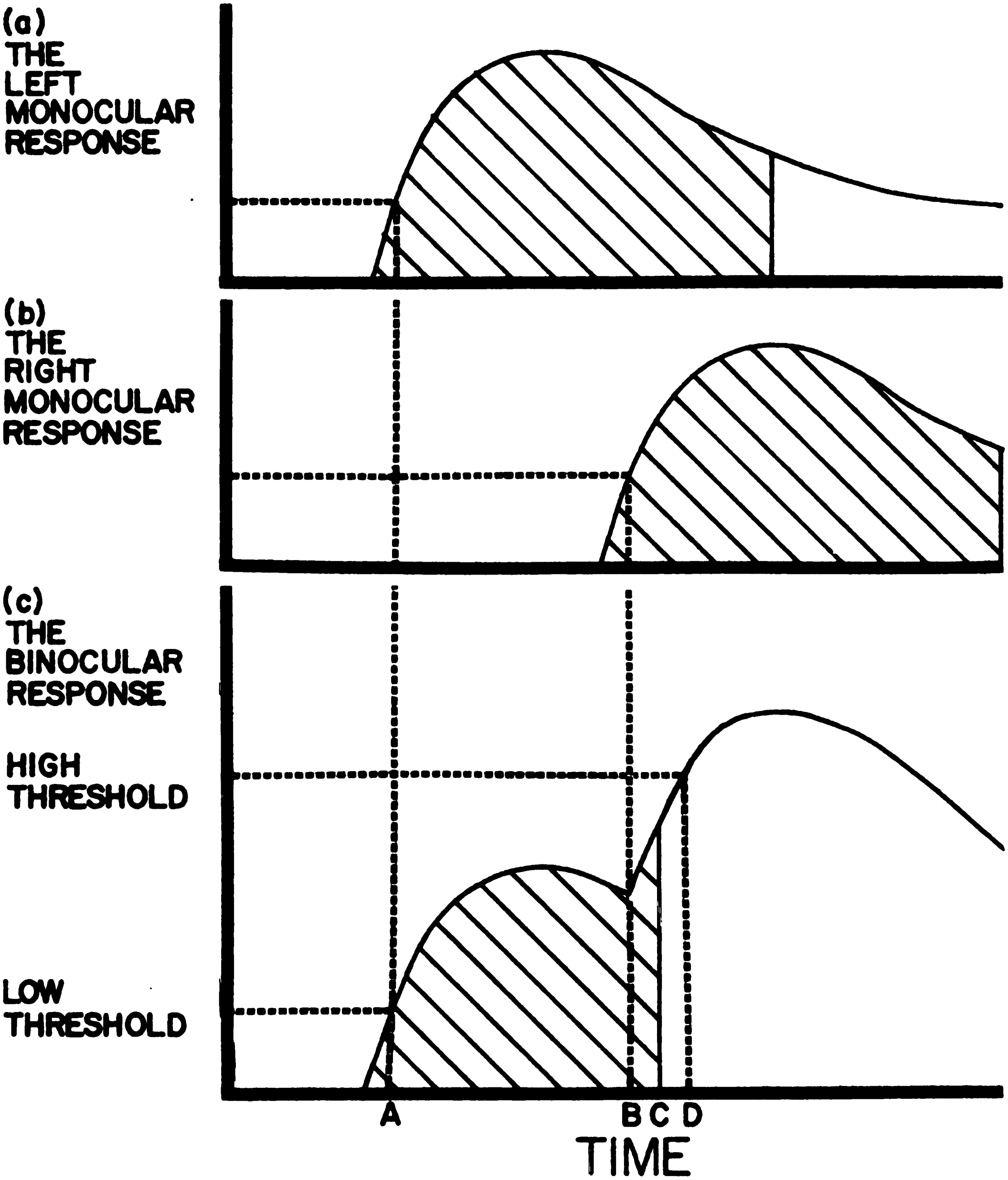

Figure 4. The relationship between the arrival of the first signal (A), the arrival of the second signal (B), the prediction of the low-threshold summation model (A), the high-threshold summation model (D), and Ueno's (1976) model applied to the binocular temporal order judgment (C)

$$\sigma_B^2 = 0.6817\,\sigma_L^2 + \sigma_o^2 \tag{23}$$

where μ_o and σ_o^2 are constant and probably very small. If σ_o^2 is neglible, then equation 23 predicts a considerable increase in the binocular sensitivity to temporal order, but not as great as that predicted by the preceding models. This should not seem surprising if there is any substance to the claim made earlier that these other models represent the best possible performance which could be expected given that the decision is based on four independent inputs.

The High-Threshold Binocular Summation Model

If the binocular combination described above for the low-threshold summation model is again assumed but modified so that the binocularly tuned unit or set of units has a higher threshold, then we have outlined the basis for a *high-threshold binocular summation model*. For the high-threshold summation model it is assumed that a single monocular input is not sufficient to excite the binocular unit to threshold but does have a facilitory effect so that the arrival of the second signal is sufficient to "turn on" the binocular mechanism. The high-threshold and low-threshold versions are contrasted in figure 4. In the high-threshold version, the first signal to arrive is met with a high threshold so that it has no observable effect on the output of the binocular mechanism but does provide an effectively reduced threshold so the onset of the binocular response will correspond, at least to a first approximation, to the arrival of the second signal plus or minus some neglible latency. However, since the variance of the last arrivals of the S1 and S2 evoked signals is identical to the variance of the first arrivals which was used in the low-threshold version of this model, both versions of the model predict the same level of binocular performance as given by equation 23 (see Appendix E). The high-threshold version does, however, predict somewhat different mean binocular arrival latencies, so the predicted mean of binocular temporal order judgment, μ_B, differs slightly, being given by

$$\mu_B = \mu_L + 0.5642(\sigma_{L,1} + \sigma_{L,2}) + \mu_o \qquad (24)$$

However, it is unlikely that this can be used to discriminate between the two versions since $\sigma_{L,1}$ and $\sigma_{L,2}$ are not observable in the temporal order judgment data.

Ueno's (1976) Model of Simple Reaction Time Applied to Binocular Temporal Order Judgment

Ueno (1976) has proposed a model to account for simple reaction time to double light pulses which could be extended to apply to the double signal arrivals in the binocular judgment of temporal order (see figure 4). He assumes

> (1) the visual system elicits at some levels a response caused by each of the two stimuli. ... The response represents the bivalent response which is temporally spread.
>
> (2) At one level of analysis, the system linearly superposes these response functions with appropriate weighting coefficients. ...
>
> (3) In order to detect a light pulse, the system continues to integrate from the onset of the pulse to a certain time, t_o, along the response function until the value becomes equal to a fixed criterion value, k. The time t_o may therefore be seen as a latency time that is consumed in the so called sensory decision stage. (p. 403)

This model also necessarily assumes that the signal is noise free so that the integral is equal to zero when the signal is absent.

Like the low threshold model, this model assumes that the two signals combine at some central locus and predicts a shorter arrival latency. It also assumes that either signal alone is sufficient to produce a detection. It differs from the low-threshold in many respects: The instant of detection occurs not where it would have been had only the faster signal been presented, but is displaced by the arrival of the second signal. The low-threshold model does not depend on any overlap of the signals in time.

Ueno's model requires the specification of the exact form of the internal signals to be integrated (such hypothetical functions are provided in, for example, Kelly, 1961) and of the nature of the "weighting coefficients" in the linear combination of the signals. For this reason, the model is not elaborated here for purposes of predicting temporal order judgments. This logic would be expected to predict heightened binocular sensitivity to temporal order, however.

Neural Counting and Neural Timing Models

Another potential summation model which will be outlined here is either what is termed a neural counting or a neural timing model (see McGill, 1967, for an example of a neural counting theory and Luce and Green, 1972, for a neural timing theory). Both of these models assume that the visual system's output is a set of sequences of brief events -- "neural pulses" -- which occur on "n" separate channels. Suppose that in its resting state, one of the n hypothetical sensory channels for the left and right eyes produces events which occur according to a Poisson process with rate parameters $\lambda_{L,0}$ and $\lambda_{R,0}$. Suppose also that these events simply combine at some level, i.e., the event train from the left system is simply superimposed upon the event train from the right. It can be shown that this binocular event sequence is also a Poisson process with resting rate parameter, $\lambda_{B,0,0}$, given by

$$\lambda_{B,0,0} = \lambda_{L,0} + \lambda_{R,0} \tag{25}$$

If a stimulus, say S1, is presented to both the left and right systems, their rate parameters are assumed to increase to $\lambda_{L,S1}$ and $\lambda_{R,S1}$, respectively. This leads to the rate parameter after binocular combination of

$$\lambda_{B,S1,S1} = \lambda_{L,S1} + \lambda_{R,S1} \tag{26}$$

If the stimulus S1 is presented to only one eye, say the left, then the combined rate would be only

$$\lambda_{B,S1,0} = \lambda_{L,S1} + \lambda_{R,0} \tag{27}$$

If the detection mechanism signals a detection with less variability when the combined rate is higher, then these models would predict an increased sensitivity to temporal order in binocular over monocular viewing.

Neural counting theorists would postulate that the observer detects weak signals on the basis of the number of events which occur on the hypothetical sensory channels within a fixed period of time. Thus the number of events acts as a code. Neural timing theorists, on the other hand, suppose that the response is decided by measuring the amount of time which is required for a fixed number of events to occur on the hypothetical channels. The code here is time.

Wandell (1977, p. 225) has shown that a neural timing model does predict a reduced reaction time variance for increased rate parameters -- namely, that

$$\nu_2^2 = \left(\frac{\lambda_1}{\lambda_2}\right)^2 \nu_1^2 + \sigma_m^2\left(1 - \frac{\lambda_1}{\lambda_2}\right)^2 \qquad (28)$$

where ν_2^2 is the variability in reaction time to the larger rate parameter, λ_2, and ν_1^2 is the variability in reaction time to the samller rate, λ_1, and σ_m^2 is the variability of the remaining (nonsensory) component of reaction time which does not depend on timing theory. Since the ratio λ_1/λ_2 is less than 1.0, equation 28 may predict decreased variability depending on the relative magnitude of σ_m^2. If it is zero, and we apply the rate parameter relationship given by equations 26 and 27, then it is predicted that

$$\sigma_B^2 = \left[\frac{\lambda_{R,0} + \lambda_{L,S1}}{\lambda_{R,S1} + \lambda_{L,S1}}\right]^2 \sigma_L^2 \qquad (29)$$

which predicts a nonlinear relationship as a function of monocular variance. The rate parameters decrease with stimulus intensity. When the intensity is very low, $\lambda_{L,S1}$, is nearly equal to $\lambda_{R,0}$ and σ_B should be approximately equal to σ_L. On the other hand, as stimulus intensity increases, $\lambda_{R,0}$, becomes less important and the contents in the brackests approaches one-quarter, the greatest increase in binocular sensitivity of any model thus far considered.

Either neural counting or neural timing models would result in not only shorter arrival latencies, but probably also reduced arrival latency variance with binocular viewing. These models, however, require specification of many free parameters -- the number of channels, the resting rate, the relationship between stimulus intensity and rate, the criterion number of events for detection, etc. -- and the derivations become very complex even for the simplest single-channel cases.

Binocular Superiority Through Noise Reduction

There are other models which could predict a binocular advantage in temporal order judgment without postulating a direct combination of the two monocular signals. The phenomenon of masking level differences in vision (Henning and Hertz, 1973, 1977) has inspired such models. Binocular grating detection in spatio-temporal noise is improved when signal or noise is presented dichoptically with differing phase. Explanations for such masking level differences often involve the notion of the output of one channel being "subtracted" or "cross-correlated" with the output of the other (see Green and Yost, 1975) so that the magnitude of the uncorrelated or out-of-phase component is altered relative to the other. Indeed, Cohn and Lasley (1976) support a theory of binocular summation at threshold which supposes that there are two combination mechanisms, "one that sums signals arising from two eyes and one that computes a difference, provide information to a more central decision center" (p. 561), to account for summation effects using both incremental and decremental flashes. In temporal order judgment experiments, the reduction of noise in binocular viewing could allow for a lowered detection criterion and, as a result, a faster and less variable arrival latency.

PILOT EXPERIMENT

The ability of an observer to discriminate which of two spatially separated but otherwise identical flashes came first shows increased variability as flash intensity decreases (Zacks, 1973). This implies that latency variability can be manipulated indirectly by manipulating the stimulus intensity. In the experiment at hand, discriminability of temporal order will be measured as a function of (a) the delay between the onset of one stimulus, S1, and the onset of another stimulus, S2, of the same intensity and duration, (b) viewing condition, either right-eye, left-eye, or binocular, and (c) stimulus intensity. The observer's task is to identify which stimulus came first.

Apparatus

The two light sources were 12' of visual angle in diameter and separated by 1^{o}. This separation is well beyond the minimum separation (12') for which the dichoptic and binocular temporal order judgment performances are equal and, hence, for which retinal interaction is assumed to be neglible (Westheimer and McKee, 1977). The most likely cause of non-independent responses is thus eliminated. These light sources were arranged symmetrically about a central fixation point provided by a small incandescent bulb which was completely blackened except for a small spot optically located between the stimuli with a beamsplitter. The light sources were LEDs (Monsanto MV5322, typical wavelength of 589 nm) whose luminances were calibrated with a photopic photodiode (UDT 500) and recalibrated with neutral density filters in place. The LEDs were driven by a 5 ms rectangular current pulse. The right light source was designated as S1. The observer's head was held in a chin rest. A LINC digital computer controlled all procedures including stimulus presentation, timing and randomization, data collection and storage, and feedback to the observer.

Procedure

A two-choice, forced-choice procedure was used with trial-to-trial feedback. A single trial proceeded as follows: The dark-adapted observer, seated in a light-tight test compartment, wore a set of headphones through which he heard a moderate level of white noise. One second before the presentation of the stimuli the noise was turned off as a warning stimulus. The noise remained off while the stimuli were presented and until either (1) the observer made a response, or (2) two seconds elapsed without a response. Either event terminated that trial. The observer's task was to choose the flash which came first. He signalled his judgment by pushing a bi-directional lever switch with his right hand to the same side as the flash he saw first. This response produced, in addition to the resumption of white noise, a 500 ms tone when the observer made a correct response. The total trial length varied from 3.0 to 5.0 seconds.

Two different procedures were used to generate stimulus onset asynchronies from trial-to-trial in this experiment. Some data sets (GEN3, MWH1, and JLZ) used a method of constant stimuli in which an asynchrony from a set of twenty possible asynchronies ranging from -50 ms (S2 preceding S1) to +50 ms (S1 preceding S2) in 5 ms intervals (excluding physical simultaneity which would upset the correct feedback rule) was selected randomly from trial-to-trial. In the remaining data sets, the asynchronies were presented according to a modified double random staircase procedure (Cornsweet, 1962). Here the stimulus onset asynchrony to be presented on a given trial depended on the observer's previous responses on the staircase; if he was incorrect before, the next asynchrony would be larger in absolute magnitude, and if he was correct on the preceding two trials, the next asynchrony would be smaller in absolute magnitude. The sign of the asynchrony (i.e., S1 preceding S2, or S2 preceding S1) was varied randomly from trial-to-trial. Since two staircases are run simultaneously, also being selected randomly from trial-to-trial, the observer is not aware of the true staircase nature of the procedure. The advantage of the staircase, of course, is its efficiency in that it concentrates the sampling near the 75 per cent correct level regardless of the observer's sensitivity. Furthermore, it may reduce potential undesirable range effects (Sternberg and Knoll, 1973). The range and sampling density in the method of constant stimuli, on the other hand, is bound by the predetermined values. Three increment sizes were used in the staircase runs: 10 ms (GEN4), 5 ms (GEN1 and GEN2), and 2 ms (GEN5).

Four stimulus intensities for the 5 ms duration flashes were used: 0.4, 0.7, 1.0 and 1.1 log units above the monocular response thresholds. The observer's threshold intensity was determined using a double random staircase method (Cornsweet, 1962). Of the four thresholds resulting from the right and left eye conditions by the stimuli 1 and 2 conditions for each observer, the largest difference between the highest and lowest threshold estimates was about 0.1 log unit, or approximately one jnd. This difference was probably not significant and did not seem to depend particularly on the stimulus conditions. For this reason, the threshold luminance was equated for all four conditions at approximately the geometric mean luminance value.

The lowest luminance value, 0.4 log unit above the monocular response threshold, was chosen because this seemed to be the lowest value for which detection was approximately certain. Less than perfect detection of the stimuli would disrupt the logic underlying the temporal order judgment model. Higher intensities were not used because at these levels the absolute differences between the monocular performances and the binocular performances are expected to be small by any model given the high level of monocular performance observed by Zacks (1973) at higher luminance levels.

Inexperienced observers were not tested until they had completed a minimum of 500 responses. This training on the temporal order judgment is essential since the practice effect can be very large. The greatest learning, however, occurs during the first 320 trials (Gengel and Hirsh, 1970; Thor, 1968).

The datum obtained from the temporal order judgments is the proportion of responses made to the flash on the right of the fixation point (S1) for a fixed stimulus onset asynchrony. This datum is preferable to the per cent correct judgment which is commonly used to construct psychometric functions plotted as a function of the absolute magnitude of the asynchrony (Sternberg and Knoll, 1973). This function is meaningful only if the "S1 first" psychometric functions were mirror-symmetric about the origin (i.e., if $F(d) = 1 - F(-d)$) and if the point-of-subjective-simultaneity is zero. If these conditions are not met, then significant distortion of the function can result from this "folding" of percent correct about the origin (see Zacks, 1973). This distortion has even led Oatley, Robertson and Scanlan (1969) to postulate that the temporal order psychometric function is sigmoidal, but bounded by 0.5 at the origin. At any rate, the psychometric function based on the proportion of S1 first judgments, F(d), was summarized by two parameters: (1) the point-of-subjective-simultaneity or mean, μ, and (2) the standard deviation, σ. These parameters were attained through Probit analysis (Finney, 1952) which, in essence, finds the best-fitting ogive function through the data points. Other forms of the psychometric function have been assumed (Kristofferson, 1967; Oatley, Robertson and Scanlan, 1969) but the normal ogive tends to provide a reasonable approximation to the data.

Most of the models discussed in this paper demanded that the performance of the right and left monocular systems be nearly identical in order to simplify the derivation of their predictions. To examine how similar the monocular variances were, an F-ratio was formed by dividing the larger monocular variance by the smaller. This ratio, unfortunately, could not be tested for significance since it is not clear how its degrees of freedom should be assigned. The ratio, nevertheless, does give some qualitative measure of how well the assumption of identical monocular variances was met. Assuming that the two monocular variances from a given data set differ only by chance, a single monocular variance, σ_M^2, was formed from the average of the two monocular variances.

Observers

Three observers participated in the pilot experiment: the author, GEN, JLZ and another paid observer, MWH. Five data sets at three intensity levels were collected on observer GEN. Observer JLZ and MWH produced only one data set each. All three observers were male, brown-eyed and right-handed. GEN and JLZ both required corrective lenses and were right-eye dominant. MWH had normal acuity uncorrected and was left-eye dominant. None of the observers were stereo-blind, since stereoblind observers have been shown to be no better with two eyes than one on binocular summation measures like contrast threshold (Lema and Blake, 1977).

The Psychometric Function Results

The psychometric function was defined as the proportion of times that the observer said that S1 preceded S2 as a function of the stimulus onset asynchrony, d. Such functions that resulted in the pilot experiment are listed in Appendix F. A cumulative normal distribution function was fitted to each function using probit analysis (Finney, 1952). This assumption that the density function underlying each F(d) was normal was examined by considering the correlation between the "linearized" proportion, p_{lin}, and the stimulus onset asynchrony, d. If the density underlying F(d) is normal, then F(d) can be linearized by $p_{lin} = F_o^{-1} F(d)$ where F_o^{-1} is the inverse standard normal cumulative distribution function. Since p_{lin} is undefined for F(d) = 0.0 and F(d) = 1.0, these certain probabilities were necessarily omitted from the correlation but were included in the probit analysis. The twenty-one correlations in the pilot experiment ranged from 0.781 to 0.995 with a median value of 0.974. Seventy-five per cent of the correlations exceeded 0.936 and twenty-five percent exceeded 0.989. Figures 5, 6 and 7 depict the obtained F(d) along with the best fitting ogives for the functions with the highest (GEN4, left eye), the median (GEN1, right eye) and the lowest (GEN5, both eyes) correlation, respectively.

This correlation does, under certain circumstances, provide a measure of how well the assumption of normality was met for each function. This is true if the extreme proportions, 0.0 and 1.0, seldom occur; if the number of points making up each function is the same; and if the variability in measurement of each point is roughly the same. The conditions were not met in this experiment (see Appendix F) so these correlations can only partially index the quality of the normality assumption.

Probit analysis leads to the two parameters, the mean or point-of-subjective-simultaneity and the standard deviation. Both will be used to examine these functions. Since probit analysis weights each point by the number of observations that went into it, it circumvents some of the problems with the use of the correlation to index normality, namely, the variability in measurement due to varying

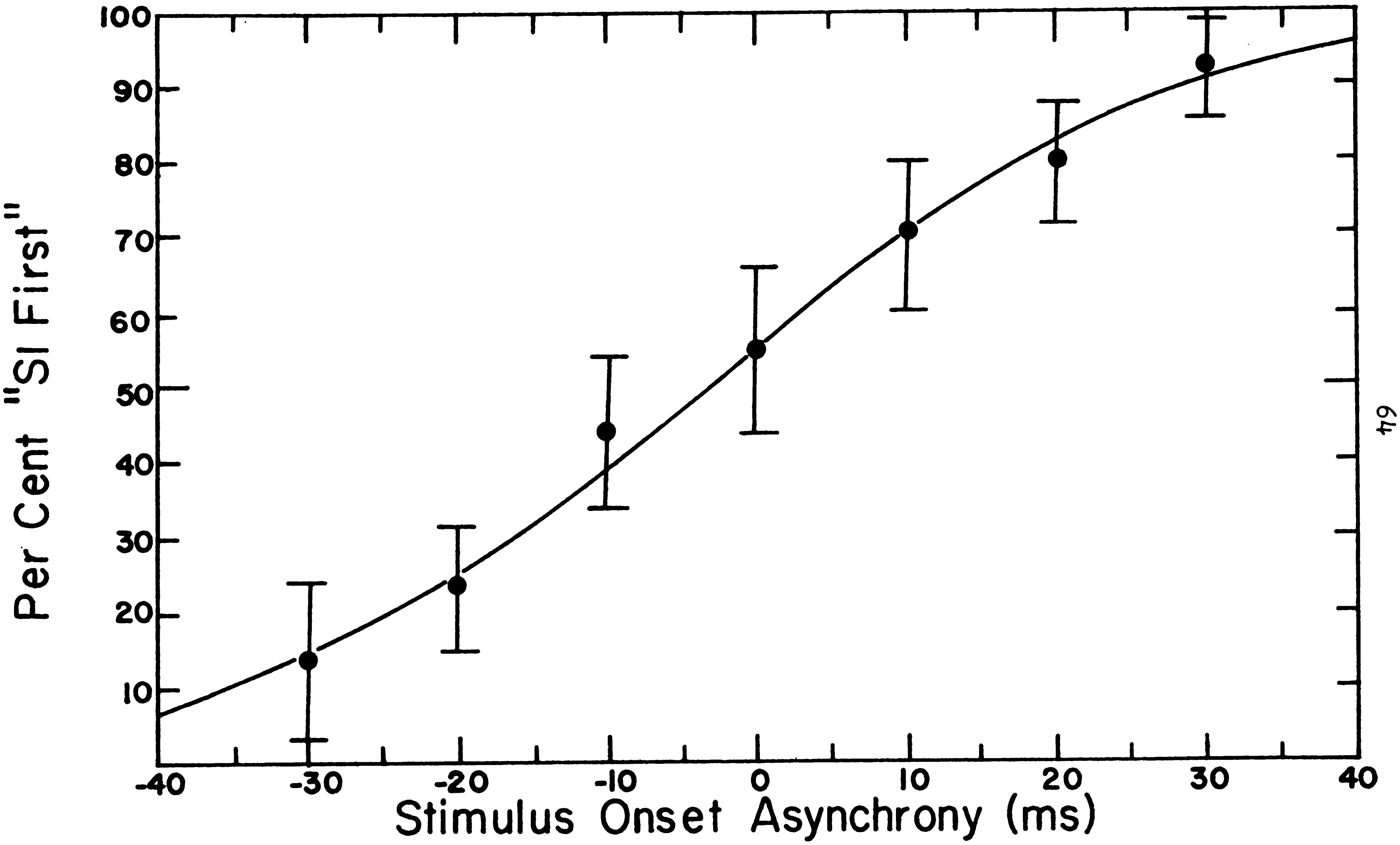

Figure 5. The psychometric function in the pilot experiment with the highest correlation between p_{lin} and d

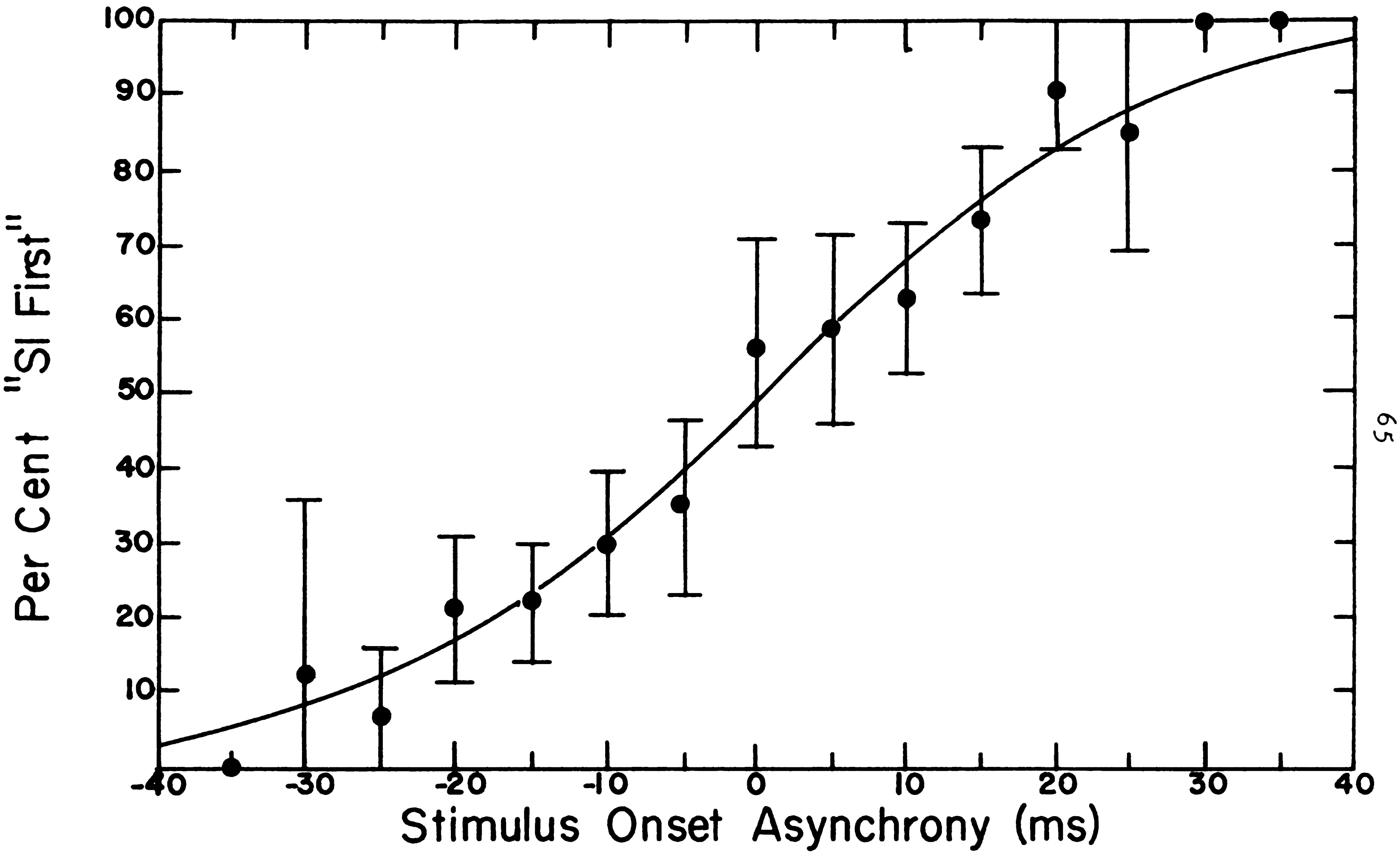

Figure 6. The psychometric function in the pilot experiment with the median correlation between p_{lin} and d

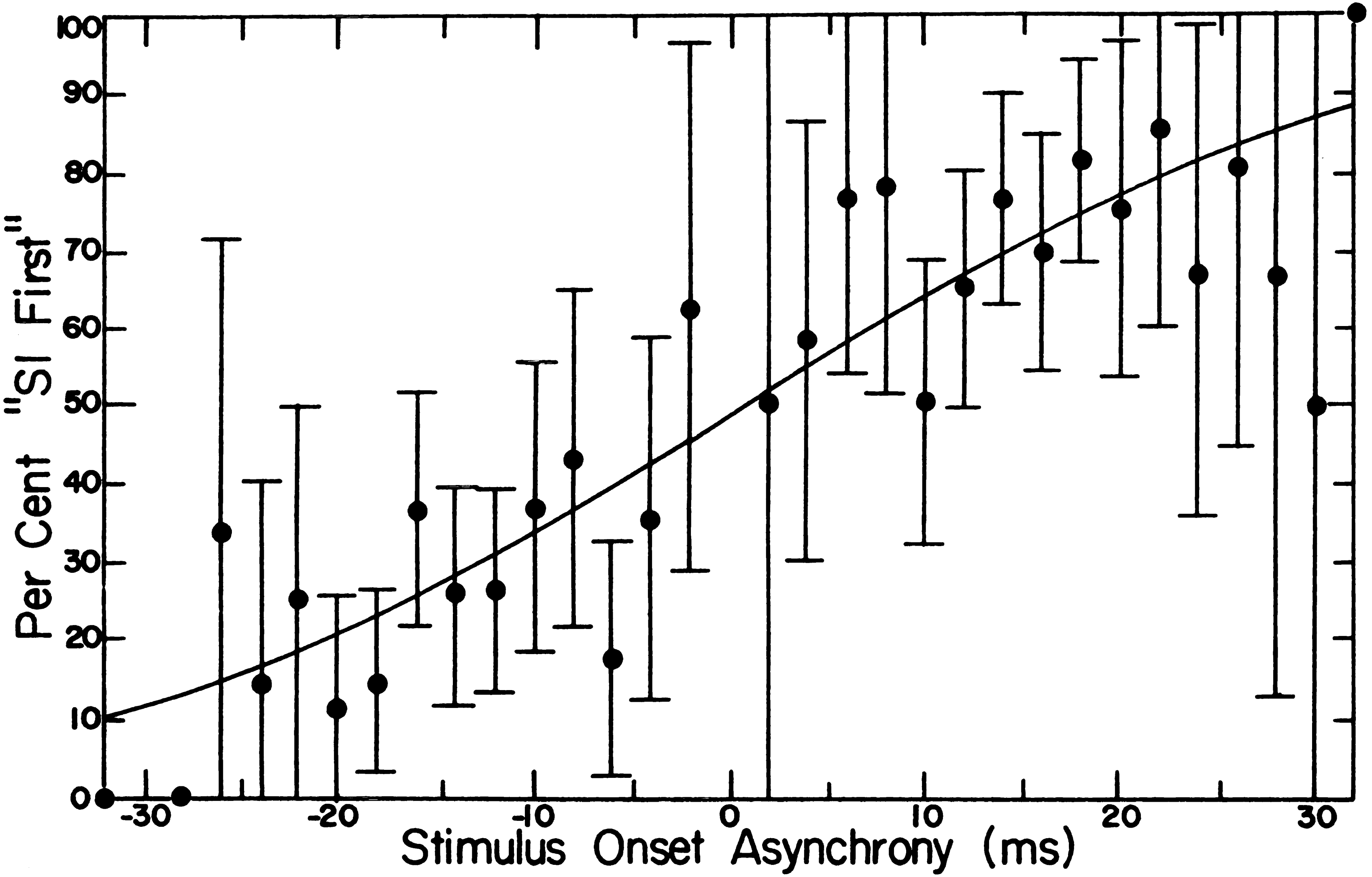

Figure 7. The psychometric function in the pilot experiment with the lowest correlation between p_{lin} and d

number of observations per point. Because of this, it probably provides a better fit than would be implied by the correlations alone.

The Standard Deviation Results

The results of applying probit analysis to both the monocular and binocular psychometric functions are listed in Table 1 (as well as in Appendix F). Also listed in Table 1 are the monocular standard deviations, σ_M (the square root of the average monocular variance); the critical ratio σ_B/σ_M; and the ratio of the larger monocular variance to smaller monocular variance which was a measure of how well the assumption of equal performance in left-eye and right-eye viewing conditions was met. Figure 8 graphically displays the observed relationship of the binocular standard deviation, σ_B, as a function of the monocular standard deviation, σ_M. Also indicated in figure 8 are the values of the binocular standard deviation predicted by the various models.

In only one instance (GEN4) did the binocular standard deviation exceed either the left-eye or right-eye standard deviations and in no case was there a failure to observe any binocular superiority as measured by the σ_B/σ_M ratio. Nor did this ratio fall below the "baseline for independence" in figure 8. This baseline was established by the likelihood ratio model, by the average arrival time model, and by the most discrepant arrival time model. The observed

Table 1

Parameters describing temporal order judgment psychometric functions in the pilot experiment

O	log above thres.	LEFT			RIGHT			BINOCULAR					
		μ_L	σ_L	n	μ_R	σ_R	n	μ_B	σ_B	n	σ_M	$\frac{\sigma_B}{\sigma_M}$	F
GEN1	1.1	-0.23	18.72	888	0.49	20.75	764	2.83	16.51	764	19.76	0.836	1.229
GEN2	1.1	-0.78	22.06	936	-3.44	21.68	956	1.67	18.19	954	21.87	0.832	1.036
GEN3	0.7	-0.96	23.33	1525	-3.73	21.77	1531	2.33	19.31	1529	22.57	0.856	1.148
GEN4	0.4	-3.38	24.35	573	-3.64	29.17	383	0.29	25.90	574	26.87	0.964	1.436
GEN5	0.4	-1.10	36.15	695	0.93	27.59	765	1.47	26.06	576	32.16	0.810	1.716
MWH1	0.7	-4.32	29.46	1068	-6.48	30.22	1066	-5.85	25.23	1077	29.84	0.845	1.053
JLZ	1.0	-9.35	35.44	606	4.41	45.27	605	-0.99	32.43	605	40.65	0.798	1.632

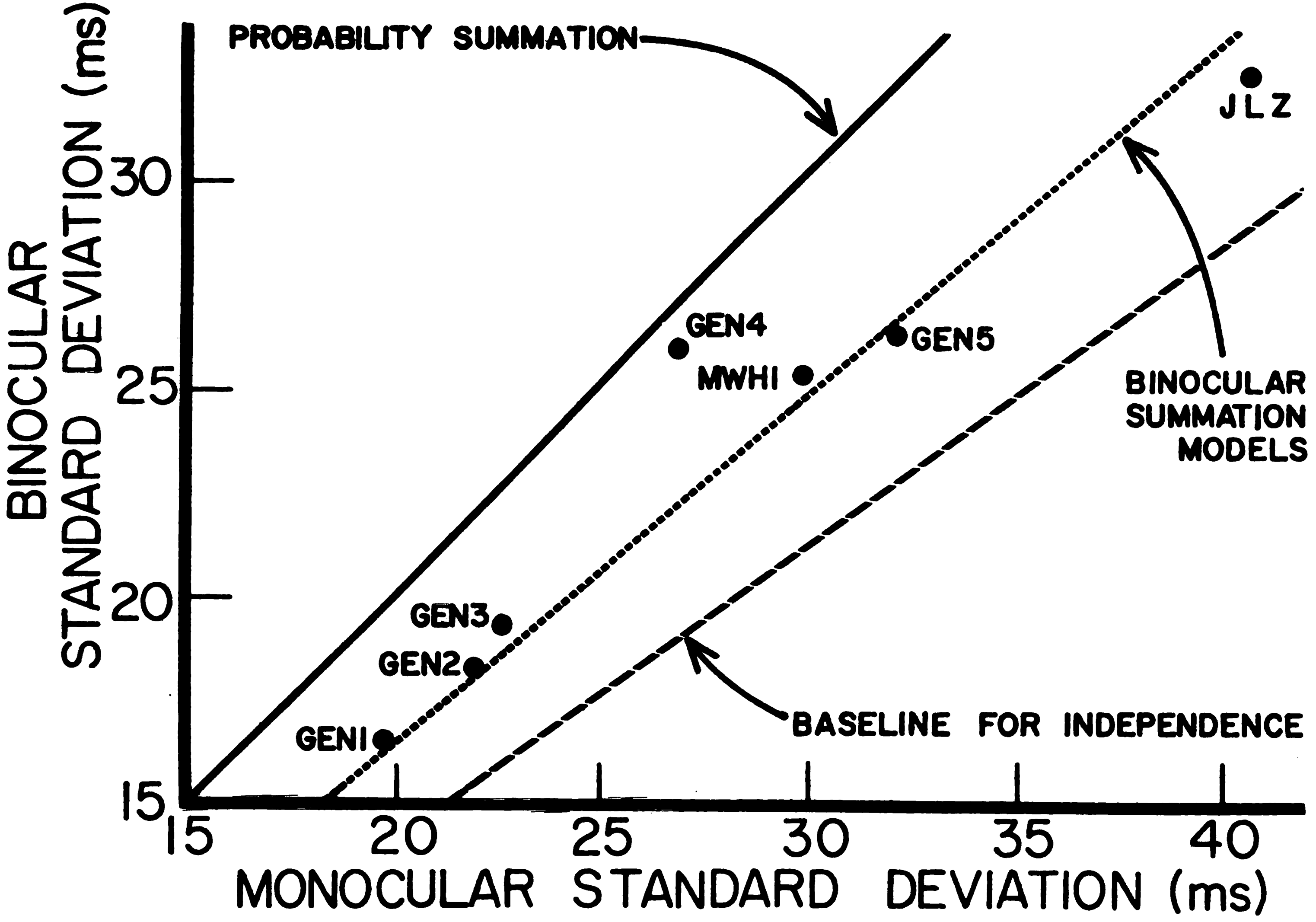

Figure 8. The binocular standard deviations plotted as a function of the monocular standard dviations in the pilot experiment

relationship between the binocular and monocular standard deviations was highly linear ($r = 0.967$, $p < 0.001$); the best-fitting line had a slope of 0.749 with a σ_B intercept of 2.660 ms. The models presented here predicted linear functions with slopes between 0.707 and 0.826 which intersect the origin. The observed ratios of binocular standard deviations to monocular standard deviations had a mean of 0.849 and standard deviation of 0.051. The average standard deviations were 28.064 ms, 27.073 ms and 23.376 ms for right-eye, left-eye and binocular psychometric functions. These average standard deviations are in a ratio of 1.00: 0.97:0.83. A multiple linear regression on these standard deviations reveals a best-fitting line of

$$\hat{\sigma}_B = 2.548 + 0.301\,\sigma_L + 0.453\,\sigma_R \tag{30}$$

and a similar multiple linear regression on the variances gives

$$\hat{\sigma}_B = \sqrt{99.862 + 0.246\,\sigma_L^2 + 0.334\,\sigma_R^2} \tag{31}$$

All of the models described in this paper predicted that the binocular variance would be a linear function of the two monocular variances. The regression on the standard deviations, however, produces a slightly smaller mean squared error (1.634 compared to 1.898). The data presented here do indicate that is some increase in sensitivity to

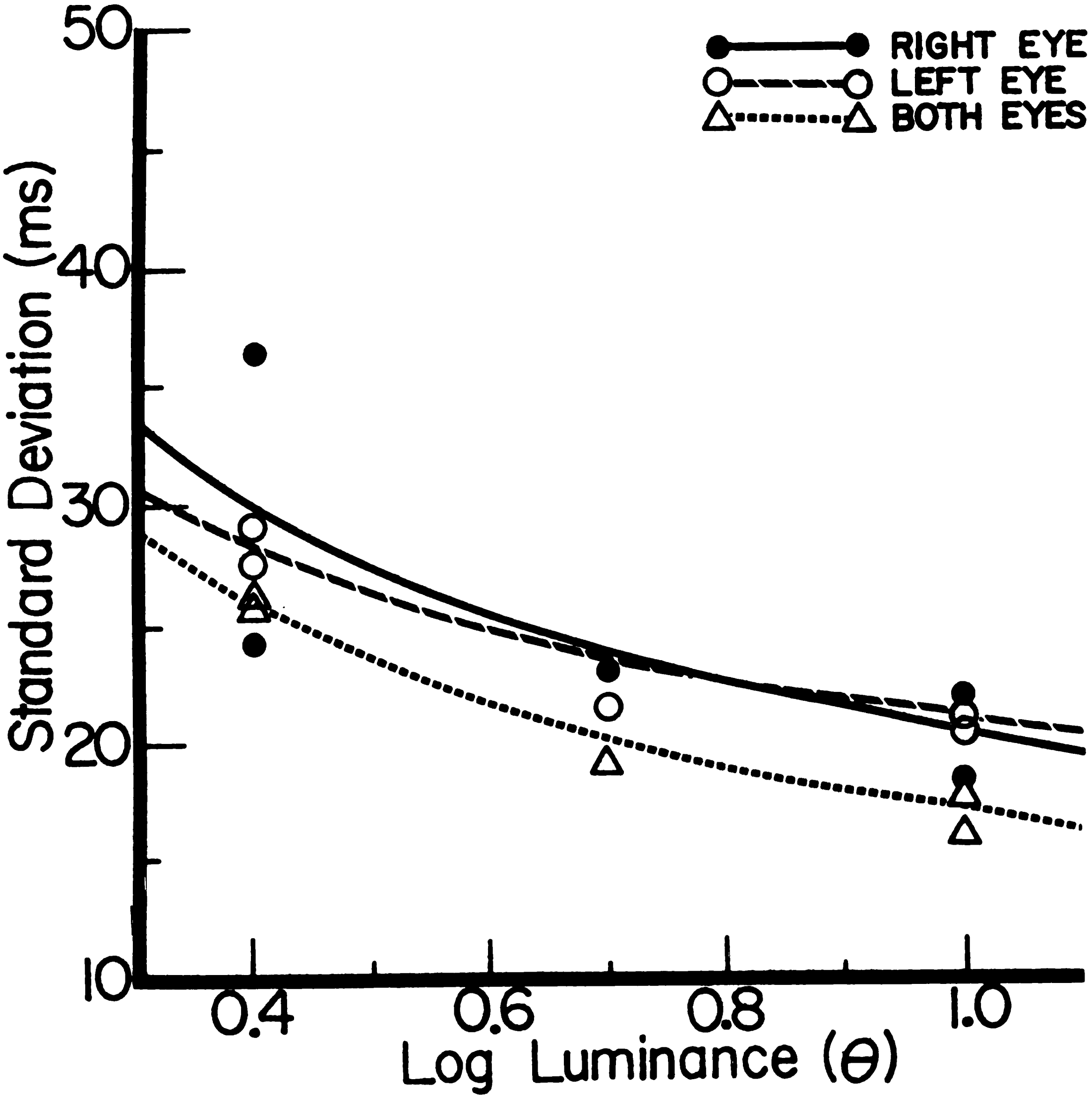

Figure 9. Standard deviations plotted as a function of log luminance for the pilot experiment with the best-fitting power functions

temporal order with binocular viewing over monocular and that this superiority is most consistent with the binocular summation models.

The relationship between stimulus intensity and the variance observed in the temporal order judgments might be mentioned. This is the same relationship noted by Zacks (1973), namely, that the variance decreases as stimulus intensity increases. The standard deviation data were approximated by power functions with coefficients of 21.314 ms and 20.918 ms for right-eye and left-eye viewing and of 17.22 ms for observer GEN's binocular viewing. Exponents for GEN's data were similar: -0.375, -0.292 and -0.403 for right-eye, left-eye and binocular viewing, respectively. These power functions are shown in figure 9. If the exponents were identical and if the power function is an appropriate function for describing the data, then the ratio of the coefficients expresses the proportional differences between the functions independent of luminance. The ratio of the binocular power function coefficients to the average monocular coefficients was 0.839.

The Point-of-Subjective-Simultaneity Results

There was little systematic effect on the point-of-subjective-simultaneity, or mean, with manipulations of stimulus intensity. There was some indication that the binocular mean may be displaced more toward positive values than were monocular means. Observer GEN's average means for the right-eye, left-eye and binocular viewing conditions were -1.88 ms, -1.29 ms and +1.72 ms, respectively. This tendency for the monocular point-of-subjective-simultaneity to be negative (i.e., S2 had to precede S1 to produce subjective simultaneity) might be accounted for either by (1) a stimulus bias, that is, the stimuli may not have been precisely calibrated or equated for luminance so that the right stimulus (S1) was slightly more intense than the left (S2) which would tend to move the mean toward negative values (see Sternberg and Knoll, 1973); (2) a response bias such as the so-called "prior entry effect" (Sternberg and Knoll, 1973; Vanderhaeghen and Bertelson, 1974); or (3) some inherent temporal processing mechanism which yields shorter arrival latencies for the right-hand as opposed to the left-hand stimulus. Such a hypothetical processing order is inconsistent with left-to-right processing of visual temporal order reported by Sekuler, Tynan and Levinson (1973) who reported that "brief visual stimuli presented in rapid sequence, one to the left and one to the right, appear to occur left first, then right, regardless of the actual order of presentation" (p. 210). Such a mechanism should tend to produce positive means in temporal order judgment. However, the variability of the means in table 1 makes any conclusion in these areas premature. The binocular mean, μ_B, can be predicted from the monocular means, μ_L and μ_R, from the following multiple regression equation for observer GEN

$$\hat{\mu}_B = 2.638 + 0.753\,\mu_L - 0.027\,\mu_R \tag{32}$$

which predicts μ_B with a mean square error of 0.115. Since observers MWH and JLZ's means differed markedly from those of GEN, they have been ommitted from this analysis.

Discussion

The data shown in figure 8 do suggest that there is some binocular processing of temporal order which leads to a small but reliable increase in sensitivity to temporal order in the binocular viewing condition. This is in contrast to the results of a binocular temporal order judgment experiment by Robinson (1967) who claimed that binocular viewing resulted in performance which was "essentially identical" to that of monocular viewing, but that sensitivity increased with dichoptic viewing of the two stimuli. Robinson, however, did not make direct comparisons of binocular monocular performances by the same observers with the same apparatus. Rather he accepted Hirsh and Sherrick's (1961) dictum that the 75 per cent correct level at 20 ms represented "normal" temporal order performance. The stimuli used by Robinson were, in addition, quite different from those described here so that it is quite possible that his belief that peripheral interactions decreased sensitivity when each eye viewed both stimuli was correct (Robinson, 1967, 1968). Other studies (Thor, 1967, 1968), however, have failed to replicate Robinson's finding. Yet other studies have suggested the opposite result. Westheimer and McKee (1977) compared binocular to dichoptic performances using line stimuli in close proximity. They found that the dichoptic viewing situation <u>increased</u> the threshold for temporal order for stimulus separations less than about 12'. By using various stimulus configurations, they further demonstrated that this decreased sensitivity in the dichoptic trials could not be due solely to random eye vergence errors, but demanded a facilitory peripheral interaction. They conclude that "the ability to detect the temporal order of adjacent stimuli can cross eyes, but the threshold never quite reaches the best values for the situation where both stimuli are seen by the same eye" (p. 890). Using large stimuli whose chromatic components were made asynchronous, Yund and Efron (1974) demonstrated that threshold onset asynchrony of the color components was higher on dichoptic than on monocular stimulus presentation (6 ms dichoptic versus 2 ms monocular).

The results in the present experiment do seem to indicate that there is an increase in sensitivity to temporal order with binocular viewing. Further, this level of performance evident in figure 8 is closest in magnitude to that predicted by the two binocular summation models. Although the effect was shown to be a reliable one, it is, nevertheless, so small that it is not surprising that other authors (e.g., Robinson, 1967; Westheimer and McKee, 1977) who have performed binocular temporal order judgment studies have failed to notice it.

Several procedural objections may be raised against the pilot experiment, however. First, no single data collection strategy was used to collect all data and, second, since all right-eye, left-eye and binocular temporal order judgments were run in groups of homogeneous types, the observer was well aware of the viewing conditions at the time the stimuli were presented. It could be argued that the binocular superiority observed in the pilot experiment was due to motivational or strategic rather than sensitivity changes which occurred from one stimulus condition to the next.

EXPERIMENT 1

Experiment 1 was planned simply to perform more comparisons of binocular to monocular performance in order to confirm the increased sensitivity to temporal order for the binocular viewing condition observed in the pilot experiment under more rigorously controlled experimental conditions. To remedy the complaints made against the design of the pilot experiment the following steps were taken in experiment 1. First, all judgments were collected under identical schedules as described below and second, the apparatus was modified to include a barrier which could be moved silently in a magnetic field, making no contacts except with the bearings which held its axle, to occlude either the right eye, left eye or neither from trial-to-trial. In this way, the observer had no systematic expectations as to the nature of the next trial and so would be less likely to employ different strategies for different trial types and would be similarly motivated for all trial types. The spot separation was also increased from 1° to 3° 23' of visual angle to further ensure the independent responses required by the independent channel model of temporal order judgment. This was approximately the separation used by Zacks (1973).

Apparatus

The optical system used in this experiment was similar to that used in the pilot experiment. Two light sources, 12' in diameter and separated by 7.5 cm, were arranged symmetrically to the left and right of a central fixation point for both eyes. This led to an optical path length of 1.270 m for the light source on the same side as a given eye and of about 1.272 m for the light source on the other side. The lights then had a separation of 3^{o} 23' for each eye. This larger separation was chosen to further minimize interactions between the flashes at the retinal level (and, hence, to eliminate the most likely source of nonindependence) and to make the results directly comparable to Zacks (1973).

The light sources were LEDs (Monsanto MV5352, typical wavelength of 585 nm) which were driven by a rectangular current pulse of 5 ms duration. The LEDs were placed into holes drilled into a 3/8 inch thich brass plate. The entire plate was painted flat black to minimize reflectance. The directional properties of the LEDs were minimized by placing several layers of frosted Mylar sandwiched between microscope slides against the front of the plate. A thin metal sheet with a hole drilled at its center was placed over the final microscope slide in order to provide a sharp image.

As in the pilot experiment a small incandescent bulb, blacked out except for a small spot, served as the fixation point for both eyes. It was located precisely between the two light sources with a beamsplitter. This beamsplitter was interposed between the observer and the occluding barrier which controlled the viewing condition so that fixation was binocular for all trials. This removed the difference in fixation conditions as source of error in the pilot experiment. Otherwise, error could arise due to differences in pupillary and accommodative responses (see Blake and Fox, 1973, and Haines, 1977).

Unlike the apparatus used in the pilot experiment, a barrier was added which could be silently moved in a magnetic field to block stimulation to the left, right, or neither eye. The luminances of the LEDs were calibrated in the same manner as described in the pilot experiment. The observer's head was again held in position with a chin rest and all procedures including stimulus presentation, timing and randomization, choice of trial condition (left, right or both eyes), data collection and storage, and feedback to the observer were controlled by a LINC digital computer.

Procedure

The timing and sequence of events that occurred in a given trial were described in the pilot experiment. However, the stimulus onset asynchronies were presented using a method of constant stimuli for all trials. Trials were arranged into blocks of 192 trials. Within a block, trials were randomized according to two classifications:

(1) stimulus onset asynchrony, and (2) left, right or binocular viewing condition. Twenty asynchronies were used ranging from -50 ms (S2 preceded S1) to +50 ms (S1 preceded S2) in 5 ms intervals, omitting physical simultaneity which would upset the correct feedback rule. These were selected randomly from trial-to-trial with the frequency distribution shown in figure 10. Six blocks were presented per day for three days at each of three intensity levels for each observer. The first block of each day was omitted from analysis to assure that a state of dark adaptation was obtained. This yielded an expected 960 observations for each ocular condition at each intensity. The observers were allowed to rest between blocks, remaining in the darkened room. Binocular dark adaptation was acheived each day prior to testing by wearing red-lensed, "night" goggles for 20 minutes and then sitting in the darkened test chamber for an additional period of approximately ten minutes before beginning the first of the blocks of trials.

Three stimulus intensities for the 5 ms stimuli were used: 0.6, 1.0 and 1.4 log units above the monocular absolute threshold. The observer's threshold was determined using a double random staircase method (Cornsweet, 1962). For the same reason as given in the procedure section of the pilot experiment, the threshold luminance was equated for the four relevant conditions: right and left eyes by right and left stimuli.

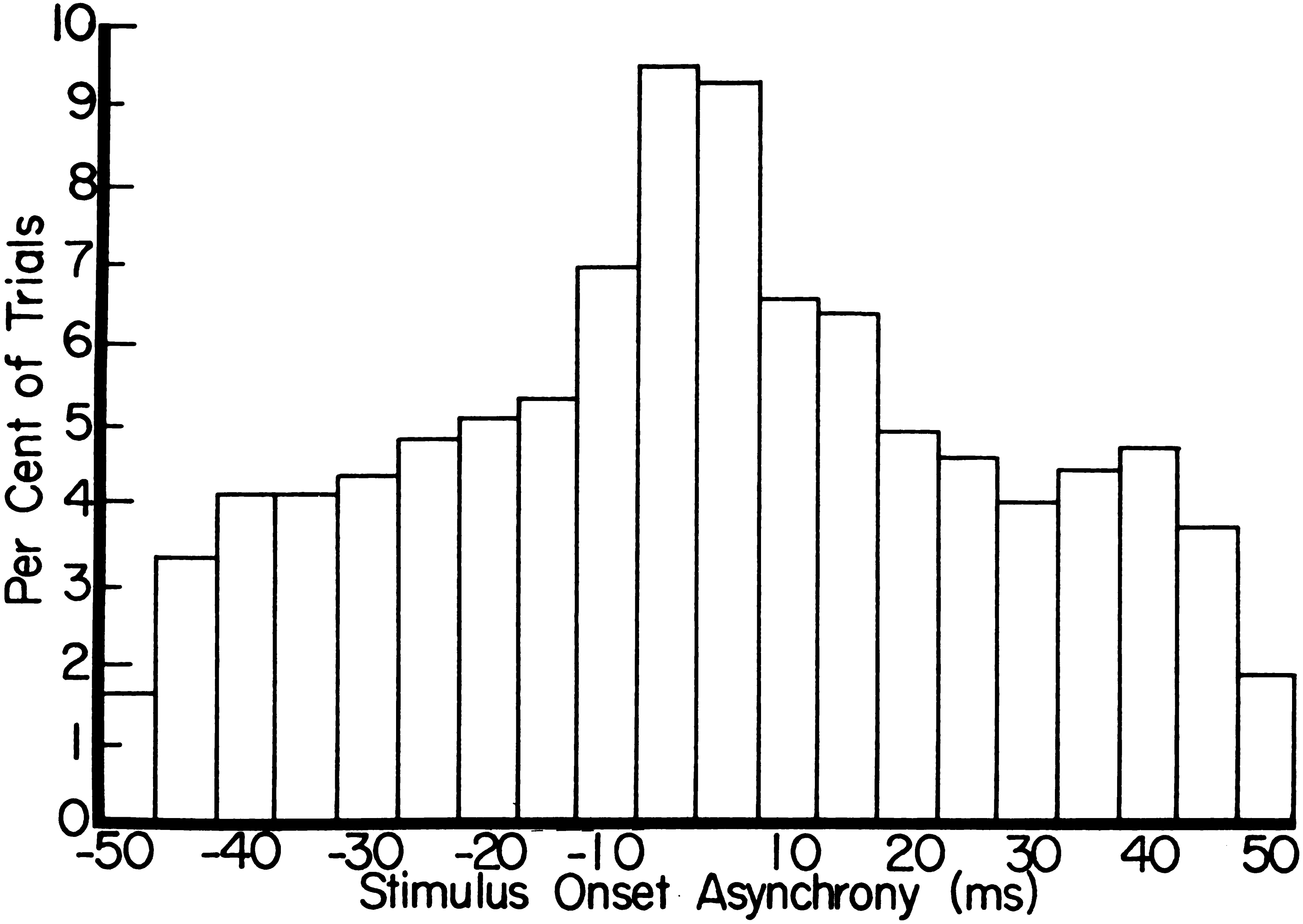

Figure 10. The sampling of stimulus onset asynchronies in experiment 1

The psychometric function and parameters for the best-fitting ogive function were obtained in the manner discussed in the pilot experiment.

Observers

The author, GEN, and one additional paid observer, MWH, served in all conditions of this experiment. Both observers were experienced in the temporal order judgments since they also took part in the pilot experiment.

The Psychometric Function Results

The psychometric functions F(d) which resulted from experiment 1 are listed in Appendix G. Probit analysis (Finney, 1952) was again used to fit cumulative normal distribution functions to these data. The assumption that the density function underlying each F(d) was normal was examined, as it was in the pilot experiment, by considering the correlation between the "linearized" proportion and the stimulus onset asynchrony (see Appendix G). The eighteen correlations ranged from 0.940 to 0.991 with a median value of 0.969. Seventy-five per cent of the correlations exceeded 0.953 and twenty-five per cent exceeded 0.986. Figures 11, 12 and 13 depict the obtained F(d) along with the best-fitting curves for the temporal order judgment data with the highest (GEN6, right eye), the approximate median (MWH3, left eye), and the lowest (GEN8, both eyes) correlation. In experiment 1, unlike the pilot experiment, these

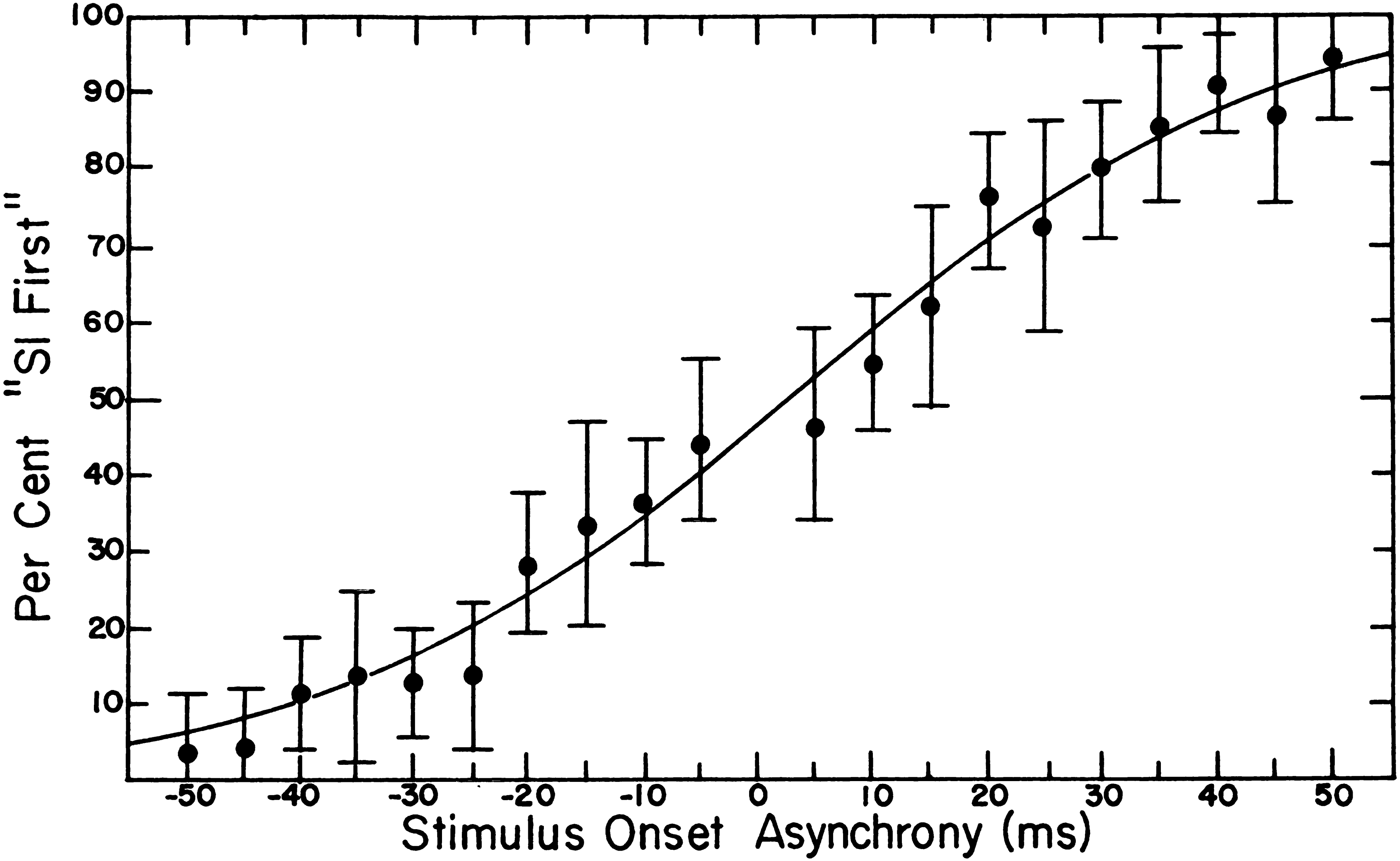

Figure 11. The psychometric function in experiment 1 with the highest correlation between p_{lin} and d

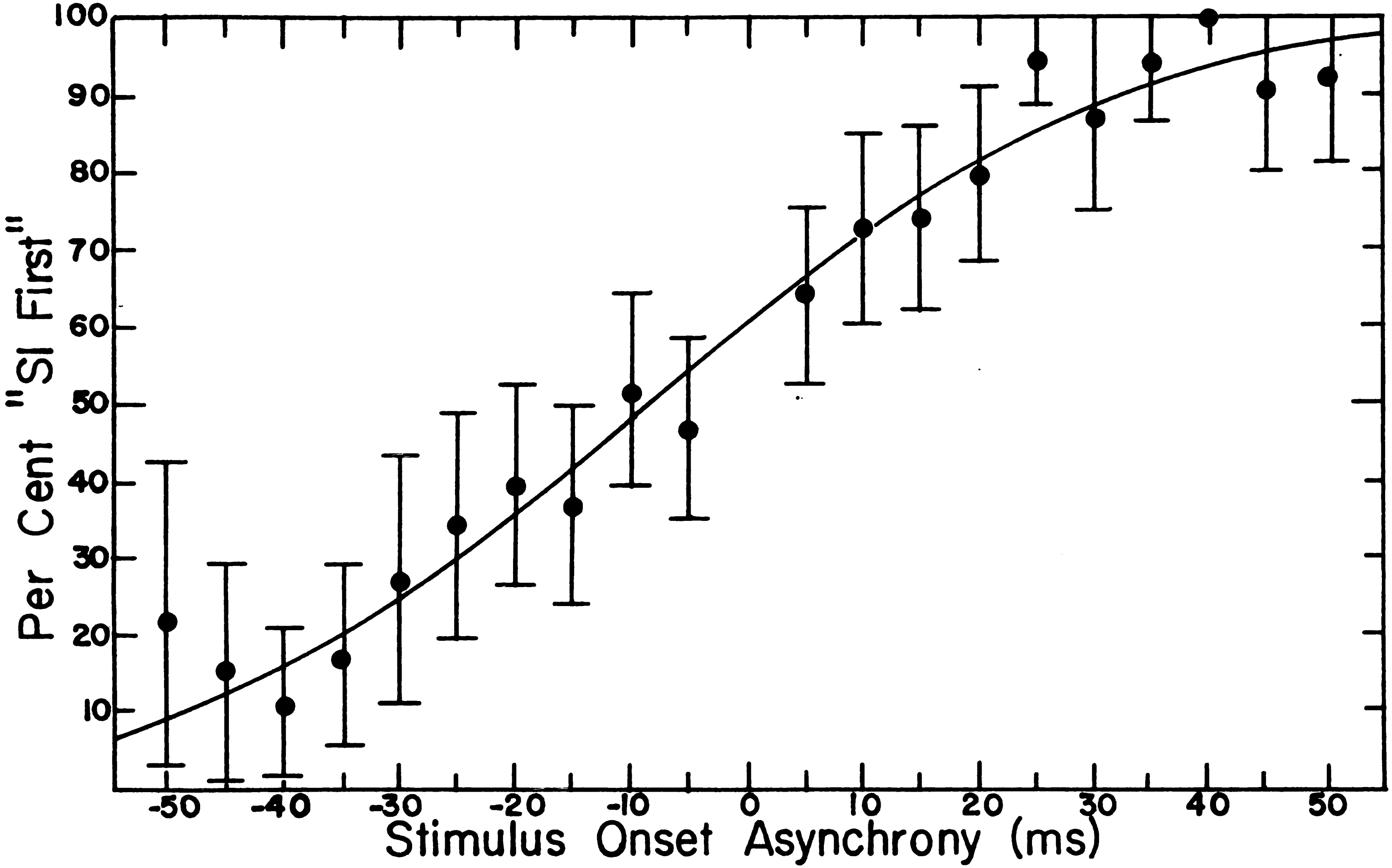

Figure 12. The psychometric function in experiment 1 with the median correlation between p_{lin} and d


100

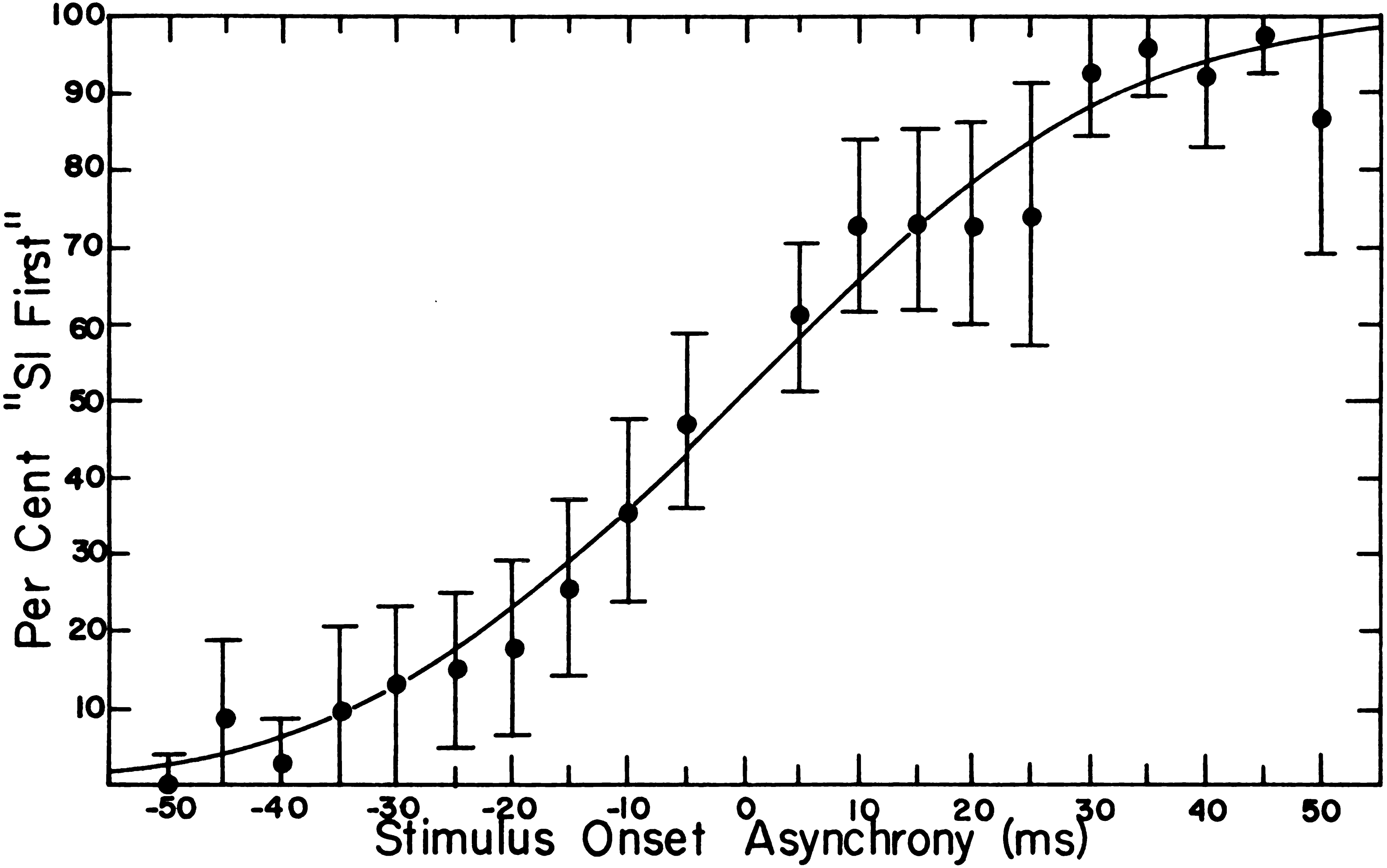

Figure 13. The psychometric function in experiment 1 with the lowest correlation between p_{lin} and d

correlations are based on the same number of data points per function, and on a nearly fixed (across but not within functions) number of observations per point. So this correlation should provide a reasonable index of the extent to which any given F(d) was adequately approximated by a normal function. The remainder of the results of experiment 1 will be discussed in terms of the mean or point-of-subjective-simultaneity and the variance of the underlying normal density functions.

The Standard Deviation Results

The results of both monocular and binocular temporal order judgments after probit analysis was applied to the psychometric functions are listed in table 2. Also listed there are the monocular standard deviations, σ_M; the ratios σ_B/σ_M; and the ratios of the larger to the smaller monocular variance. Figure 14 graphically displays the observed relationship of the binocular standard deviation to the monocular standard deviation along with the values of the binocular standard deviation predicted by the various models. In only one instance (MWH3) did the binocular standard deviation fail to be smaller than both the left and right standard deviations. But in no case was there a failure to see a reduction from σ_M^2, the average of the two monocular variances although this reduction was marginal for MWH3. Nor did any performance surpass the "baseline for independence," i.e., $\sigma_B < 0.707\ \sigma_M$, although GEN6 approached

Table 2

Parameters describing temporal order judgment psychometric functions in experiment 1

0	log above thres.	RIGHT			LEFT			BINOCULAR					
		μ_R	σ_R	n	μ_L	σ_L	n	μ_B	σ_B	n	σ_M	$\frac{\sigma_B}{\sigma_M}$	F
GEN6	1.0	2.488	31.959	1367	-12.717	30.795	1314	2.943	22.547	1416	31.382	0.718	1.077
GEN7	1.4	3.976	22.690	881	-7.482	23.668	871	-1.095	20.481	973	23.184	0.883	1.088
GEN8	0.6	0.362	27.525	928	-1.148	35.811	881	-0.615	25.806	910	31.938	0.808	1.693
MWH2	0.6	-11.773	42.387	801	-11.890	34.463	892	-10.342	31.452	910	38.629	0.814	1.513
MWH3	1.0	-2.230	34.785	963	-8.606	31.463	921	-2.081	32.672	931	33.166	0.985	1.222
MWH4	1.4	-8.374	27.016	886	-12.930	29.292	811	-7.625	24.406	885	28.177	0.866	1.176

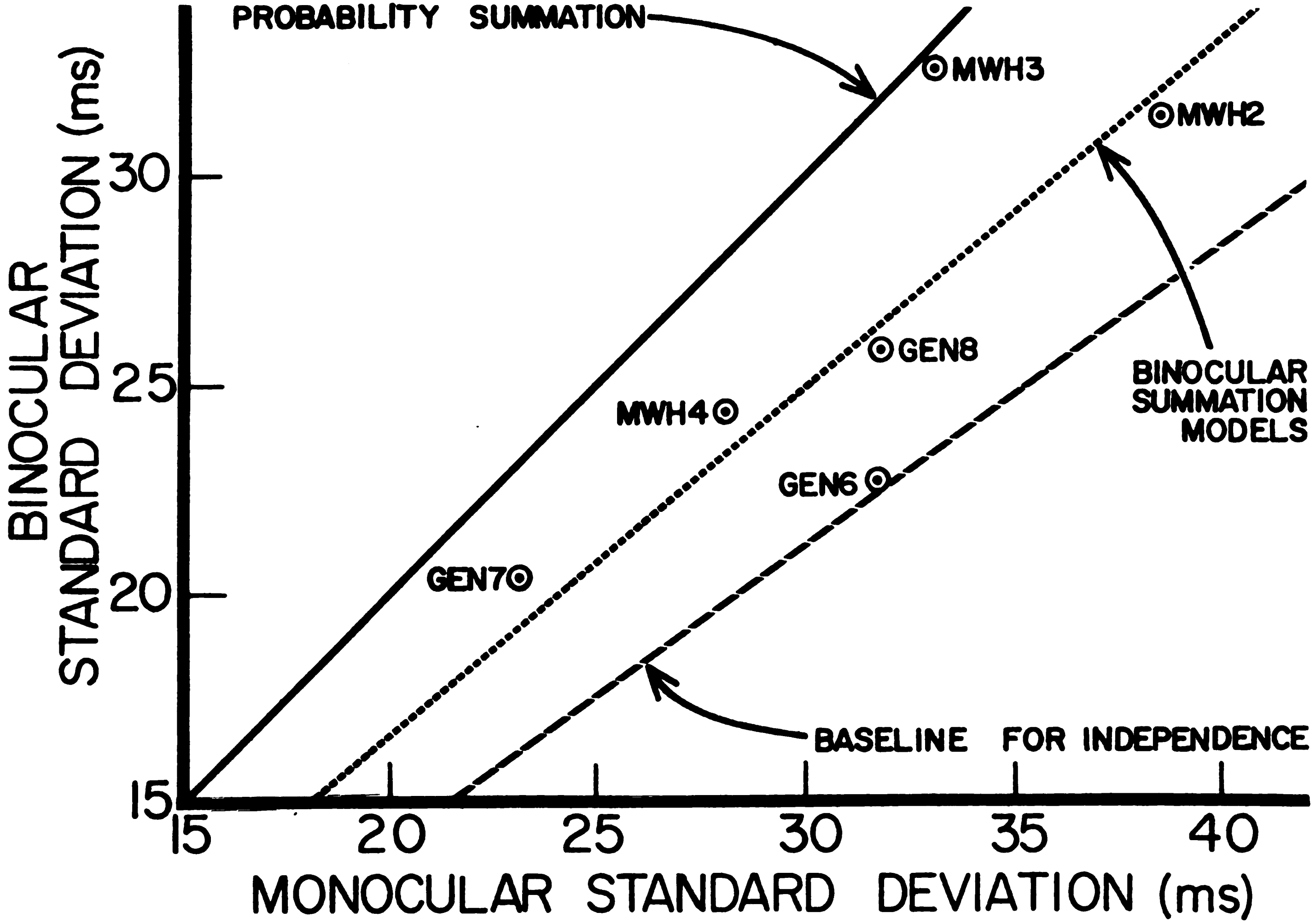

Figure 14. The binocular standard deviations plotted as a function of the monocular standard deviations for experiment 1

this level. In general, the relationship observed in experiment 1 between binocular and monocular standard deviations is less well approximated by a linear function than was the relationship observed in the pilot experiment ($r = 0.812$). However, when the data from both the pilot experiment and experiment 1 are plotted together as in figure 15 they do seem to belong to the same function. In fact, the best fitting line through the data of experiment 1 had slope of 0.768 (compared to 0.749 in the pilot experiment) and a σ_B intercept of 2.365 ms (2.660 in the pilot experiment). The observed ratios of binocular to monocular standard deviations had a mean of 0.845 which is very similar to that found in the pilot experiment and is very close to that predicted by the binocular summation models of 0.826. Although the slope of the best-fitting line is between that predicted by the two summation models and that by the so-called "baseline for independence," examination of figures 8, 14 and 15 strongly supports the binocular summation models over the alternatives presented here (the effect of "noise" in the data is to decrease the slope of the regression line.)

The average standard deviations were 31.060 ms, 30.915 ms and 26.227 ms for right-eye, left-eye and binocular viewing conditions, respectively. These are in a ratio of 1.00:1.00:0.84. A multiple linear regression on these standard deviations yields

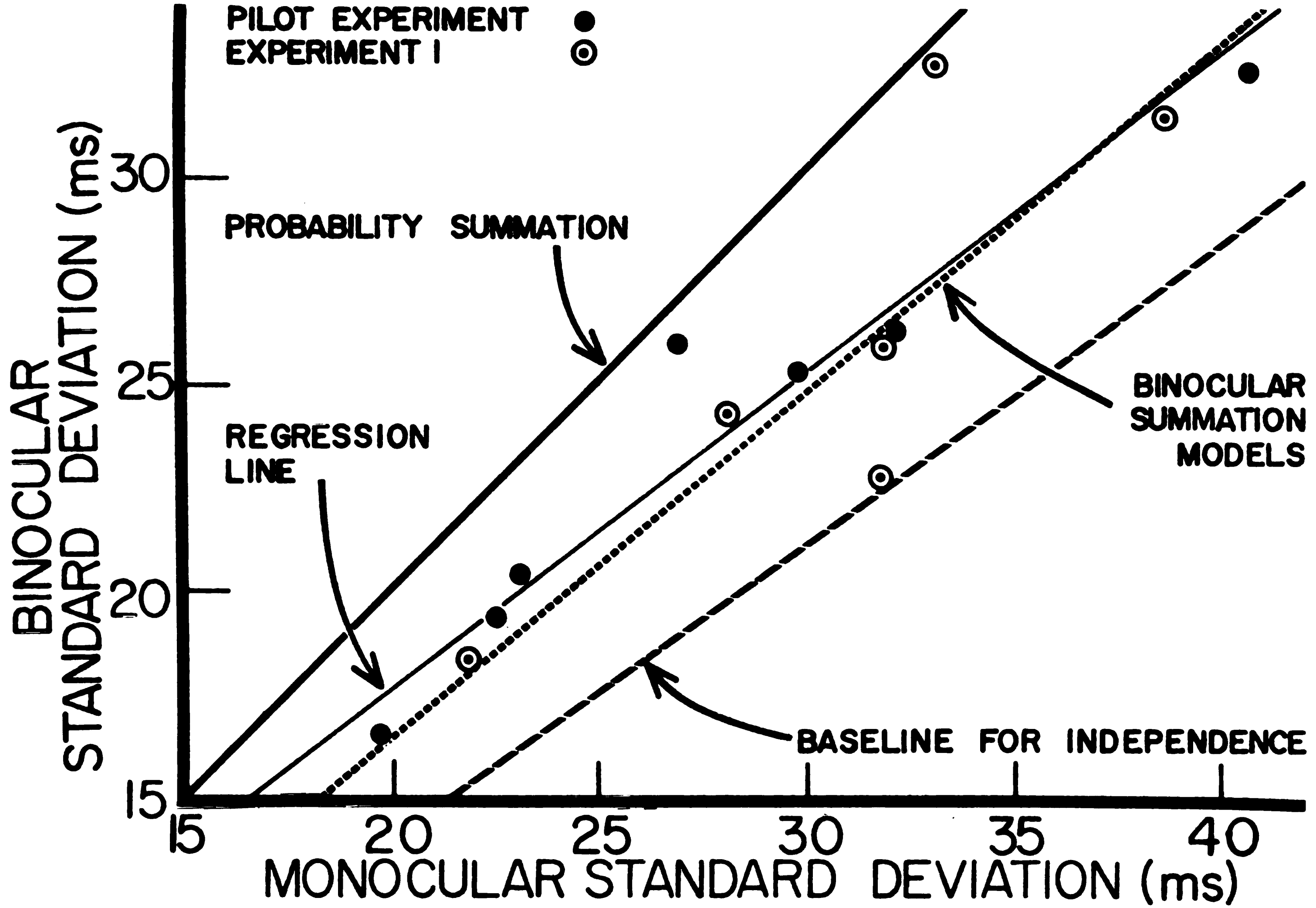

Figure 15. The binocular standard deviations plotted as a function of the monocular standard deviations of both the pilot experiment and experiment 1

$$\hat{\sigma}_B = 6.061 + 0.380\,\sigma_L + 0.287\,\sigma_R \tag{33}$$

and a multiple linear regression on the variances gives

$$\hat{\sigma}_B = \sqrt{167.364 + 0.133\,\sigma_L^2 + 0.438\,\sigma_R^2} \tag{34}$$

Unlike the case in the pilot experiment, the regression based on variance provided a slightly better fit (mean squared error of 6.957 compared to 7.611). If the data from the pilot experiment and experiment 1 are similar, then their combination is justifiable. The regression equations based on the combined standard deviations and variances are given by

$$\hat{\sigma}_B = 2.993 + 0.375\,\sigma_L + 0.378\,\sigma_R \tag{35}$$

$$\hat{\sigma}_B = \sqrt{102.147 + 0.236\,\sigma_L^2 + 0.369\,\sigma_R^2} \tag{36}$$

respectively. These two equations are approximately equal in their ability to account for the data (mean squared error of 4.114 and 4.168, respectively).

An analysis of variance was applied to the standard deviations of experiment 1 (see Appendix H). It did not, however, yield any significant difference between the viewing conditions. Given the small size of the effect, such an analysis would be incapable of detecting it given the variability inherent in the data. The only significant effect out of the observer, stimulus luminance and ocular viewing condition, was the stimulus luminance effect ($p < 0.05$). The mean standard deviations were found to be 32.907 ms, 30.704 ms and 24.592 ms for stimulus luminances 0.6, 1.0 and 1.4 log units above threshold.

As was done in the pilot experiment, a power function of stimulus intensity was fitted to each observer's standard deviations. These are shown in figure 16. This produced coefficents of 27.829 ms and 22.482 ms for GEN's combined monocular performances and binocular performance, respectively, and coefficients of 32.315 ms and 28.817 ms for observer MWH. Exponents corresponding to these were -0.329, -0.272, -0.355 and -0.269, respectively. These result in ratios of binocular to monocular coefficients of 0.808 for GEN and 0.895 for MWH. These are clearly in the range expected by the binocular summation models.

The ratio of the coefficients of the binocular and monocular power function is meaningful, however, only to the extent that the data are adequately approximated by a power function and to the extent that the exponent for monocular and binocular viewing is the same. A power function did in general provide a somewhat better fit when the alternatives were linear, exponential and logarithmic functions. The finding of Zacks (1973) also supports the power functions presented here. Using approximately the same stimulus conditions as were used in experiment 1,

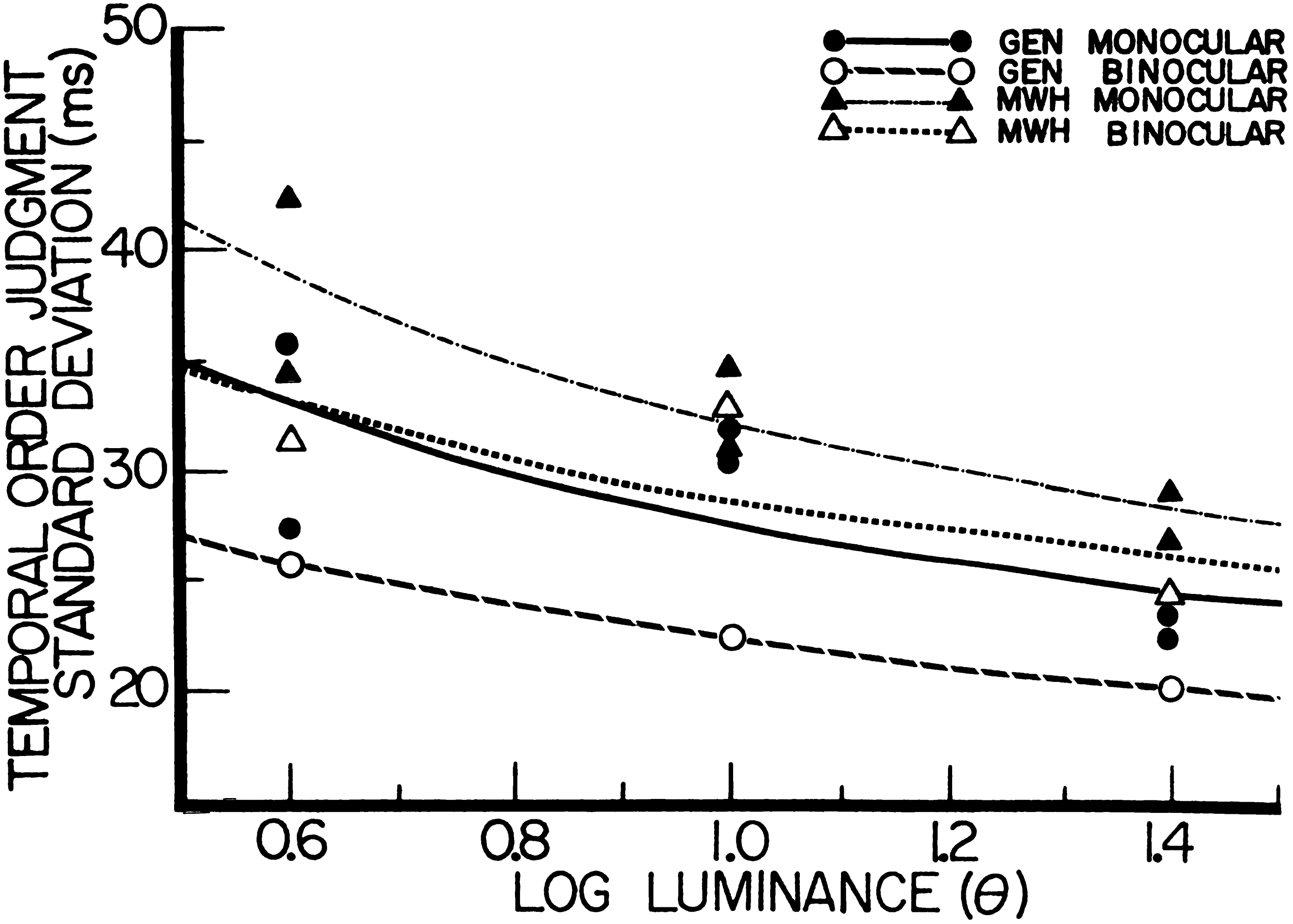

Figure 16. Standard deviations plotted as a function of log luminance for experiment 1 along with the best-fitting power functions

Zacks found standard deviations of 19.1 ms at 2.4 log units above threshold and 10.6 ms and 17.4 ms at 4.0 log units above threshold. The power functions described here would predict standard deviations at these higher luminances of 20.9 ms and 17.7 ms, respectively, for observer GEN, and 22.8 ms and 19.9 ms for observer MWH. Since these predictions are not inconsistent with those found by Zacks at these much higher luminances, a power function probably does provide a reasonable fit for the range of 0.4 to 4.0 log units above threshold.

The exponents for the monocular and binocular power functions are similar showing a mean value of -0.329 (standard deviation of exponents was 0.048). There may, however, be some tendency for the exponent of the binocular power function to be slightly nearer zero than that of the monocular functions, but there is insufficient data here to make any such determination. If this speculation is true, however, then the ratio of the power functions would also be a power function of luminance and not a constant. On the basis of the data available, this power function would be expected to have a coefficient of about 0.847 (the average ratio of coefficients) and exponent of about +0.023 (the difference between the average binocular exponent minus the average monocular exponent). When a power function is fit to the σ_B/σ_M ratio as a function of luminance, it results in a coefficient of 0.843 and exponent of -0.012. In addition to the exponent showing the wrong sign, this function provides a very poor fit to the data (see figure 17). Further, the power function of the ratio, even if it were correct, would differ markedly from constant only for very small and for very large luminances. The present models were intended only for a more restricted range of stimulus intensity.

Before leaving these power functions, it might be noted in passing that experiment 1 yielded larger standard deviations than did the pilot experiment. For observer GEN, the coefficient of the power function in the pilot experiment was 75.9 per cent and 78.8 percent of the monocular and binocular standard deviations found in experiment 1. In fact, as can be seen in figure 15, GEN's performance at 1.4 log units above threshold is exceeded by GEN's performance at only 0.7 log units above threshold in the pilot experiment. A similar effect is evident for observer MWH. Since the stimuli in the pilot experiment were separated by only 1° of visual angle, whereas those in experiment 1 were separated by 3° 23', this finding is consistent with the observations of Oatley, Robertson and Scanlan (1969) that there is some decreased sensitivity to temporal order with increased spot separation about a central fixation point and attributed this to a "slight increase in the variability of conduction time from extra-foveal as compared with near foveal regions of the retina" (p. 175). It is, however, not possible to determine whether the decreased sensitivity found here is due to this retinal location

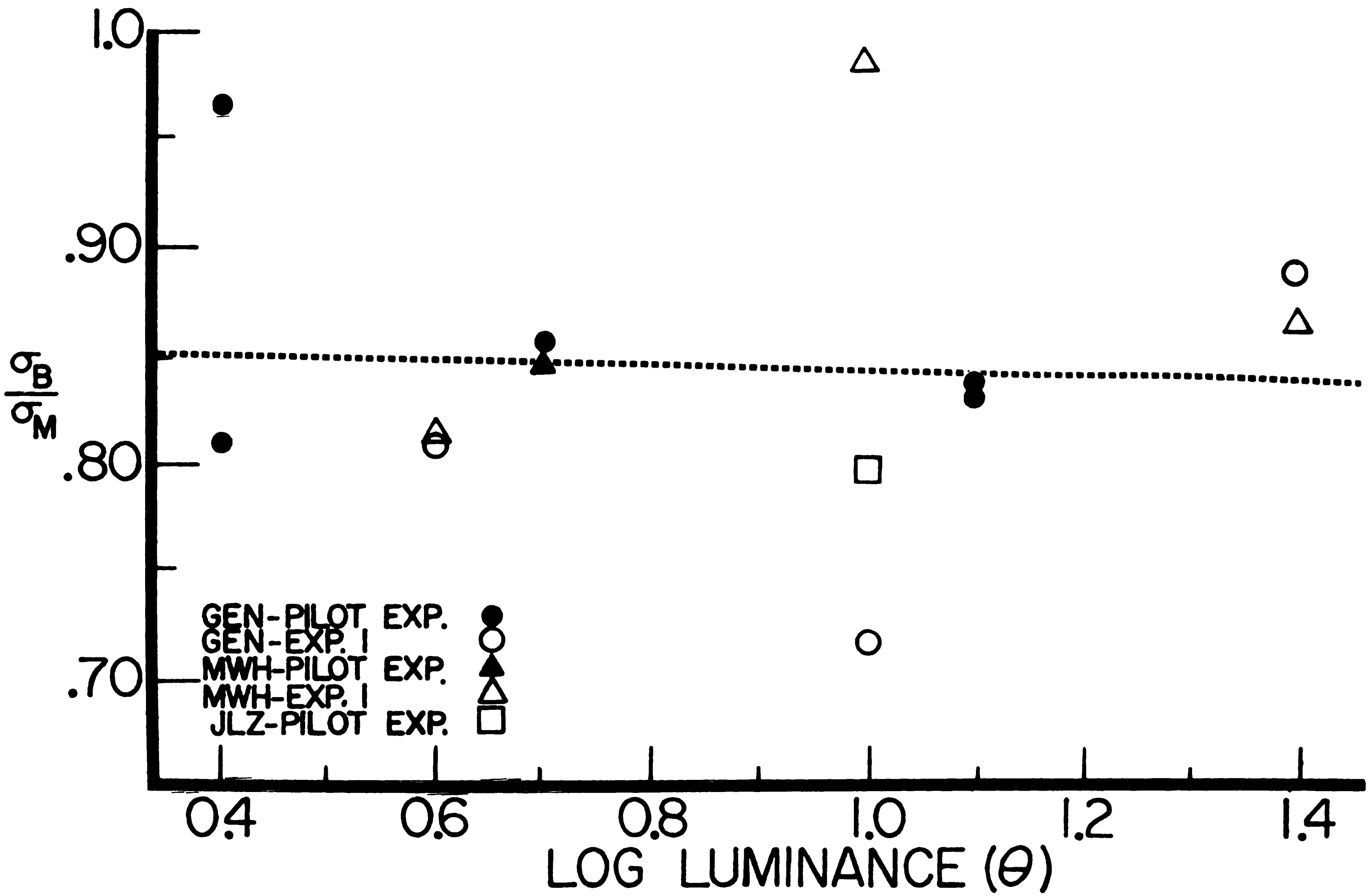

Figure 17. The ratio σ_B/σ_M plotted as a function of log luminance for both the pilot experiment and experiment 1 along with the best-fitting power function

effect, or to the spot separation itself, since both of these are confounded in this experiment. Westheimer and McKee (1977) examined judgments to temporal order with closely adjacent stimuli. Their stimuli were two vertical lines, 6' long and 15" wide and 10 ms in duration. They found an optimal separation of 2' to 6' where the threshold for temporal order was only 3 to 4 ms. Performance declined with larger and smaller separations. The decline with closer separations could be attributed to substantial overlap of the image spread of the two stimuli. These authors further examined the effect of orientation on this effect and concluded that "Insofar as it can be satisfied, relative orientation and position do not play a significant role" (p. 889). Hence, it may be tentatively concluded that the decreased sensitivity with spot separation was not due to a retinal location effect, but more likely to a spot separation effect.

Although the luminance effect was the only significant effect observed in the analysis of variance, a large F-ratio was also obtained for the observer effect, i.e., observer GEN's average standard deviation of 26.809 ms was lower than MWH's of 31.993.

The Point-of-Subjective-Simultaneity Results

The point-of-subjective-simultaneity or the mean of the psychometric functions was predominantly negative as were the monocular means of the pilot experiment. These means were also of about the same magnitude. An analysis of variance (see Appendix H) was applied to the temporal order means and yielded two significant main effects and one significant interaction. First, the observers differed significantly ($p < 0.0025$) in that observer GEN's average point-of-subjective-simultaneity was -1.476 ms while observer MWH's average was -8.428 ms.

The viewing condition main effect was also significant ($p < 0.02$) with averages of -2.592 ms, -9.129 ms and -3.136 ms for right-eye, left-eye and binocular viewing, respectively. Recall that a more negative psychometric function mean is interpreted in the context of the independent channel model of temporal order judgment as meaning the S1 had a shorter arrival latency than did S2; conversely, a more positive mean implies that S2 had a shorter latency. These results then imply that although all of the average means were negative, S2 had a relatively shorter average arrival latency compared to S1 under right-eye viewing than under left-eye viewing where the converse held. Binocular viewing tended to yield an intermediate value which was much closer to the mean for the right-eye viewing. Since S1 is the stimulus to the right of the central fixation point and S2 is the stimulus to the left, this result can be restated that for monocular viewing, the stimulus in the nasal hemifield which was imaged on the temporal retina had a shorter mean arrival latency than did a similar stimulus in the temporal hemifield which was imaged on the nasal retina.

This result is opposite to that which would have been hypothesized on the basis of documented properties of nasal hemiretina. Maddess (1975) notes a number of tasks which show nasal hemiretinal superiority. In particular, a stimulus delivered to nasal hemiretina is thought to exhibit a shorter latency than a similar stimulus delivered to temporal hemiretina. There is much evidence that simple reaction time is considerably faster in nasal than temporal hemiretina (Gilliland and Haines, 1975; Maddess, 1975; Poffenberger, 1912; Rains, 1964). This effect of retinal position on simple reaction time will be discussed in a later section of this paper. In the present temporal order judgment results, the effect suggests a latency difference of 3.3 ms (half of the difference between the monocular means) in favor of faster latency in temporal retina.

The only significant interaction was the observer by luminance interaction shown in figure 18. Significance for this interaction does not appear to be the result of stimulus manipulations but is probably a procedural artifact. All of the blocks of data for a given observer were collected in successive sessions. In the general model for monocular temporal order judgment presented earlier, the observer's decision criterion was zero (see equation 12). That is, if the response to S1 preceded the response to S2, then the observer said that S1 came first. If the criterion were other than zero, the effect on the psychometric function would be to move it toward the displaced

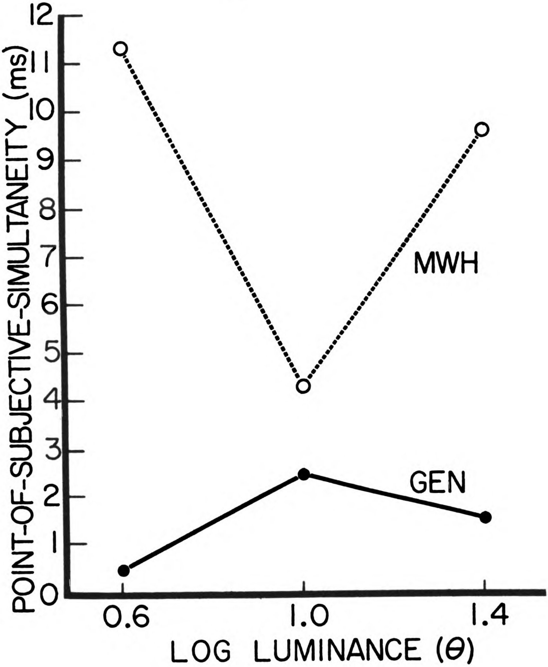

Figure 18. The significant luminance by observer interaction in experiment 1

criterion while maintaining the same shape. In other words, a nonzero criterion would produce shifts in the mean of the psychometric function but would have no effect on the standard deviation. If such a decision criterion were to change slowly over time then it would produce large changes in the point-of-subjective-simultaneity between seesions in experiment 1 and, hence, a significant observer by luminance interaction with no systematic relation to luminance. If the decision criterion in temporal order judgment is nonzero and if it depends on motivational and other nonsensory variables as does the criterion in signal detection theory, then the notion of measuring the sensory latency difference with the point-of-subjective-simultaneity may be suspect.

The binocular psychometric function means for both observers in experiment 1 could have best been predicted by the following multiple linear regression equation

$$\hat{\mu}_B = 0.769\,\mu_R - 0.096\,\mu_L - 2.022 \qquad (37)$$

which predicts μ_B with a mean squared error of 2.277 ms. It is worth noting again that the value of the binocular mean seems to be most closely related to the right-eye mean.

Discussion

The results of experiment 1 by confirming the results of the pilot experiment have established that binocular viewing does, indeed, produce a small but reliable increase in sensitivity to temporal order over monocular viewing as measured by the variance underlying the psychometric function. The question now arises, "Given this small binocular superiority, which model, if any, of those described here best accounts for the data?" In terms of the psychometric function standard deviations, we can unreservedly say that the best prediction is made by the low- and high-threshold binocular summation models which predict the same level of performance. We can intuitively reject the alternative hypotheses of probability summation, on one hand, and of that proposed by the likelihood ratio model, the average arrival time model, and the most discrepant arrival time model by simply referring to figure 15. In fact, these alternative models seem to provide only the lower and upper bounds of the effect. That is, no data set produced a σ_B / σ_M ratio which was greater than one or which was less than the "baseline for independence." These points, furthermore, tend to cluster very close to the line predicted by the two binocular summation models. The overall average ratio of binocular standard deviations to monocular standard deviations was 0.847, compared to 1.0 predicted on the basis of probability summation, 0.826 predicted by the two summation models and 0.707 predicted by the remaining models. The fact that the observed average ratio was slightly larger than expected by the binocular summation models is consistent with the notion of a small, additive, constant variance σ_0^2 which was included in equation 23, but was assumed to be neglible when the ratio 0.826 was determined.

One argument against accepting the binocular summation model over the other models presented could go something like this: If in the judgment of temporal order there is another source of variance besides the sensory variance which arises from errors in performance, then the additional variance would not vary with ocular presentation and would tend to make binocular and monocular performances relatively more similar so that the ratio σ_B / σ_M would be displaced closer to 1.0. Such an argument could account for a given displacement of some larger effect toward the level predicted by probability summation in figure 15. If this performance error variance can be represented by an additive model as would be expected for variance arising in the decision mechanism because of a non-deterministic decision rule (Sternberg and Knoll, 1973), then the ratio of the observed binocular variance, $\sigma_{B,obs}^2$, and the observed monocular variance, $\sigma_{M,obs}^2$, would be given by

$$\frac{\sigma_{B,obs}^2}{\sigma_{M,obs}^2} = \frac{\sigma_{B,sensory}^2 + \sigma_{error}^2}{\sigma_{M,sensory}^2 + \sigma_{error}^2} \tag{38}$$

where $\sigma^2_{B,\text{sensory}}$ and $\sigma^2_{M,\text{sensory}}$ are the true sensory variances which have been described throughout this paper and σ^2_{error} is the additive error variance that arises in performance. Notice that equation 38 implies that error variance is constant regardless of viewing condition. Notice also that if the ratio of the true binocular to true monocular variance were indeed 0.5 as predicted by a number of models in this paper, then the actual observed ratio could take any value from 0.5 to 1.0, depending on the relative contribution of the performance error variance. However, if the performance error variance is constant, then equation 38 predicts a nonlinear function of the observed monocular variance. When the true sensory variances are small, then the observed variance ratio should be very nearly equal to 1.0 since equation 38 would be dominated by the performance error variance in both numerator and denominator. When the sensory variances get very large, on the other hand, the effect of the performance error variance becomes increasingly more negligible and the observed variance ratio approaches 0.5. The data of figure 15 do appear quite linear and so lend little support for this interpretation.

In addition to this variance due to performance of the decision mechanism, there may be other sources of additive variance which do maintain linearity. For example, the effect of small involuntary eye movements or momentary lapses in convergence upon the fixation point would tend to cause the stimuli to fall on non-corresponding regions of the retinae on some proportion of the binocular trials. On most tasks, stimulation on non-corresponding areas sufficiently disparate to cause diplopia eliminates binocular summation or reduces it to levels expected on the basis of probability summation (Battersby and Defabaugh, 1969; Blake and Fox, 1973; Eriksen and Greenspon, 1968; Shaad, 1935; Thorn and Boynton, 1975; Westendorf and Fox, 1977). Since probability summation is not expected to play a role in binocular temporal order judgment, these disparate viewings could lead to binocular performance which is not any better than monocular trials. If it is assumed that the proportion, p, of trials on which the observer is convergent is the same across all binocular trials and at all intensity levels, and that the effect is dichotomous so that the observer is either performing at the monocular or at the theoretical binocular level, then the observed binocular to monocular ratio can be modeled as

$$\frac{\sigma_{B,obs}^{2}}{\sigma_{M}^{2}} = p^{2}\,\frac{\sigma_{B,sensory}^{2}}{\sigma_{M}^{2}} + q^{2} \tag{39}$$

where q = 1 - p. Notice that this predicted relationship is linear and could serve as an explanation for the small effect observed in the temporal order judgments. If this explanation is correct, however, then the observer would have to have been non-convergent on about 84 per cent of the trials in order for the performance to drop from the level predicted by the likelihood ratio model, the average arrival time model and the most discrepant arrival time model to the average level observed here. Further, reasonably large disparities (6') which still allow fusion do not eliminate binocular summation (Westendorf and Fox, 1977). It does seem more reasonable, then, to accept the binocular summation models.

Thus far throughout this paper the assumption has been made that the two monocular standard deviations, σ_L and σ_R, were identical and so could be combined to a single monocular standard deviation. This was done primarily because the data seemed to support it (see the F-ratios in tables 1 and 2) and because it allowed for a convenient two dimensional graphical display of the relationship (see figures 8, 14 and 15). The multiple linear regression equations on the parameters of the psychometric functions in the preceding results sections may allow an examination of the parameters of the binocular psychometric function as a joint function of the parameters of the two monocular functions. Multiple linear regression on the variances is particularly important because some of the models (for which predictions of binocular performance with unequal monocular performances were derived in the appendices) do describe binocular variance as a linear combination of the left-eye and right-eye variances. Before considering these, however, one should be reminded that because of "noise" variability in the data, the linear regressions will tend to (1) underestimate the coefficients and (2) to overestimate the constant term. These effects were evident for the case of the linear regression in figure 15.

The average arrival time model (see Appendix B) predicts that

$$\sigma_B^2 = \omega^2 \sigma_R^2 + (1 - \omega)^2 \sigma_L^2 \tag{40}$$

where ω is positive and less than one. The most discrepant arrival time model predicts this relationship as well, but because it has no mechanism to bias the inputs as it is described here, it requires that ω be equal to 0.5 (see Appendix C). No predictions were developed for either the low-threshold or the high-threshold versions of the binocular summation model (see Appendices D and E) for the binocular variance when monocular variances are unequal. It is, however, unlikely that these versions would lead to predictions of σ_B which are linear combinations of σ_R and σ_L. Little can be said about the appropriateness of any of the models in regard to binocular temporal order judgment variance.

A similar linear regression was performed on the mean of the psychometric functions in Experiment 2 and led to equation 37. Here it was noted that the binocular point-of-subjective-simultaneity seemed to depend most heavily on the right-eye performance. The average arrival time model (see Appendix B) predicted a weighted average according to

$$\mu_B = \omega\mu_R + (1 - \omega)\mu_L \tag{41}$$

where ω was positive and less than one. The most discrepant arrival time model (see Appendix C) predicted a similar average with $\omega = 0.5$. Again, no predictions were developed for either the low-threshold or the high-threshold versions of the binocular summation model (see Appendices D and E) for the binocular performance when the monocular performances are other than identical. No conclusions can be drawn from the data to favor any one model over the others.

Before concluding this discussion, some comparison of the present effect on the variance underlying the psychometric functions with other suprathreshold binocular summation phenomena is in order. Here will be considered only the putative binocular enhancement of brightness mentioned in the introduction to this paper. Until recently, the experiments of DeSilva and Bartley (1930) and of Fry and Bartley (1933) constituted the main evidence for the existence of binocular brightness enhancement (Blake and Fox, 1973). In the experiment of Fry and Bartley (1933) observers viewed two small rectangular targets separated by one degree of visual angle. The luminance of the top target, which was monocularly viewed, was set at various levels, and at each level the observer adjusted the intensity of the binocularly viewed lower target. At all intensity levels measured, the binocular target luminance had to be reduced in order to obtain a brightness equal to that of the monocular target. These results, replotted on log-log coordinates, are shown in figure 19. This relationship is highly linear ($r = 0.996$) and can be summarized by the following linear regression equation:

$$\log I_B = 0.835 \log I_M - c \qquad (43)$$

where I_B is the luminance of the binocular target required to match the brightness of the monocular target of luminance

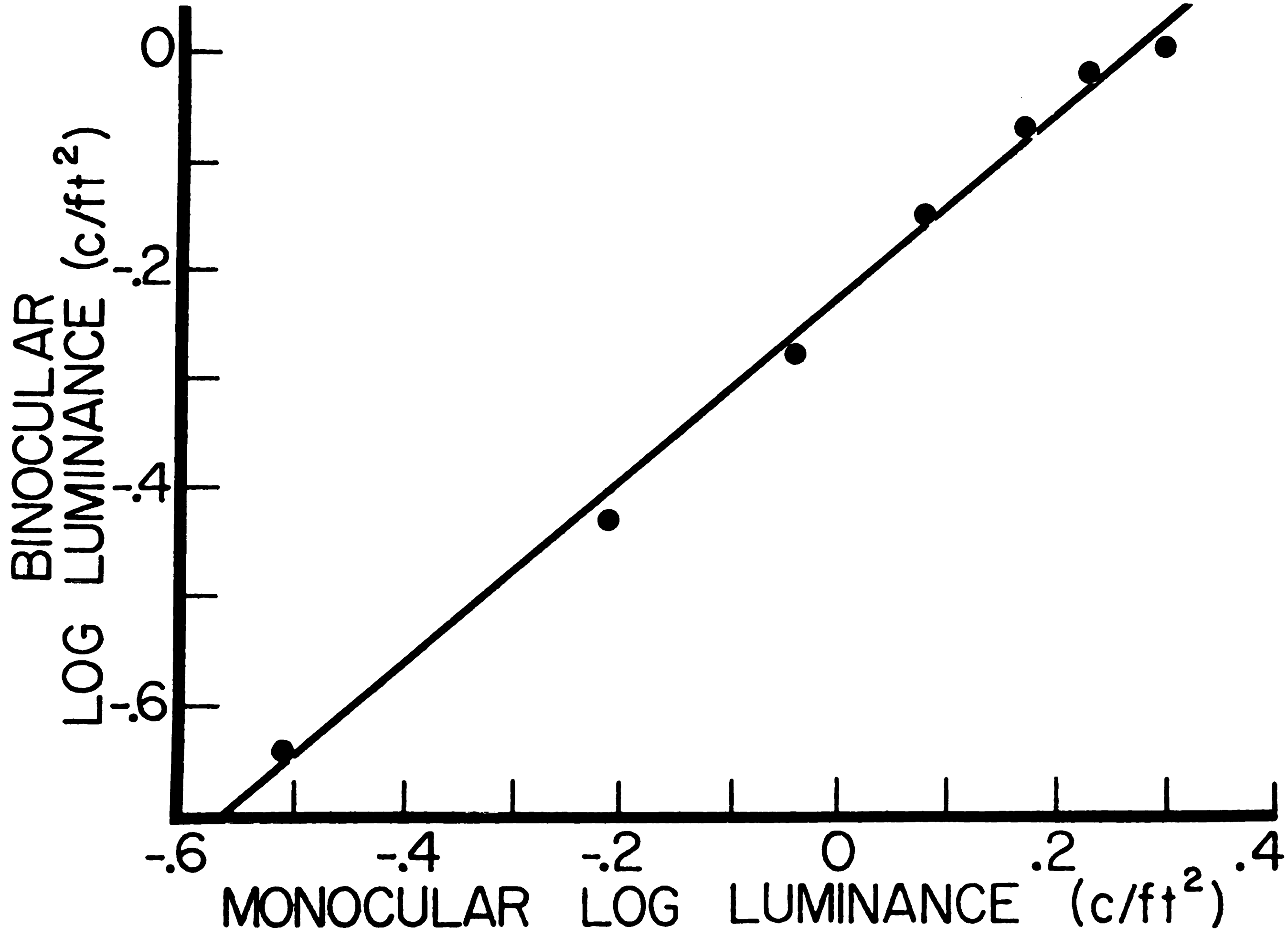

Figure 19. Log luminance required to match the brightness of a binocularly viewed target to the brightness of a monocularly viewed target (after Fry and Bartley, 1933)

I_M, and where c is a constant (for a given observer) given by

$$c = 0.230 + 0.165 \log \theta \qquad (44)$$

where θ is the observer's threshold (in c/ft^2). The effect of the constant c is to make both I_B and I_M measured relative to threshold.

If equation 43 is solved for I_M and if the variable I_B is relabeled simply as I, then we find the approximate intensity, I_M, that the monocular target would have to have been increased to in order to match the brightness of a given binocular target of intensity I. That is,

$$\log I_M = 1.198 (\log I + c) \qquad (45)$$

The question now arises, "Is the increased sensitivity to temporal order with binocular viewing consistent with the binocular enhancement of brightness?" or, more specifically, "Is the level of sensitivity with binocular viewing approximately that which would have been expected for monocular viewing if the luminance were simply increased to produce the same brightness as would have resulted for binocular viewing?" In the pilot experiment and experiment 1 it was shown that the monocular psychometric function's standard deviation could be satisfactorily modelled by a power function, at least for a restricted range of stimulus intensity.

So for a stimulus intensity I, the monocular standard deviation could be estimated according to

$$\sigma_M = a \log^b I \tag{46}$$

where a and b are the coefficient and exponent, respectively. Binocular viewing should produce a brightness equivalent to increasing the stimulus intensity according to equation 45. The binocular standard deviation would be estimated by

$$\sigma_B \approx a(1.198 \log I + 1.198\ c)^b \tag{47}$$

and the critical ratio found by dividing the appropriate halves of equation 46 by those of equation 47 would be estimated by

$$\sigma_B / \sigma_M \approx (1.198)^b \left[1 + \frac{c}{\log I}\right]^b \tag{48}$$

In general, equation 48 is not constant with changes in I unless c is zero (which could occur only if equation 43 intersected the origin, or if "brightness threshold" intensity was the same for both binocular and monocular viewing). A lower limit can be placed on equation 48, however. Since c is always negative because log [illegible] is negative and less that 0.230 (see figure 19), since log I is always positive and, finally, since b is negative, it can be shown that

$$\widehat{\sigma_B/\sigma_M} \geq (1.198)^b \tag{49}$$

For the values of b estimated in experiment 1 (-0.329 for GEN and -0.355 for MWH), σ_B/σ_M must be greater than 0.942 and 0.938 for GEN and MWH, respectively. (These would have been 0.945 and 0.941 had figure 19 been "inverted" prior to applying the linear regression). These ratios are clearly larger than those found in this study or those predicted by the various models in this paper. Therefore, the binocular brightness enhancement must be a "smaller" effect than the binocular effect in temporal order judgment.

Most investigators have not found the results of DeSilva and Bartley (1930) and of Fry and Bartley (1933) conclusive. An alternative hypothesis which expects no binocular brightness enhancement of brightness is that the binocular brightness is the product of an averaging process between the two eyes (Levelt, 1965; Blake and Fox, 1973; deWeert and Levelt, 1974). Some workers have questioned the early studies on procedural grounds (Levelt, 1965; see also Blake and Fox, 1973). Further, more recent reports have not found the large binocular brightness enhancement reported by Fry and Bartley (1933). J. C. Stevens (1967), for example, had observers give direct magnitude estimations to scale both monocular and binocular brightness. The resulting functions of intensity were virtually identical, differing by only one-tenth of a log unit which is roughly one jnd. Stevens felt that this slight difference might be artifactual, but conceded that the method of magnitude estimation may lack the precision to reveal a summation effect. At any rate, the levels of binocular enhancement reported by Fry and Bartley (1933) are probably the greatest which have been reported. It follows then that the estimates of σ_B/σ_M based on Fry and Bartley's data must be viewed as exceptionally low estimates; values based on other reports would be considerably closer to 1.0. So, in the study of small binocular summation effects, the small increase in sensitivity with binocular viewing may be considered as a relatively "large" effect.

The fact that the increase in sensitivity to temporal order exceeds the level consistent with binocular brightness enhancement rules out one explanation for the effect observed here. It assures, at least, that the binocular superiority cannot simply be accounted for by the brightness enhancement effect. Had it been otherwise, one may have argued that the brightness enhancement in some way facilitated the binocular temporal order judgment. In terms of the general model of temporal order judgment offered in an earlier section, this view would imply that brightness somehow "causes" visual latency variability. Given these results, it makes more sense to think of brightness and latency as independent variables which correlate only because both covary with stimulus intensity. In this view, one cannot be said to cause the other.

EXPERIMENT 2

A Two-Component Serial Model of Simple Reaction Time

The pilot experiment and experiment 1 have presented evidence that binocular viewing increases sensitivity to temporal order over monocular viewing. However, the conclusion that this increased sensitivity results from a reduction in the sensory latency variability of the response hinges on the validity of the interpretation given by the general independent channel model. Even if this model is basically correct, it does not follow that the reduction in variability occurred in the sensory channels but may have been produced by a reduction in the variability of a probabilistic decision rule (Sternberg and Knoll, 1973). Experment 2 was intended to examine this sensory latency using a different method in order to see if converging evidence will support this interpretation.

Numerous techniques have been proposed for measuring perceptual latency means but there has been little effort to measure the variance of latency. Perhaps th most direct latency estimation is that of the simple reaction time where the experimenter simply measures the time from the onset of a stimulus until the time that the observer makes a response. Models of simple reaction time are typically postulated with reaction time being composed of at least two serial latency components (Gibbon and Rutschmann, 1969; Mansfield, 1973; Rutschmann, 1966; Rutschmann and Link, 1964; Sternberg and Knoll, 1973; Vaughan, Costa and Gilden, 1966): First, a <u>sensory-detection component</u> is proposed which, like the the arrival latency in models of temporal order judgment, includes the delay from the onset of the stimulus to the time it is detected. Vaughan, Costa and Gilden (1969) suggest that since both neurophysiological measures of visual latency and simple reaction time conform to similar power functions of luminance, the visual reaction time is determined at least in part by a visual latency component. For a given observer, the average reaction time is virtually a constant (with respect to intensity) from the visual evoked response lag which presumably corresponds to the time of detection. They also conclude, however, that since there are intersubject differences despite similar visual evoke response latencies, the sensory processes do not account for all of the average reaction time effects. Mansfield (1973), for example, used reaction time extensively to measure visual latency functions by simply subtracting out an "irreducible constant latency."

The second reaction time component is often called the <u>motor component</u>. The motor component has no counterpart in the model of temporal order judgment. It is defined as the delay from the detection to the time that the response is recorded. The motor latency is typically assumed to be variable but not systematically influenced by stimulus manipulations (Mansfield, 1973) so that if binocular viewing were to reduce reaction time variability over monocular viewing, the reduction would be attributed to a reduction in the variability of the sensory-detection component and not of the motor component. Although this motor component is assumed to be independent of stimulus properties it is also assumed to be affected by variability in "preparatory set for motor responses," and so to vary from individual-to-individual (Vaughan, Costa and Gilden, 1966) and so may be subject to "criterion" effects. If the reaction time sensory-detection component can be equated with the arrival latency in temporal order judgment, then an experiment comparing binocular to monocular visual reaction time variability could supply evidence to bear on the results of the pilot experiment and experiment 1.

Because of the hypothetical added motor latency in simple reaction time there are objections to directly equating the mean and variance of the reaction time distribution to those of the arrival latency distribution in temporal order judgment. First, the additional motor component in the reaction time must certainly require that the mean reaction time be greater than the mean arrival latency. As a result, reaction times should overestimate the arrival latency mean. If, however, the motor component is truly independent of stimulating conditions, including whether S1 or S2 is presented, then the mean motor latency in the mean reaction time to S1, denoted $\overline{RT}_1$, and in the mean reaction time to S2, denoted $\overline{RT}_2$, must be equal and so cancel in the difference which would be used to estimate the point-of-subjective-simultaneity in temporal order judgment (see equation 14). Second, since the motor component of reaction time is assumed to be independent of and additive to the sensory-detection component, it should act to inflate estimations of the arrival latency variability in temporal order judgment. The motor variability should, in fact, act like the additive error variance which was described in the preceding discussion (see equation 38). To summarize, what is expected _a priori_ is that reaction time differences should correspond to the point-of-subjective-simulataneity but that reaction time variance should tend to overestimate the arrival variance in temporal order judgment.

Examination of the literature, however, reveals that reaction differences are very poor estimators of the point-of-subjective-simultaneity (Gibbon and Rutschmann, 1969; Rutshmann, 1967; Rutschmann and Link, 1964; Sanford, 1971; Silverstein, 1976; Sternberg and Knoll, 1973). In the bisensory experiment of Rutschmann and Link (1964) this test failed dramatically: Whereas the mean auditory reaction time was about 45 ms shorter than mean visual reaction time, the auditory stimulus had to be presented about 43 ms earlier than the visual stimulus to produce subjective-simultaneity. This led to the conjecture that the decision mechanism for temporal order was about 88 ms "further" from the auditory than the visual channel. Sternberg and Knoll (1973) point out, however, that this may as well be accounted for by assuming that the mean motor latencies were not equal. In the flash pair experiment, Gibbon and Rutschmann (1969) found that one of the two subjects showed a similar but smaller discrepancy whereas the other subject showed good agreement between reaction time means and the point-of-subjective-simultaneity. In the former case they speculate that the discrepancy may be due to a response bias which was present in the temporal order judgments but not in the reaction time data. Silverman (1976) compared the mean reaction time (both choice and simple) differences and the point-of-subjective-simultaneity for ten subjects at two different intensity levels where the stimuli were a rotary acceleration of the observer in an Ames Man-Carrying Rotation Device and the displacement of a visual stimulus on a cathode-ray tube. He found a significant correlation between mean choice reaction time differences and the point-of-subjective-simultaneity, although the sign of the correlation changed with intensity level, +0.63 in one case and -0.63 in the other. Similar but nonsignificant correlations were observed between simple reaction time and the point-of-subjective-simultaneity. In short, there seems to be only a weak relationship, if any at all, between the point-of-subjective-simultaneity and mean reaction time differences.

If reaction time variances are used to estimate the variance in temporal order judgment according to equation 15 then the additional variance in the reaction time task would be expected to inflate the estimates. However, Rutschmann and Link (1964) and Gibbon and Rutschmann (1969) found the shapes of temporal order judgment psychometric functions predicted from reaction time distributions to be surprisingly similar to obtained functions. This led them to conclude that the motor component in simple reaction time adds little variance relative to the variability in the sensory-detection component. However, Rutshmann (1967) found that an increase in stimulus intensity decreased the mean and variance of reaction time to brief shocks presented to the two hands, but had no systematic effect on the "slope" of the psychometric function for temporal order judgment and either no effect or an effect in the opposite direction on the point-of-subjective-simultaneity.

Sternberg and Knoll (1973) argue that even if the motor component's contribution is constant, the reaction time task and the temporal order judgment task may not even be related by the latencies of the same internal events. To have the same internal basis requires, first, that there be a "final common path" for reactions to stimulus onsets so that inputs to the temporal order judgment decision mechanism are the same as the inputs to the reaction time decision mechanism. Second, even if such a "final common path" exists, different features of the signals may serve for onset markers in reaction time and temporal order judgment. Afterall, the reaction time task requires speed with low false alarm rates, whereas temporal order judgment requires low variance to maximize precision. Thus, temporal order judgment might depend on the estimated time of the peak response and reaction time upon the delay before a response first exceeds a criterion level; The estimation of the time of a peak might have less sampling variance than the delay until a criterion response is reached because the latter, but not the former, may vary with trial-to-trial fluctuations in sensitivity. Vanderhaeghen and Bertelson (1974) also conclude that there are "parallel independent analyzing stages, respectively for (choice) RT (reaction time) and TOJ (temporal order judgment)" (p. 571). On the other hand, Siverstein (1976) concludes that for the stimuli he used, "the criterion-signal arrivals ... are similar to, if not the same as, those used to cue a judgment of temporal order" (p. 90).

It is apparent that the relationship between latency measured by temporal order judgment and reaction time is not settled. In experiment 2 reaction time distributions will be collected not so much to precisely estimate the parameters of the arrival time distribution, but only to assess the general direction of the effects of stimulus manipulation. There is ample evidence that reaction is influenced by the same set of stimulus parameters as is temporal order judgment and for most cases in the same directions as would be inferred by temporal order judgment (Gibbon and Rutschmann, 1969; Hansteen, 1971; Lewis, Dunlap and Matteson, 1972; Mansfield, 1973; Rutschmann and Link, 1964; Silverstein, 1976; Teichner and Krebs, 1972).

Binocular Summation and Simple Reaction Time

Blake and Fox (1973) highly recommend the investigation of binocular summation with reaction time (but as an indirect measure of brightness summation -- they assume that reaction time and brightness are somehow "linked"). Poffenberger (1912) concluded that reaction time under binocular viewing was faster than for monocular. Poffenberger's results have been questioned on methodological grounds, however (Teichner, 1954). Miles (1936), in an investigation of those conditions which decrease visual reaction time, found that viewing the stimulus binocularly significantly reduced reaction time compared to viewing it monocularly. More recently, Minucci and Connors (1967) examined the relationship between monocular and binocular simple reaction time to a brief light flash varied over a four log unit range. Not only did they find that reaction time decreased with intensity increments, but they confirmed the earlier results of Poffenberger and of Miles. They concluded that decrease in binocular reaction time was greater than that anticipated from complete summation. However, it is far from clear what kind of reaction time should be expected from the possible effects of probability summation (Blake and Fox, 1973). Minucci and Connors also found that monocular dominant-eye reaction time was faster than non-dominant-eye reaction time, although still slower than the binocular reaction time. In a review over an eighty-three year period, Teichner and Krebs (1972) compared studies of binocularly mediated foveal reaction with studies of monocularly mediated foveal reaction time and found the "the monocular condition did, indeed, tend to produce longer RT" (p. 352). Gilliland and Haines (1975) compared monocular to binocular reaction time for a range of stimulus positions from 90° arc left to 90° arc right of the line of sight along the horizontal meridian in 10° increments. They found reduced binocular reaction times in all positions with differences as large as 40 ms and as much as 22 ms faster than the fastest monocular reaction time, even for stimuli located far to the periphery (beyond 60°) which are assumed to be monocularly mediated due to nasal occlusion. Haines (1977) replicated this study, comparing reaction time along other meridians other than horizontal and using colored as well as "white" stimuli. Binocular reduction of reaction time was again found, even in the non-binocular region. The decrease in the absolute difference in reaction time between monocular and binocular viewing with retinal eccentricity was attributed to the decreased degree of retinal image correspondence, as well as to possible changes in sensitivity.

Gilliland and Haines (1975) did not find the reaction time advantage for the dominant eye as did Minucci and Connors (1967), but they did note a viewing condition by stimulus position interaction: Right-eye reaction times were faster than left eye reaction times when the stimulus was in the right hemifield, and left faster than right when in the left hemifield. That is, the reaction time was shorter for a visual stimulus delivered to the nasal retina than to the temporal retina. Both Hall and von Kries (1879, cited in Maddess, 1975) and Poffenberger (1912) had reported this nasal superiority in reaction time much earlier. It has also been reported by Rains (1963), Payne (1967), Mansfield (1973) and Maddess (1975). This result implies that faster reaction time results from stimulation of the portion of the retina which crosses to the contralateral brain regions. Maddess (1975) also found reduced reaction time for dominant hand reactions and overall superior performance for right-eye dominant and for right-hand dominant observers. The reduced reaction time for nasal stimulation cannot be accounted for by the conduction velocities of the fibers in the optic tract because these crossed fibers show, if anything, longer latency (Kirk, Cleland and Levick, 1975; Bishop, Jeremy and Lance, 1953) at least in the cat. It seems counterintuitive to imagine that the contralaterally projecting fibers are shorter than the ipsalateral. Nor can the effect be attributed to asymmetries in time perception due to cerebral lateralization (Polzella, DaPolito and Hinsman, 1977) since the enhancement occurs from both nasal retinae. Nasal superiority has been reported for a number of other tasks as well (see Maddess, 1975).

This nasal superiority for reaction time, however, does not apply to a spot about 15° horizontally from the fovea corresponding to the location of the blind spot in the other eye. Here, reaction time is shorter in the temporal retina (Rains, 1963; Payne, 1967). Further, the upper retina yields faster reaction time than does the lower retina (Payne, 1967). The commonly held explanation for all of these effects is that the distribution of reaction time to stimuli across the retina is highly correlated with the density distribution of the rods across the retina (Gilliland and Haines, 1975; Payne, 1967; Poffenberger, 1912; Rains, 1963). Payne (1967), for example, measured reaction at small intervals along a circle about the fovea and found that many details of the reaction time magnitude corresponded to the best available information on the rod density in the human retina. The complex reaction time functions found by Haines (1977) who measured reaction time in 10° increments across five retinal meridians might also be explained by receptor density.

The major concern of this paper is not, however, in the reaction time mean, but rather in the reaction time variance; It is a well-established observation that binocular viewing reduces reaction time means, but does it also reduce reaction time variance? Although they comment no further, Gilliland and Haines (1975) report standard deviations averaged over all of the stimulus conditions of 33.1 ms (range 28 to 42 ms) for the binocular testing condition and of 36.8 ms (range 28 to 59 ms) and 35.2 ms (range 30 to 49 ms) for the right-eye and left-eye monocular test conditions, respectively. This difference between binocular and monocular variance may have been significant for some of the stimulus conditions (the average values include those stimulus positions where nasal occlusion prevented full binocular view) and so the present quest is somewhat encouraged.

To summarize, although there are serious doubts about equating even the sensory components of simple reaction time and the arrival latency in temporal order judgment, experiment 2 will look for similarities between temporal order judgment predicted by reaction time distributions and the performance observed in the pilot experiment and experiment 1. This amounts to treating reaction time distributions "as if" they were arrival distributions in the model of temporal order judgment. If this is successful, the notion that reaction time measures provide a converging method to support the results of the temporal order judgment experiments is strengthened. Any variance effects may then be meaningfully compared to those inferred from the temporal order judgment experiments. Experiment 2 also provides a replication of the documented reaction time mean effects described above.

Models for Binocular Simple Reaction Time

To the extent that the arrival latency of the signal evoked in the judgment of temporal order is identical to the sensory-detection component of reaction time, then some of the models of binocular temporal order judgment have counterparts in models of binocular reaction time. Others do not. The likelihood ratio model, since it defines the response to a specific stimulus onset asynchrony as an abstract "state," does not specifically invoke the notion of arrival latency and so, as it stands, is not extendable to reaction time. The most discrepant arrival time model deals specifically with the arrival time differences from each eye and not with the individual arrival latencies of the signals from a given eye. As a result, it applies only to the judgment of temporal order. There is, then, no prediction offered by either the likelihood ratio model or the most discrepant arrival time model for the binocular reaction time.

Although it is mathematically equivalent to the most discrepant arrival time model, the average arrival time model is formally different; it allows for the arrivals of the two monocular signals referenced to the same stimulus in visual space to be averaged prior to being differenced. Such an average arrival time model could also be applied to the sensory-detection component of reaction time, where reaction time is assumed to be the simple sum of an independent sensory-detection component and motor component.

Let the monocular reaction time distribution for the S1 evoked signals be described by the means, $\overline{RT}_{L,1}$ and $\overline{RT}_{R,1}$, and the variances, $\nu^2_{L,1}$ and $\nu^2_{R,1}$, for the left and right systems, respectively. Then,

$$\overline{RT}_{L,1} = \mu_{L,1} + m \tag{50}$$

$$\overline{RT}_{R,1} = \mu_{R,1} + m \tag{51}$$

$$\nu^2_{L,1} = \sigma^2_{L,1} + \sigma_m^2 \tag{52}$$

$$\nu^2_{R,1} = \sigma^2_{R,1} + \sigma_m^2 \tag{53}$$

where m and σ_m^2 are the mean and the variance of the motor component. This motor component is assumed to be the same for all signals although it may vary from observer-to-observer. The mean latencies, $\mu_{L,1}$ and $\mu_{R,1}$, and the latency variances, $\sigma^2_{L,1}$ and $\sigma^2_{R,1}$, are assumed to be exactly the same as was put forth in the earlier models of temporal order judgment. The average arrival time model would then predict the binocular reaction time mean, $\overline{RT}_{B,1}$, and variance, $\nu_{B,1}$, as

$$\overline{RT}_{B,1} = 0.5\,\mu_{L,1} + 0.5\,\mu_{R,1} + m \tag{54}$$

$$\nu_{B,1} = 0.25\,\sigma^2_{L,1} + 0.25\,\sigma^2_{R,1} + \sigma_m^2 \tag{55}$$

In short, the average arrival time model predicts, in addition to a large reduction in the reaction time variance, that the binocular reaction time mean will just be the average of the two monocular reaction time means. That is, the mean binocular reaction time must lie between the two monocular means. This prediction conflicts with the numerous investigations already cited which have shown that the binocular reaction time mean is reliably less than either monocular reaction time mean -- a situation which could not be arrived at by averaging reaction time means.

The binocular summation models, since they also operate at the level of the signals arriving from different eyes in response to the same stimulus, may also be applied to the sensory-detection component of reaction time to predict a reduced binocular reaction time variance. If the monocular reaction time to S1 for the left and right systems is again described by equations 50 through 53, then the low-threshold summation model would predict that the sensory-detection component in the binocular reaction time would be the shorter of the two monocular sensory latencies and so the binocular reaction time distribution would be described by the mean, $\overline{RT}_{B,1}$, and the variance, $\nu^2_{B,1}$, as

$$RT_{B,1} = \overline{RT}_{L,1} - 0.5642\,\sigma_{L,1} \tag{56}$$

$$\nu^2_{B,1} = \nu^2_{L,1} - 0.3183\,\sigma^2_{L,1} \tag{57}$$

where the response to S1 is again assumed to be identical in each eye. The low-threshold model predicts a smaller reaction time variance decrement with binocular viewing than did the average arrival time model, but it also is consistent with the observed reduction in reaction mean with binocular viewing.

As with the low-threshold model, the high-threshold version also makes a prediction about the change in reaction time in the binocular condition. Now the sensory-detection component in the binocular reaction time would be the longer of the two monocular sensory-detection latencies so that the binocular reaction distribution would be described by the mean

$$\overline{RT}_{B,1} = \overline{RT}_{L,1} + 0.5642\,\sigma_{L,1} \qquad (58)$$

where again the two responses to S1 are assumed to be identical. Binocular reaction time variance would again be given by equation 57. Notice that this high-threshold version predicts a mean reaction time which longer than the monocular reaction time. This is, as noted, inconsistent with the observations made by several other investigators. Thus, if the sensory-detection component of reaction time is identical to the arrival latency in temporal order judgment, then all of the temporal order judgment models except the low-threshold binocular summation model may be rejected on the basis of the documented reduction of the reaction time mean with binocular viewing. Since the viewing conditions were randomized from trial-to-trial in experiment 2, a reduction in reaction time variance (if it is observed) could not be attributed to a reduction in the variance of the motor component through a "preparatory set for response," and so in the framework of the two-component, serial model of reaction time, must be attributed to a reduction in the variance of the sensory-detection latency. If this reaction time mean effect with binocular viewing is observed in experiment 2 and if there is some reduction in variability of the binocular reaction time, then not only is converging evidence supplied to support the explanation of the temporal order judgment effects of the pilot experiment and experiment 1, but the low-threshold model may be selected over the other alternatives presented.

Apparatus

The same apparatus was used in experiment 2 as was used in all data sets but one in experiment 1 with one minor modification: the bi-directional switch was replaced with a short-throw Microswitch which was connected to a short, spring-loaded lever. One data set was gathered with the same apparatus as was used in the pilot experiment with the same minor modification.

Procedure

For the reaction time trials, the observer was binocularly dark-adapted according to the method described in experiment 1 and seated in a darkened test chamber and presented with a moderate level of white noise through headphones. He pressed a lever connected to a microswitch to initiate a trial. This lever press was accompanied by a 100 ms high-pitched tone and the cessation of the white noise. The stimulus, either S1 (the right LED) or S2 (the left LED), was presented to either the left eye, right eye or both eyes randomly with equal probability of being any possible stimulus-eye combination on any trial. A random (rectangularly distributed) delay between 1048 and 2095 ms separated the cessation of the warning stimulus and the onset of the test stimulus. The observer released the lever immediately upon detection of the flash. This lever release was accompanied by a low-pitched 100 ms tone and the resumption of the white noise. Thus the observer controlled the intertrial interval. Fifteen blocks of 192 trials were performed on three days at each stimulus intensity. As in experiment 1 the first block of each daily session was omitted from further analysis to assure that a state of complete dark adaptation had been achieved and to provide the observer with some warm-up trials. Thus approximately 384 responses per condition were expected. For each stimulus by observer by intensity case, the reaction time distribution was characterized by its mean and

and variance. The reaction time itself was measured from the onset of the stimulus to the offset of the lever response.

The stimulus intensities for the 5 ms flashes were the same as those used in experiment 1, namely 0.6, 1.0 and 1.4 log units above the monocular response threshold.

In the one data set using the apparatus modified from the pilot experiment, the ocular condition was not randomized from trial-to-trial but was fixed within a block of 192 responses. The stimulus intensity for this single data set was 0.7 log unit above threshold. This data set, since it was not collected under comparable circumstances, was omitted from the analyses of variance performed on the other reaction time measures.

Observers

The same observers who participated in experiment 1 also served as observers in experiment 2, namely, the author (GEN) and one paid observer (MWH). Both produced data sets at the same intensities and with the same apparatus as was used in experiment 1. Observer GEN also produced one data set with the apparatus from the pilot experiment at 0.7 log units above threshold.

A Comparison of Reaction Time to Temporal Order Judgment Results

The means and standard deviations of the reaction time distributions from left-eye, right-eye and binocular viewing for both observers in all stimulus conditions are given in table 3. In table 4 are the parameters of temporal order judgment performance, μ_L, μ_R, μ_B, σ_L, σ_R, σ_B, σ_M, σ_B/σ_M, and the ratio of the larger monocular variance divided by the smaller, as estimated by simply replacing the arrival latency means and variances with reaction time means and variances in the appropriate equations. These values should be compared to the obtained temporal order judgment parameters listed in tables 1 and 2. The first question to ask is, "Do temporal order judgment and reaction time reflect decisions based on the same underlying sensory input?" If so, then reaction time must behave in a manner analogous to temporal order judgment for differences in the viewing conditions, luminance values and observers. The correlations between the point-of-subjective-simultaneity predicted by reaction time means and the observed point-of-subjective-simultaneity were -0.902 ($p < 0.05$), -0.509 and +0.451 for right-eye, left-eye and binocular viewing, respectively. The overall insignificant correlation was 0.003. There seems to be very little relationship between the point-of-subjective-simultaneity and reaction time means. This conclusion is also evident in figure 20 which graphically displays the obtained point-of-subjective-

Table 3

Performance on the simple reaction time task in Experiment 2

O	log above θ		RIGHT			LEFT			BINOCULAR		
			$\overline{RT}$	ν	n	$\overline{RT}$	ν	n	$\overline{RT}$	ν	n
GEN3	0.7	S1	248.41	39.53	784	249.62	40.03	744	238.35	42.16	720
		S2	263.68	53.01	732	279.40	56.25	766	244.16	44.38	777
GEN6	1.0	S1	246.957	40.746	540	243.580	39.273	526	226.974	38.791	535
		S2	271.362	53.762	530	282.100	55.372	519	243.200	38.735	580
GEN7	1.4	S1	233.062	40.079	387	234.035	40.073	402	219.302	40.114	354
		S2	256.028	55.350	386	265.594	57.513	382	232.935	43.303	369
GEN8	0.6	S1	259.464	52.625	390	258.935	50.171	372	233.716	43.988	359
		S2	279.595	51.483	390	305.869	68.642	413	252.590	44.091	356
MWH2	0.6	S1	333.390	89.878	362	317.681	77.026	404	284.710	60.036	324
		S2	340.908	75.833	404	340.138	70.103	370	311.317	77.180	388
MWH3	1.0	S1	274.202	61.810	381	276.064	57.067	409	263.263	68.050	361
		S2	284.218	69.628	377	289.550	68.640	393	270.415	61.432	359
MWH4	1.4	S1	273.139	65.061	404	266.768	59.356	400	257.809	51.767	372
		S2	278.198	63.537	364	280.185	71.184	384	267.376	72.654	356

Table 4

Temporal order judgment parameters as predicted

by the reaction time data of Experiment 2

O	log above θ	RIGHT		LEFT		BINOCULAR				
		$\hat{\mu}_R$	$\hat{\sigma}_R$	$\hat{\mu}_L$	$\hat{\sigma}_L$	$\hat{\mu}_B$	$\hat{\sigma}_B$	$\hat{\sigma}_M$	$\frac{\hat{\sigma}_B}{\sigma_M}$	$\hat{F}$
GEN3	0.7	-15.27	66.13	-29.78	69.04	-5.81	61.21	67.60	0.906	1.090
GEN6	1.0	-24.41	67.46	-38.52	67.89	-16.23	54.82	67.67	0.810	0.987
GEN7	1.4	-22.97	68.34	-31.56	70.10	-13.63	59.03	69.22	0.853	1.052
GEN8	0.6	-20.13	73.62	-46.94	85.02	-18.87	62.28	79.53	0.783	1.354
MWH2	0.6	-7.52	117.60	-22.46	104.15	-26.61	97.78	111.08	0.880	1.275
MWH3	1.0	-10.02	93.11	-13.49	89.26	-7.15	91.68	91.21	1.005	1.088
MWH4	1.4	-5.06	90.94	-13.42	92.68	-9.57	89.21	91.82	0.972	1.039

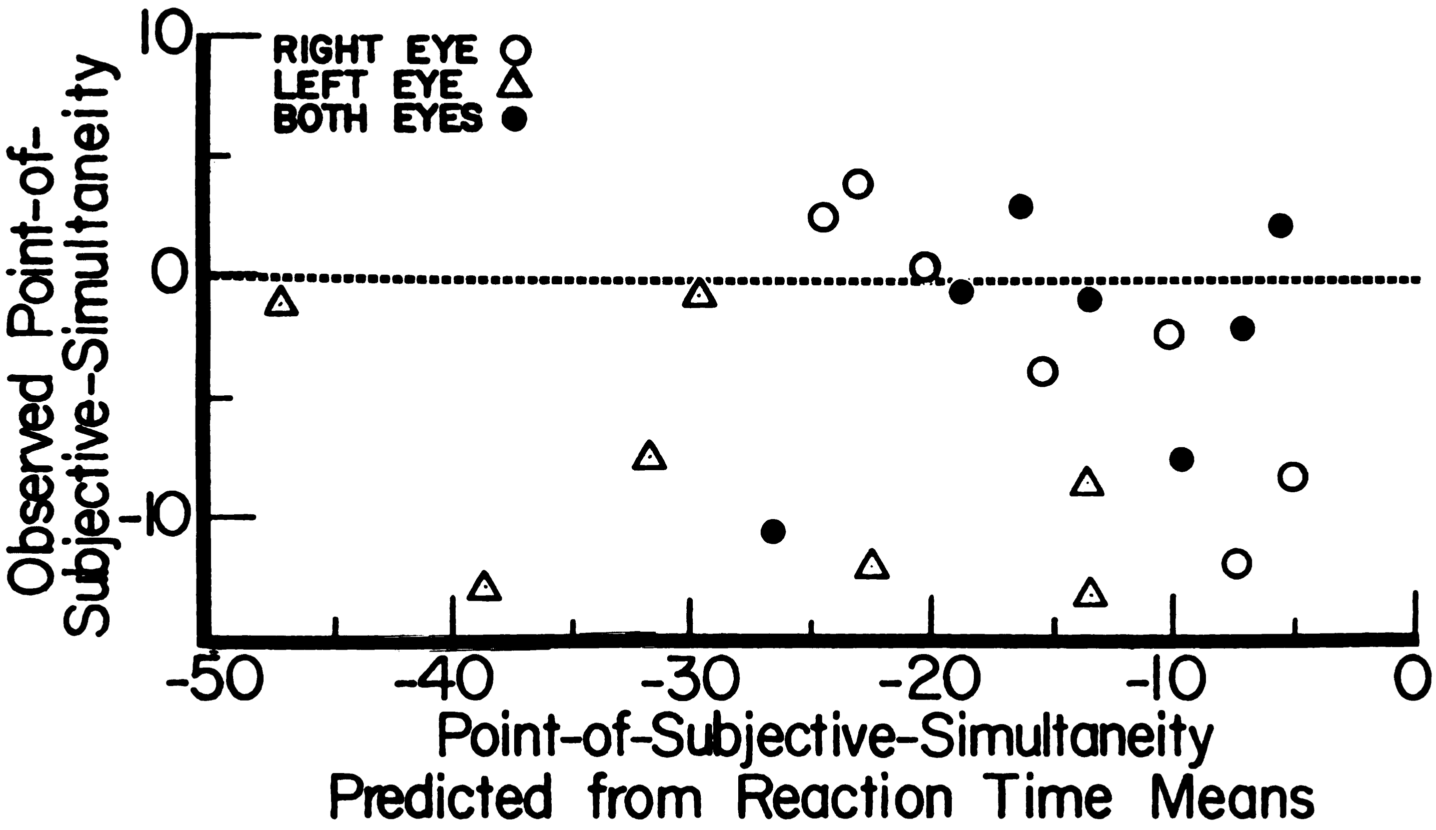

Figure 20. The relationship between the point-of-subjective-simultaneity as obtained in the pilot experiment and experiment 1 and as estimated from the reaction time means in experiment 2

simultaneity as a function of the predicted point-of-subjective-simultaneity. This is in accord with the other studies comparing reaction time to temporal order judgment (Gibbon and Rutschmann, 1969; Rutschmann and Link, 1964; Silverstein, 1976) in finding little relationship between mean reaction time and the point-of-subjective-simultaneity.

Correlations between temporal order judgment standard deviations predicted from simple reaction time variances and those obtained experimentally were 0.831 ($p < 0.05$), 0.665 and 0.810 ($p < 0.05$) for the right-eye, left-eye and binocular viewing conditions. The overall correlation of 0.795 was highly significant ($p < 0.001$) and implied that about 47 per cent of the variability in the obtained temporal order judgment standard deviations can be accounted for by a similar change in the reaction time with the same changes in conditions. Obtained temporal order judgment standard deviations are plotted as a function of the predicted standard deviations in figure 21 along with the linear regression equation. The correlation between the observed and predicted monocular standard deviation was 0.772 ($p < 0.05$). Further, the predictions based on reaction time failed only one time out of seven to predict which of the two standard deviations, left-eye or right, would be the larger for a given subject at any intensity level. In fact, the predicted temporal order judgment "F-ratio" in table 4 correlates with the obtained temporal order judgment "F-ratios" at 0.980 ($p < 0.001$).

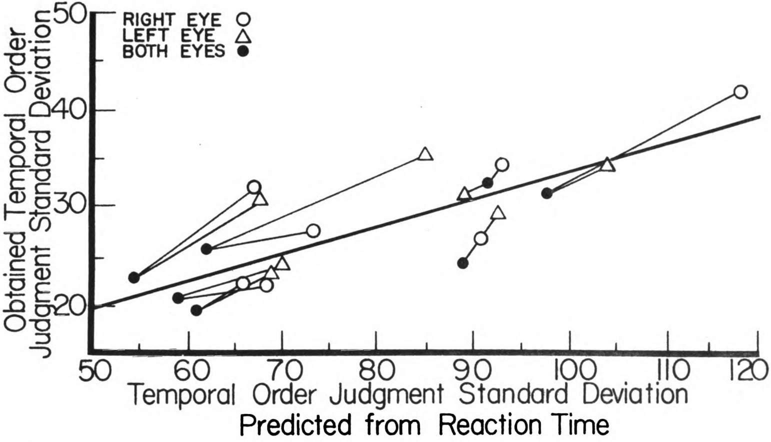

Figure 21. The obtained temporal order judgment standard deviations plotted as a function of the standard deviations predicted by the reaction time data along with the regression line, $y = 0.281x + 5.931$

In general, the temporal order judgment variance estimates based on reaction time results behave very similarly with experimental manipulations except that they are considerably larger -- 2.9 times as large on the average. This, of course, contradicts Gibbon and Rutschmann's (1969) claim that the motor component in reaction time adds little variance relative to the sensory variability. If the reaction time latency is, indeed, composed of independent sensory-detection and motor latencies and if the sensory-detection component is nearly the same as the arrival latency in temporal order judgment, then the motor variance should be equal to half of the difference between the temporal order judgment variance and the temporal order judgment variance estimates based on reaction time. These estimated reaction time motor standard deviations are listed in table 5. These estimates turn out to be relatively constant for a given observer: for observer GEN its mean is 43.798 ms (standard deviation of 4.744 ms) and for observer MWH 64.165 ms (standard deviation of 5.597 ms). There may be some tendency for the motor standard deviation to increase for the lowest intensities and for monocular viewing. Indeed, some differences between the motor variance estimated in this way and the "true" motor variance were expected even if the reaction time and temporal order judgments were based on the same sensory signal because the estimates in table 5 also include differences in decision criteria between the two tasks. All in all, the fact

Table 5

Estimated motor standard deviations
in the simple reaction time dats of Experiment 2

	O	log above	RIGHT	LEFT	BINOCULAR	MEAN
	GEN3	0.7	44.155	45.947	41.072	43.724
	GEN6	1.0	42.008	42.785	35.332	40.042
	GEN7	1.4	45.583	46.657	39.148	43.796
	GEN8	0.6	48.282	54.525	40.079	47.629
MEAN			45.007	47.478	38.908	43.798
	MWH2	0.6	77.567	69.497	65.467	70.843
	MWH3	1.0	61.070	59.066	60.718	60.285
	MWH4	1.4	61.400	62.176	60.674	61.417
MEAN			66.345	63.580	62.237	64.165

remains that the estimated motor variability is reasonably constant for a given observer and the observed temporal order variance corresponds to the variance estimated from reaction time variance. This lends strong support to the notion that temporal order arrival latencies and the reaction time sensory-detection component are very closely related.

The Reaction Time Mean Results and Discussion

Because of significance on an F-maximum test for homogeneity of variance ($p < 0.01$), analysis of variance was applied separately to the mean and standard deviations of the reaction times gathered under each condition. For the analysis on the reaction time means (see Appendix H), all four main effects (observer, luminance, stimulus S1 or S2, and ocular viewing condition) were significant at the $p < 0.0005$ level or better. Observer GEN's reaction times were significantly shorter than those of MWH (overall averages of 252.5 ms and 289.4 ms, respectively). Reaction times to S1 were significantly faster than reaction times to S2 (average means of 261.3 ms and 280.6 ms, respectively). A Scheffè test revealed that reaction time means for the lowest luminance of 0.6 log unit above threshold (293.2 ms average) was significantly longer than those to the higher luminances of 1.0 and 1.4 log units above threshold (264.3 ms and 255.3 ms, respectively) which did not differ significantly from each other. Such an increase in mean reaction time with decreased stimulus intensity has been well documented (Mansfield, 1973; Teichner and Krebs, 1972; Vaughan, Costa and Gilden, 1966). A Scheffé test similarly reveals that the mean reaction time in the binocular condition (255.3 ms) was significantly faster than those in the right-eye and left-eye viewing conditions (277.5 ms and 280.0 ms, respectively) which were not significantly different. Decreased reaction time in binocular observation has often been reported (Gilliland and Haines, 1975; Haines, 1977; Miles, 1936; Minucci and Connors, 1967; Poffenberger, 1912; Teichner and Krebs, 1972).

Five two-way interactions and one three-way interaction were found to be significant in the analysis of variance on mean reaction time. These are all depicted graphically in figures 22 through 27. Figure 22 depicts the significant ($p < 0.005$) interaction between <u>observer and stimulus</u>: Although both observers reacted more rapidly to S1 (the right stimulus) than to S2 (the left stimulus), the difference between observer GEN's reaction times were greater than those of MWH (25.8 ms versus 12.8 ms). The <u>observer by luminance interaction</u> (figure 23, $p < 0.0005$) is similarly explained: The overall reaction time for GEN decreased gradually in steps of 12.7 ms then 12.4 ms as stimulus intensity was increased from 0.6 to 1.0 to 1.4 log units above threshold, while MWH's decreased much more

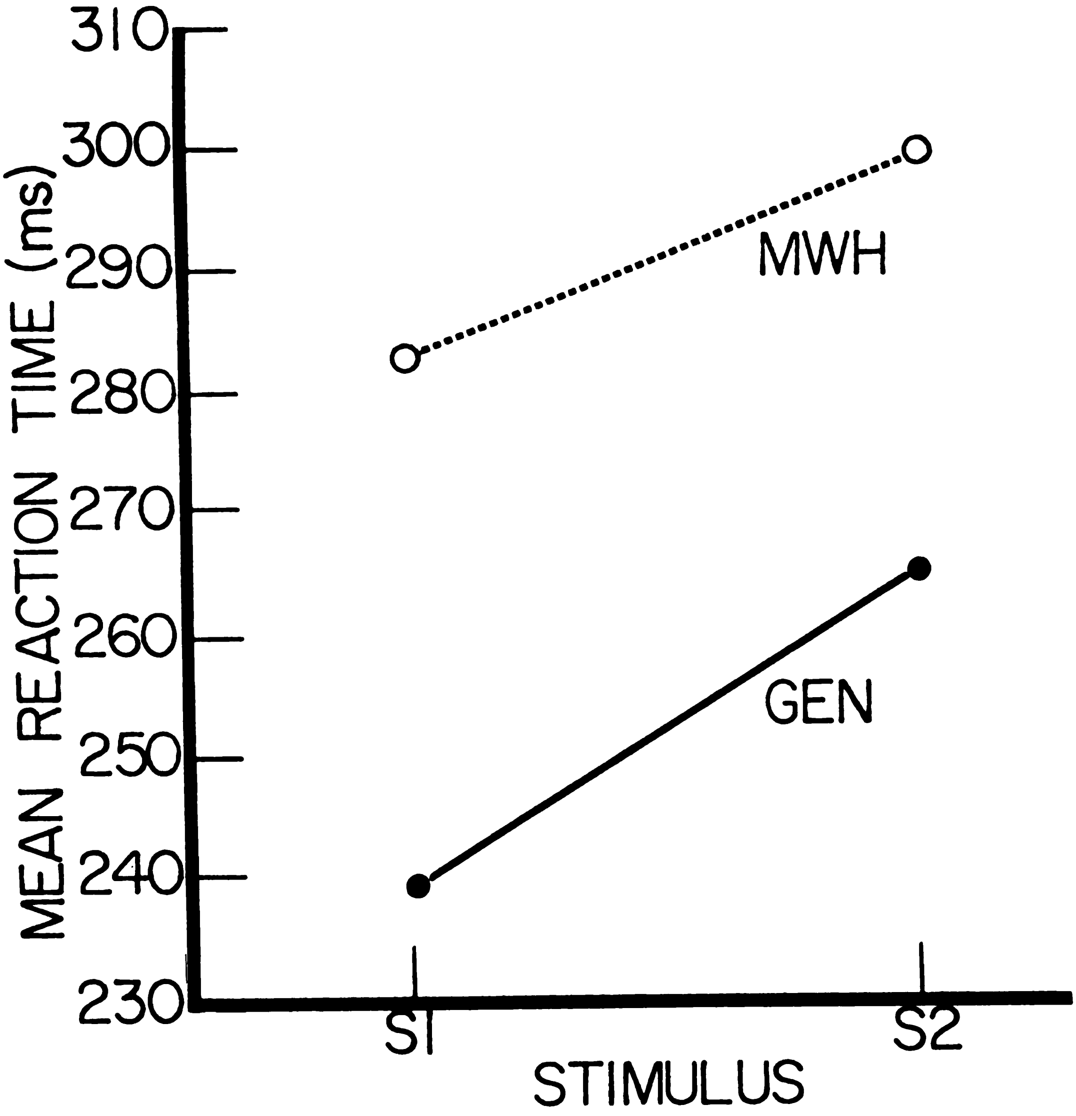

Figure 22. The significant observer by stimulus interaction in the mean reaction time

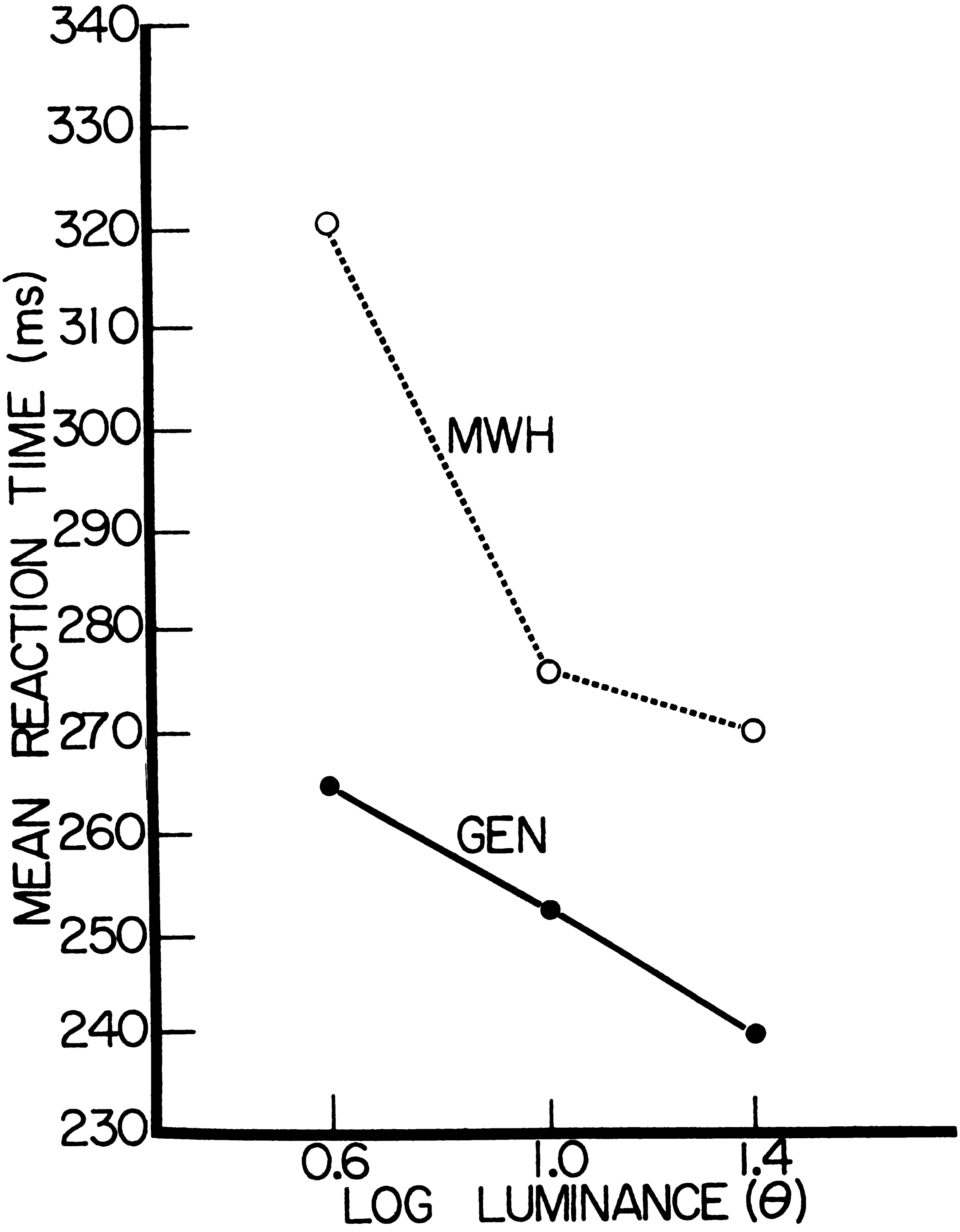

Figure 23. The significant observer by luminance interaction in the mean reaction time

dramatically at first with 45.1 ms then 5.7 ms.

Unlike the anomolous results of these first two interactions, the observer by viewing condition interaction (figure 24, $p < 0.05$) has an immediate interpretation in terms of the effects discussed in the introduction to experiment 2: Although both observers responded in the binocular condition and with about the same amount of reduction in reaction time (26.5 ms for GEN and 20.4 ms for MWH faster than the average monocular reaction time), observer GEN showed a faster reaction time in the right-eye viewing condition (by 7.4 ms) and observer MWH showed a slightly faster reaction time in the left-eye viewing condition (by 2.3 ms). Since GEN was right-eye dominant and MWH was left-eye dominant, these results are in line with the findings of Minucci and Connors (1967) who found that dominant eye reaction times are faster than those with the non-dominant eye. Others, however, have failed to replicate this effect (Gilliland and Haines, 1975).

The significant ($p < 0.05$) intensity by ocular condition interaction (figure 25) indicates a differential rate of decrease in reaction time with increased luminance for the ocular conditions: the reaction time decreased in average steps of 34.2 ms then 9.3 ms for the right and 32.8 ms then 11.2 ms for the left eye as luminance was increased from 0.6 to 1.0 to 1.4 log units above threshold, whereas the binocular reaction time decreased only in steps of 19.6 ms then 6.6 ms.

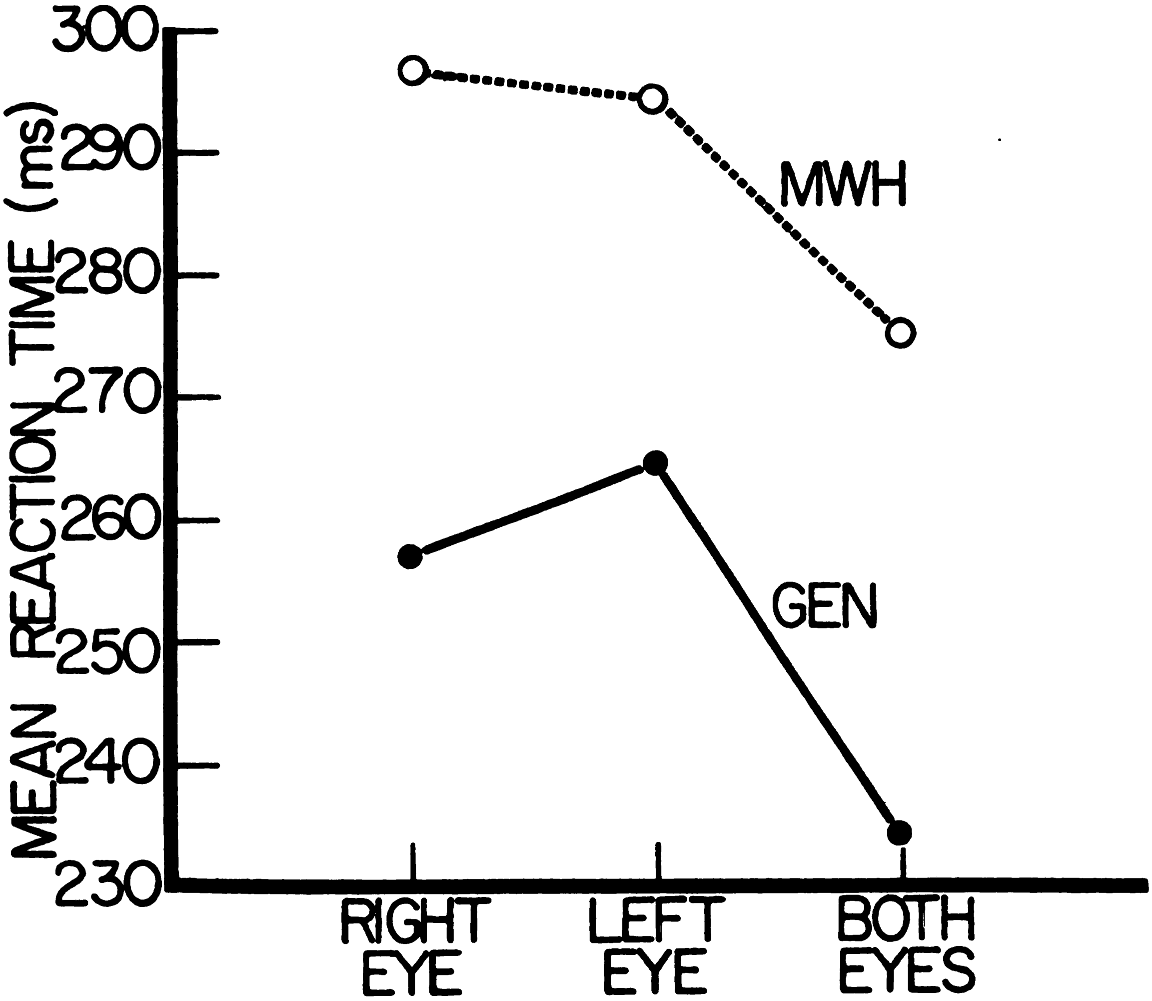

Figure 24. The significant observer by viewing condition interaction in the mean reaction times

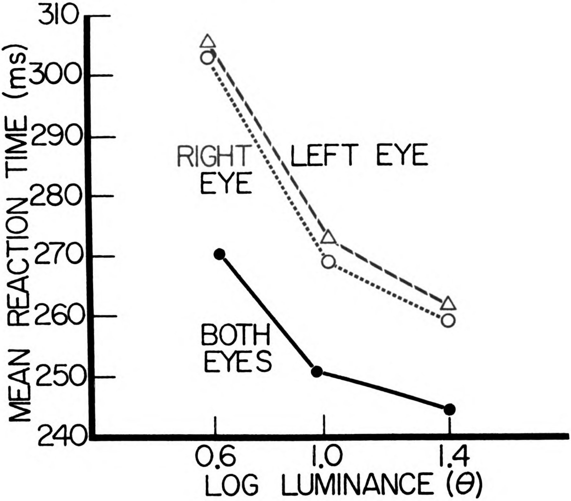

Figure 25. The significant luminance by viewing condition interaction in the mean reaction time

The stimulus by ocular condition interaction was also significant (figure 26, $p < 0.01$) and can be interpreted as a superiority of the nasal hemifield (temporal retina) for producing shorter reaction time. Although the mean reaction time was similar for both stimuli with binocular viewing (20.5 ms shorter than the mean monocular reaction for S1 and 26.4 ms shorter for S2), the right stimulus (S1) produced shorter reaction time for left-eye viewing than for right-eye viewing (266.2 ms versus 270.0 ms) whereas the left stimulus (S2) produced shorter right- than left-eye reaction time (284.9 ms versus 293.9 ms). The magnitude of the effect was about 6.4 ms in favor of shorter reaction time for stimuli imaged on the temporal retina. This effect was not greatly different from the latency difference inferred from the temporal order judgment data of experiment 1. Both results conflict with the generally held notion that visual latency is shorter in the nasal retina (Gilliland and Haines, 1975; Maddess, 1975; Poffenberger, 1912; Rains, 1964). This generalization is not always correct, however; at certain positions along the horizontal meridian stimuli delivered to the temporal retina show reaction time shorter than stimuli delivered to nasal retina (Payne, 1967; Rains, 1963). This position is at a spot in the temporal retina which corresponds to the blind spot in the other eye (about 15° peripheral of the fovea). Many such positions can be found along other than horizontal meridians (Haines, 1977; Payne, 1967). The

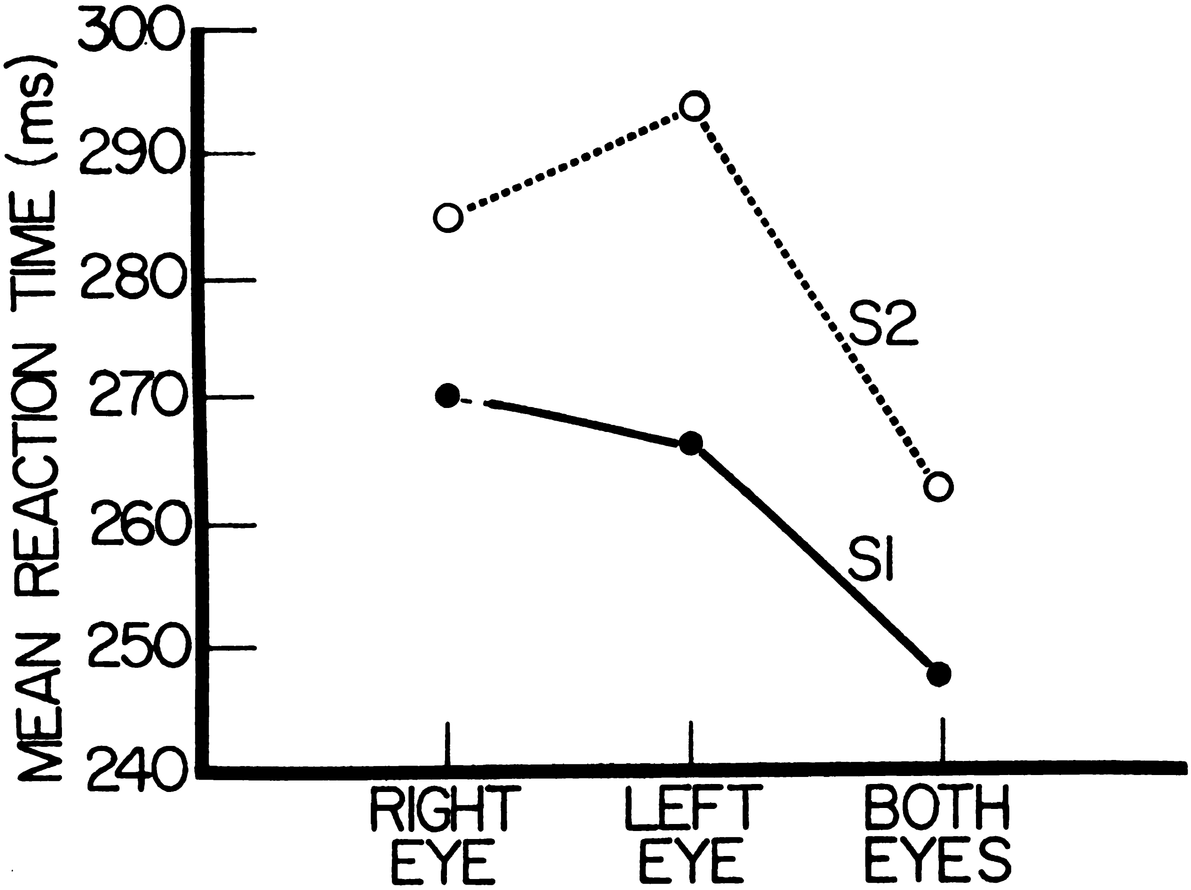

Figure 26. The significant stimulus by viewing condition interaction in the mean reaction times

stimuli used in the present experiments were only $1^{\circ}\ 42'$ peripheral to the fovea. This is a much more central retinal position than was used by investigators who were interested in comparing reaction time to stimuli in the two hemifields (Maddess used stimuli 4° peripheral, but this is the most central position reported). Given the complex relationship between retinal position and reaction time (see Haines, 1977, and Payne, 1967), it would not be especially surprising to find a superiority of the temporal retina at this stimulus separation.

The significant ($p < 0.05$) three-way interaction was the <u>observer by stimulus by viewing condition interaction</u> shown in figure 27. Both observers showed the nasal hemifield superiority just described for the stimulus by viewing condition interaction, but they showed it to a greater extent in their dominant eyes. The right-eye dominant observer GEN showed only a very slight decrease in reaction time the right stimulus (S1) in left-eye over right-eye viewing (less than 1.0 ms), but a large decrease in reaction time to the left stimulus (S2) in right- over left-eye viewing (15.9 ms). The left-eye dominant observer MWH, on the other hand, showed a larger decrease in reaction time to the right stimulus (S1) in left- over right-eye viewing (6.7 ms) but a smaller decrease in reaction time to the left stimulus (S2) in right- over left-eye viewing (2.2 ms). The stimulus by luminance interaction and the luminance by observer by viewing condition interaction also produced

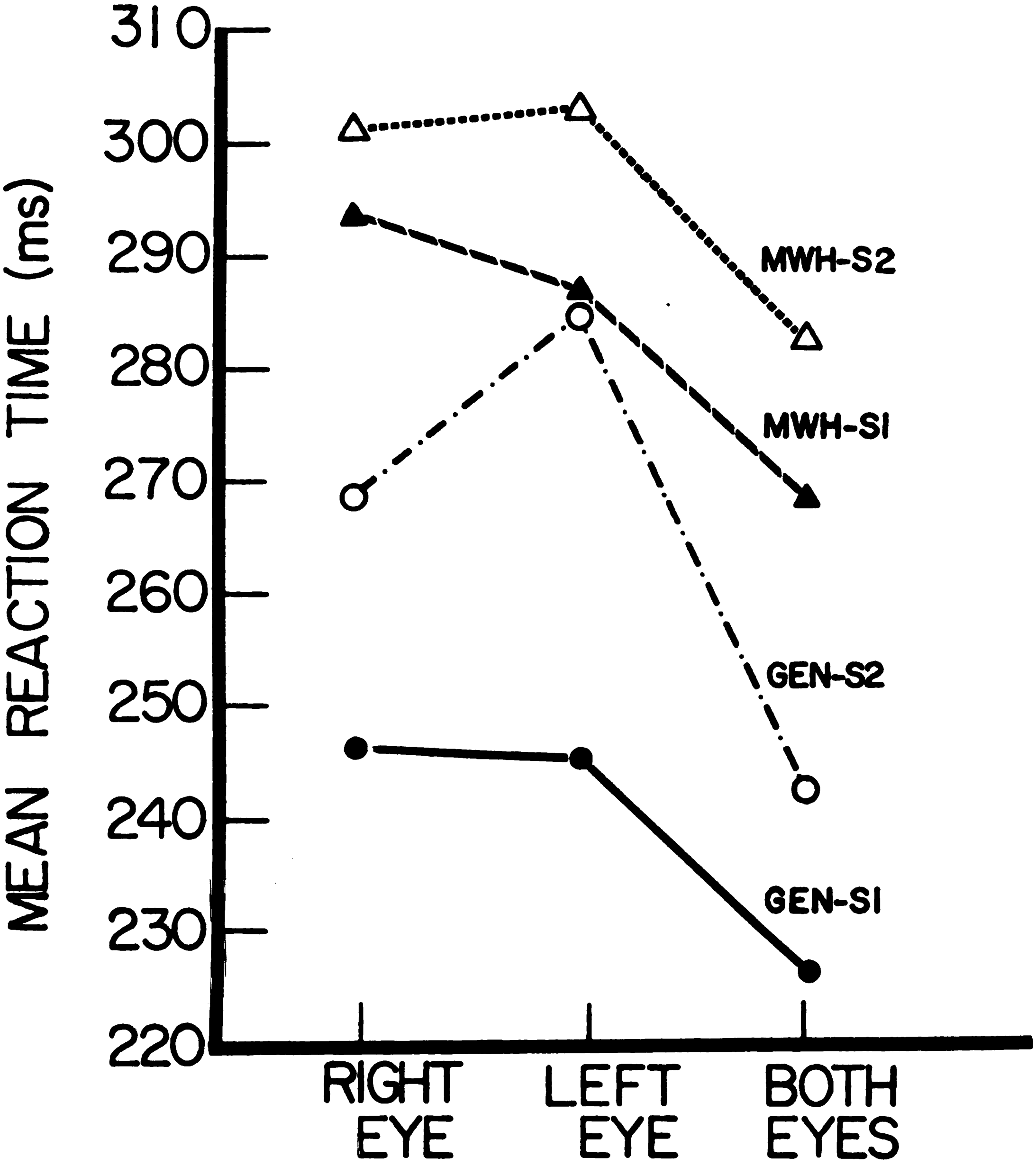

Figure 27. The significant observer by stimulus by viewing condition interaction in the mean reaction times

large F-ratios, but since none reached significance they are not discussed here.

Reaction Time Standard Deviation Results and Discussion

The analysis of variance applied to the reaction time standard deviations (ommitting GEN3) revealed only three significant main effects (see Appendix H). Observer GEN's average standard deviation (47.5 ms) was significantly ($p < 0.001$) smaller than observer MWH's average standard deviation (67.8 ms); the average standard deviation to S1 (54.2 ms) was significantly ($p < 0.05$) smaller than those to S2 (61.0 ms); and the average standard deviations in the lowest luminance level (63.4 ms) were significantly ($p < 0.05$) larger than those in the middle and highest luminances (54.4 ms and 55.0 ms, respectively). There were no significant interactions although the observer by stimulus by viewing condition interaction's F-ratio was fairly large.

The noticeable exception in this analysis of variance is the insignificance of the viewing condition on standard deviations. Table 3 reveals that in nine out of fourteen cases (including GEN3) was the binocular reaction time standard deviation less than both monocular standard deviations. The F-ratio for the viewing condition in the analysis of variance did not quite reach significance ($p < 0.115$). The average reaction time standard deviations for right-and left-eye viewing were 60.0 ms and 59.5 ms while binocular viewing yielded an average reaction time standard deviation of 53.3 ms, so that there was some decrease in the standard deviations for the binocular case. An F-test for independent variances, on the other hand, would indicate that the variance of the combined binocular reaction times is significantly less than that of the combined monocular reaction times ($F_{9888,\ 4712} = 1.2590$, $p < 0.002$). This is a very small F-ratio and reaches significance only because of the enormous degrees of freedom. This raises the question as to how large an F-ratio would have been predicted *a priori* on the basis of the reaction time models presented earlier. These models postulate a reduction in the variability of a sensory-detection component for binocular viewing between 68.2 and 50.0 per cent of monocular variability, so if the motor variability were zero, this should be reflected in F-ratios from 1.47 to 2.00. However, it has already been noted that a large motor variability relative to the sensory-detection variability was present in the reaction time. This motor variability represented about 69 percent of GEN's and 77 per cent of MWH's overall reaction time variance. If the motor component were, say, 73 per cent of the overall variance then the observed F-ratio should lie between 1.09 and 1.27. Thus the observed reduction in reaction time variability with binocular viewing was insufficient to achieve significance in the analysis of variance but was within the expectations of the models.

If the models are correct, then the failure to achieve significance must be attributed to the small magnitude of the effect relative to other sources of variance.

The reaction time results presented so far present some support for the notion that binocular viewing reduces the variability in reaction time relative to monocular viewing. When the reaction variances are used to predict the variances in the temporal order judgment psychometric functions and then to estimate the ratio σ_B/σ_M as was done in table 4, then some additional support for the hypothesis is provided. In every case but one this ratio was less than one. Furthermore, these predicted ratios correlate significantly ($p < 0.05$) with the obtained ratios at 0.782. Thus it can be concluded that although there was not a dramatic decrease in reaction time variability with binocular viewing, the reaction time data did behave similarly to the temporal order judgment data in all respects except in the prediction of the point-of-subjective-simultaneity given that the reaction time has an additional motor variance. This does, then, tend to confirm the theorizing put forth in the pilot experiment and experiment 1.

CONCLUSION

The small binocular reduction in reaction time variability found in experiment 2 as well as the confirmation of the binocular reduction in mean reaction time provides not only converging support for the binocular reduction in arrival latency but also for the low-threshold binocular summation model in particular. Of those models whose predictions were considered, only this model predicted both the slight reduction in reaction time variance and in reaction time mean. Furthermore, this model provided the best accounting of the temporal order judgment effects observed in the pilot experiment and experiment 1.

The implications if the low-threshold binocular summation model is validated are significant. The models which postulate a binocular temporal order judgment resulting from a "higher" processing of the monocular sensory information -- the likelihood ratio model, the average arrival time model and the most discrepant arrival time model -- do not postulate any special binocular system. Rather, the visual system is assumed to be inherently monocular, at least for the perception of onset, and the cyclopean view is synthesized from the monocular views through a rather complex processing. The binocular summation models, on the other hand, do require a special binocular system for the perception of onset. This system receives inputs from both eyes and combines then at some well-defined spatial locus. The high-threshold version postulates two separate visual systems to handle monocular and binocular perception of onset, at least when luminance level is low as it was in these experiments. It assumes that a single monocular input alone is insufficient to excite the binocular system. The binocular system would then be inactive during purely monocular viewing. Cortical neurons with just the property have been described in area 18 of the monkey by Hubel and Wiesel (1970). These binocular "depth" cells do not usually respond to stimulation of either eye separately, but appropriate stimulation of the two eyes together evokes a brisk response. The low-threshold model, on the other hand, requires only a single visual system to handle the perception of onset because the same units would be active in either binocular or monocular viewing.

At the level of the visual cortex of the brain, single neurons do receive excitatory inputs from both eyes. Hubel and Wiesel demonstrated this for the first time in 1959 by recording from single neurons in the visual cortex of the cat. For almost every cell studied, two areas could be defined where light stimuli evoked a response, one associated with each eye. The most vigorous response was produced by simultaneous stimulation of the two eyes. The proportion of binocularly activated neurons found has increased as the technique of presenting stimuli has become more and more refined. Recently, Bishop, Coombs and Henry (1973) showed that all the cells in the striate cortex of the cat receive an excitatory input from both eyes. Several lines of psychophysical investigation lead to the conclusion that the human brain also contains such binocular cells (eg., Blake and Levinson, 1977). Thus the striate cortex can be regarded as a "cyclopean retina" and a possible neurophysiological locus for the operation of the binocular summation models.

The low-threshold model, if it is to represent optimum binocular performance, places some limitations on the observer's ability to use the onset information present in the signals arising from the two eyes. Most investigations of temporal order judgment assume, as has the present one, that for the time relation between two signals to be judged, their representations must be brought together somewhere in the brain -- at a "central timing mechanism" (eg., Sternberg and Knoll, 1973). Colavita (1977) has implicated the insular-temporal cortex, at least in the cat, as "the critical brain region for discrimination of changes in temporal patterns of stimuli." He further provides evidence that this region receives multimodal sensory input, including visual, via cortico-cortical connections. Whatever the locus of such a central timing mechanism, it is clear that the low-threshold model if it is to describe optimum performance demands that signals from the two eyes be combined and so confounded prior to this region.

The cortical neurons of the cat tend to meet these requirements. Certainly any information passed on to central timing mechanism from the visual cortex of the cat via cortico-cortical connections must have left- and right-eye information combined. The situation is not so clear in the monkey whose visual system is assumed to be most similar to man's. In monkey striate cortex (area 17), cells activated by both eyes are fewer than in the cat, and even those that do receive binocular input are typically dominated by one eye or the other (Hubel and Wiesel, 1970). So, either only those cells which receive a nearly balanced input are passed on to the central timing mechanism, or the visual cortex must not be the final point where monocular signals combine prior to the temporal order judgment, or the low-threshold summation model does not represent optimum performance. If the monocular signals were available for temporal order judgment, then superior performance, say, at the level of the average arrival time model, could have been achieved.

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APPENDICES

APPENDIX A

THE LIKELIHOOD RATIO MODEL

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Suppose that the density function $f_1(s_1)$ defines the conditional point probability that a given sensory experience, s_1, on some hypothetical sensory continuum arose given that S1 came first and that $f_1(s_1)$ is normal with mean, μ_{S1}, and variance, σ^2. Similarly suppose that $f_2(s_2)$ gives the conditional probability of the s_2 experience when S2 preceded S1, and that $f_2(s_2)$ is normal with mean, μ_{S2}, and variance equal to the variance of $f_1(s_1)$. Let us further assume that s_1 and s_2 are independent and the observer's criterion is 1.0. Then in the monocular case, the observer judges that "S1 came first" whenever on a given trial $\ell_L(s_L)$ is greater than 1.0, where $\ell_L(s_L)$ is defined by equation 7 of the text. By supplying the appropriate forms of the density functions assumed above, equation 7 can be shown to be equivalent to

$$\exp\left\{-\tfrac{1}{2}\left(\frac{s - \mu_{S1}}{\sigma}\right)^2 + \tfrac{1}{2}\left(\frac{s - \mu_{S2}}{\sigma}\right)^2\right\} > 1.0 \qquad (59)$$

which can be further reduced by the the natural logarithm of both sides, by assuming that $\mu_{S1} - \mu_{S2}$ is positive and by doing some algebraic manipulations. Then the observer will say that S1 preceded S2 whenever

$$s > \tfrac{1}{2}\mu_{S1} + \tfrac{1}{2}\mu_{S2} \qquad (60)$$

Since a "hit" is defined as reporting that S1 came first when S1 actually did come first, the probability of a hit in the left monocular case is given by

$$P_L(\text{"hit"}) = 1 - F_1(\tfrac{1}{2}\mu_{S1} + \tfrac{1}{2}\mu_{S2}) \qquad (61)$$

where $F_1(s)$ is the cumulative distribution function of $f_1(s)$. By making the change of variable $z = (s - \mu_{S1})/s$ in the integral form of equation 61, it can be shown that equation 61 is equivalent to

$$P_L(\text{"hit"}) = 1 - F_o(-\tfrac{1}{2}\, d'_L) = F_o(\tfrac{1}{2}\, d'_L) \qquad (62)$$

where F_o is the cumulative standard normal distribution function and $d'_L = (\mu_{S1} - \mu_{S2})/\sigma$. If it is assumed that the sensitivity of the right monocular system, d'_R, is identical to left sensitivity, d'_L, then the probability of a hit in the right system, $P_R(\text{"hit"})$, is equal to the probability of a hit in the left system, $P_L(\text{"hit"})$, as given by equation 62.

Now since the psychometric function for temporal order judgment with the left eye, $F_L(d)$, is a nondecreasing function bounded by 0.0 and 1.0, it makes sense to approximate it with an ogive function which can be characterized by the two parameters μ_L and σ_L. $F_L(d)$ is then approxiated by

$$F_L(d) = \int_{-\infty}^{(d - \mu_L)/\sigma_L} f_o(x)dx \tag{63}$$

where $f_o(x)$ is the standard normal distribution and x is a variable of integration. Any particular value of $F(d)$ is equal to the hit rate for the stimulus onset asynchrony of value d. An analogous equation to 63 exists for hits in the right-eye viewing condition, $F_R(d)$, differing only in that the subscript "R" replaces the subscript "L."

Equating equation 63 with the integral form of equation 62, it can be shown that the two indefinite integrals can be equated if and only if

$$\tfrac{1}{2}\, d'_L = (d - \mu_L)/\sigma_L \tag{64}$$

for which again there exists an analogous equation for the right system differing only in the subscripts.

In the binocular temporal order judgment, this model assumes that the judgment is made according to the likelihood ratio, $\mathcal{L}_B(s_1, s_2)$, given by equation 8 of the text, which, if the two monocular systems are independent, can be reduced to

$$\mathcal{l}_B(s_1, s_2) = \mathcal{l}_L(s_1)\mathcal{l}_R(s_2) \tag{65}$$

the product of the two monocular likelihood ratios. If, as was done for the monocular cases, the appropriate forms of the density functions are supplied in equation 65, then the binocular observer will say S1 came first when

$$\exp\left\{ \begin{array}{l} -\frac{1}{2}\left(\frac{s_1 - S1}{\sigma}\right)^2 - \frac{1}{2}\left(\frac{s_2 - S1}{\sigma}\right)^2 \\ +\frac{1}{2}\left(\frac{s_1 - S2}{\sigma}\right)^2 + \frac{1}{2}\left(\frac{s_2 - S2}{\sigma}\right)^2 \end{array} \right\} > 1.0 \tag{66}$$

Equation 66 can be reduced (be taking the natural logarithm of both sides, assuming that $\mu_{S1} - \mu_{S2}$ is greater than zero, and doing some algebraic manipulation) so that the observer will say that S1 came first when

$$s_2 > \mu_{S1} + \mu_{S2} - s_1 \tag{67}$$

The probability of a hit in the binocular case, P_B("hit"), is then given by

$$P_B(\text{"hit"}) = 1 - \int_{-\infty}^{+\infty} F_1(\mu_{S1} + \mu_{S2} - s_2)f_1(s_1)ds_1 \tag{68}$$

By making the transformation $z_1 = (s_1 - \mu_{S1})/\sigma$ and $z_2 = (s_2 - \mu_{S1})/\sigma$, it can be shown that equation 68 is equivalent to

$$P_B(\text{"hit"}) = 1 - \int_{-\infty}^{+\infty} F_o(-d'_L - z_1) f_o(z_1) dz_1$$

$$= \int_{-\infty}^{+\infty} F_o(d'_L + z_1) f_o(z_1) dz_1 \qquad (69)$$

where z_1 is a variable of integration. The binocular temporal order judgment psychometric function can represented as it was for the monocular case by an ogive identical to equation 63 except with the subscript "B" replacing the subscript "L." If equation 64 is solved for d and substituted into equation 63, then the psychometric function for binocular temporal order judgment, $F_B(d)$, is given by

$$F_B(d) = \int_{-\infty}^{v(d')} f_o(x) dx \qquad (70)$$

where $v(d') = (\frac{1}{2}\sigma_L d'_L + \mu_L - \mu_B)/\sigma_B$ and where x is a variable of integration. If the right half of equation 69 is equated with the right half of 70, and if both sides of the resulting equation are differentiated with respect to d', then the following results

$$\frac{\sigma_L}{2\sigma_B} f_o(v(d')) = \int_{-\infty}^{+\infty} f_o(d' - x) f_o(x) dx \qquad (71)$$

by a series of manipulations, the right side of equation 71 can be simplified to

$$\frac{\sigma_L}{2\sigma_B} f_o(v(d')) = \frac{1}{\sqrt{2}} f_o\left(\frac{d'}{\sqrt{2}}\right) \qquad (72)$$

which can be true if and only if

$$\mu_B = \mu_L = \mu_R \tag{73}$$

and

$$\sigma_B = \frac{\sigma_L}{\sqrt{2}} = \frac{\sigma_R}{\sqrt{2}} \tag{74}$$

which then relates the parameters of the binocular psychometric function to those of the monocular psychometric functions. Equations 73 and 74 combined with equation 64 and its counterpart for the right and binocular sensitivities can be shown to satisfy equation 9 of the text. These same results are also valid when $\mu_{S1} - \mu_{S2}$ is negative and in the limit as $\mu_{S1} - \mu_{S2}$ approaches zero, although the derivation is slightly modified for these cases.

The assumption that $d'_L = d'_R$, $\mu_L = \mu_R$, and $\sigma_L = \sigma_R$ was reasonably well approximated by the data presented in the pilot experiment and in experiment 1. For the general binocular case, the hit rate is given by

$$P_B(\text{"hit"}) = \int_{-\infty}^{+\infty} F_o\left[\frac{d'_R}{d'_L}\,x + \frac{(d'_R)^2}{2d'_L} + \frac{d'_L}{2}\right] f_o(x)\,dx \tag{75}$$

where x is a variable of integration. This equation is analogous to equation 69 and it may be reduced in a similar manner. This solution is, however, sufficiently complex that it will not be attempted here.

APPENDIX B

THE AVERAGE ARRIVAL TIME MODEL

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THE AVERAGE ARRIVAL TIME MODEL

The decision rule for the average arrival time model, G_{aa}, may be written in analogy to equation 12 as

$$G_{aa}(x_{L,1}, x_{L,2}, x_{R,1}, x_{R,2}) = \begin{cases} 1.0 & \text{if } \bar{x}_1 - \bar{x}_2 < 0.0 \\ 0.0 & \text{if } \bar{x}_1 - \bar{x}_2 > 0.0 \end{cases} \tag{76}$$

where $\bar{x}_1$ is the average of the signals evoked by S1 and is equal to $\omega x_{R,1} + (1 - \omega)\, x_{L,1}$; $\bar{x}_2$ is the average arrival of the signals evoked by S2 and is equal to $\omega x_{R,2} + (1 - \omega)\, x_{L,2}$; and where ω is positive and less than one. The parameter, ω, represents a biasing factor which allows for the preferential weighting of one of the eyes. It would not be expected to vary dramatically for a given observer. If there is no eye preference, then $\omega = 0.5$. Equation 76 differs from equation 12 in that the weighted average of the arrival times of signals elicited by S1 and S2 replace the arrival times of the single signals in the monocular case.

In a similar way, equation 77 which describes the psychometric function $F_B(d)$ relating the binocular temporal order judgment for a value of the stimulus onset asynchrony, d, is given in analogy to equation 13 as

$$F_B(d) = P\left\{\left[\omega r_1 + (1 - \omega)l_1\right] - \left[\omega r_2 + (1 - \omega)l_2\right] \leq d\right\} \qquad (77)$$

A final analogy, this time to equations 14 and 15 is given to described the parameters of $F_B(d)$ in equations 78 and 79.

$$\mu_B = E\left[\omega r_1 + (1 - \omega)l_1\right] - E\left[\omega r_2 + (1 - \omega)l_2\right] \qquad (78)$$

$$\sigma_B^2 = \mathrm{VAR}\left[\omega r_1 + (1 - \omega)l_1\right] + \mathrm{VAR}\left[\omega r_2 + (1 - \omega)l_2\right] \qquad (79)$$

If r_1, r_2, l_1, and l_2 are independent normally distributed random variables with means $\mu_{R,1}$, $\mu_{R,2}$, $\mu_{L,1}$, and $\mu_{L,2}$, respectively, and variances $\sigma_{R,1}^2$, $\sigma_{R,2}^2$, $\sigma_{L,1}^2$, and $\sigma_{L,2}^2$, then equations 78 and 79 may be reduced to

$$\begin{aligned}\mu_B &= \omega(\mu_{R,1} - \mu_{R,2}) + (1 - \omega)(\mu_{L,1} - \mu_{L,2}) \\ &= \omega\mu_R + (1 - \omega)\mu_L\end{aligned} \qquad (80)$$

$$\sigma_B^2 = \omega^2(\sigma_{R,1}^2 + \sigma_{R,2}^2) + (1 - \omega)^2(\sigma_{L,1}^2 + \sigma_{L,2}^2)$$

$$= \omega^2\sigma_R^2 + (1 - \omega)^2\sigma_L^2 \tag{81}$$

In the text (equations 20 and 21) is given the special case of these equations where ω is equal to 0.5 and where σ_R^2 and σ_L^2 are the same.

A model, conceptually very similar to the average arrival time model, could be called the <u>average arrival difference model</u>. Here it is assumed that the two arrival time differences, $x_{L,1} - x_{L,2}$ and $x_{R,1} - x_{R,2}$ which will be denoted x_L' and x_R', are averaged in the binocular decision rule given by

$$G_{aad}(x_L', x_R') = \begin{cases} 1.0 \text{ if } \omega x_R' + (1 - \omega)x_L' < 0.0 \\ 0.0 \text{ if } \omega x_R' + (1 - \omega)x_L' > 0.0 \end{cases} \tag{82}$$

This model postulates a different sequence of processing than the average arrival time model in that the differencing precedes the averaging, but it is mathematically indistinguishable from it as can be shown by simple rearranging the terms in equation 82 until equation 76 emerges.

APPENDIX C

THE MOST DISCREPANT ARRIVAL TIME MODEL

APPENDIX C

THE MOST DISCREPANT ARRIVAL TIME MODEL

This decision rule may be expressed by

$$G_{mda}(x_{L,1}, x_{L,2}, x_{R,1}, x_{R,2})$$

$$= \begin{cases} 1.0 \text{ if} & \begin{array}{l} x_{R,1} - x_{R,2} < x_{L,1} - x_{L,2} < 0 \\ x_{R,1} - x_{R,2} < -(x_{L,1} - x_{L,2}) < 0 \\ x_{L,1} - x_{L,2} < x_{R,1} - x_{R,2} < 0 \\ x_{L,1} - x_{L,2} < -(x_{R,1} - x_{R,2}) < 0 \end{array} \\ 0.0 \text{ if} & \begin{array}{l} 0 < x_{R,1} - x_{R,2} < x_{L,1} - x_{L,2} \\ 0 < -(x_{R,1} - x_{R,2}) < x_{L,1} - x_{L,2} \\ 0 < x_{L,1} - x_{L,2} < x_{R,1} - x_{R,2} \\ 0 < -(x_{L,1} - x_{L,2}) < x_{R,1} - x_{R,2} \end{array} \end{cases} \tag{83}$$

Since the inequalities of equation 83 are disjoint, the predicted binocular psychometric function is given by

$$\begin{aligned} F_B(d) = {} & P\left[l_1 - l_2 < r_1 - r_2 < d\right] \\ & + P\left[l_1 - l_2 < -(r_1 - r_2) < d\right] \\ & + P\left[r_1 - r_2 < l_1 - l_2 < d\right] \\ & + P\left[r_1 - r_2 < -(l_1 - l_2) < d\right] \end{aligned} \tag{84}$$

which may be simplified to equation 85 if it is assumed as it has been with the other models that the arrival latency distributions are normal. After making several transformations of the variables of integration and after differentiating with respect to d, the psychometric function may be described by the underlying <u>density</u> function

$$F_B(d) = \frac{1}{\pi\,\sigma_R} \int_{-\infty}^{+\infty} \exp\left\{ -\tfrac{1}{2}\left[\frac{\sigma_L}{\sigma_R}\, y + \frac{2d - \mu_L - \mu_R}{\sigma_R} \right]^2 - \tfrac{1}{2}y^2 \right\} dy \tag{85}$$

where y is a variable of integration. Finally if we consider the special case where $\sigma_L = \sigma_R$, then equation 85 reduces to a recognizable normal density function characterized by equations 20 and 21 of the text.

Very little detail has been devoted to the derivation of the most discrepant arrival time model because it can, in fact, be shown to be mathematically equivalent to the average arrival time mode. Examine the defining conditions of the decision rule (equation 83) and notice that the same decision would have been made to the sign of the following: $0.5(x_{L,1} - x_{L,2}) + 0.5(x_{R,1} - x_{R,2})$. The sign of the larger arrival time difference would dominate and the final decision would be based only on the sign of this larger difference. This is, of course, the defining condition of the average arrival time decision rule (see equation 76). This equivalence of the two models only holds when ω is 0.5 and when the decision rule is a deterministic one.

APPENDIX D

THE LOW-THRESHOLD BINOCULAR SUMMATION MODEL

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THE LOW-THRESHOLD BINOCULAR SUMMATION MODEL

The total arrival latency at the decision locus will be assumed to be composed of the sum of two independent latencies: (1) the latency, a, from the onset of the stimulus to the arrival of the signal at the binocular combination point, and (2) the latency, b, from the arrival of the signal at the combination point to the arrival of the combined signal at the decision mechanism. One of the monocular latencies, l_1, for example, may be expressed as

$$l_1 = a_{l_1} + b_1 \tag{86}$$

with a similar expression for each of the three other monocular latencies.

These latencies are used by the general temporal order judgment model (Sternberg and Knoll, 1973) in equation 13 to predict the psychometric function. In the left monocular experiment, the performance manifested by the psychometric function, $F_L(d)$, can be characterized by combining equations 14, 15 and 86. This results in the point-of-subjective-simultaneity, μ_L, given by

$$\mu_L = E(a_{l_1}) - E(a_{l_2}) + E(b_1) - E(b_2) \tag{87}$$

where E is the expected value operator. The variance underlying $F_L(d)$ is similarly given by

$$\sigma_L{}^2 = VAR(a_{l_1}) + VAR(a_{l_2}) + VAR(b_1) + VAR(b_2) \tag{88}$$

where VAR is the variance operator. An analogous set of equations could be demonstrated for the right monocular system.

In the binocular temporal order judgment, the total binocular latency from the onset of S1 to the arrival at the decision locus would be given by

$$Y_1 = \min\,(a_{l_1},\, a_{r_1}) + b_1 \tag{89}$$

where min is the minimum operator. In equation 89 it specifies the smaller of a_{l_1} and a_{r_1}. A similar equation would apply to the total binocular latency for the signal evoked by S2, namely Y_2. Binocular performance could then be characterized in analogy to equations 87 and 88 by the mean, μ_B, given by

$$\mu_B = E\left[\min(a_{l_1},\, a_{r_1})\right] - E\left[\min(a_{l_2},\, a_{r_2})\right] + E(b_1) - E(b_2) \tag{90}$$

and variance, σ_B^2, given by

$$\sigma_B^2 = \mathrm{VAR}\left[\min(a_{l_1}, a_{r_1})\right] + \mathrm{VAR}\left[\min(a_{l_2}, a_{r_2})\right] + \mathrm{VAR}(b_1) + \mathrm{VAR}(b_2) \qquad (91)$$

If the latencies b_1 and b_2 are identical and constant, then the terms in which they appear in equations 90 and 91 vanish. In this case, the minimum latency, $\min(a_{l_1}, a_{r_1})$ hereafter denoted a_1, has the probability density function, g_1, which is given by

$$g_1(a_1) = \left[1 - F_{L,1}(a_1)\right] f_{R,1}(a_1) + \left[1 - F_{R,1}(a_1)\right] f_{L,1}(a_1) \qquad (92)$$

where the density functions $f_{L,1}$ and $f_{R,1}$ are as defined previously except that they exclude the post-combination latency, b. The capital "F" designates, as usual, the cumulative distribution function. Similarly, the minimum latency, $\min(a_{l_2}, a_{r_2})$ hereafter denoted a_2, has density g_2 given by

$$g_2(a_2) = \left[1 - F_{L,2}(a_2)\right] f_{R,2}(a_2) + \left[1 - F_{R,2}(a_2)\right] f_{L,2}(a_2) \qquad (93)$$

Since the density functions within equation 92 were assumed to be normal, the nth noncentral moment of a_1 is given by

$$E_1(a_1{}^n) = E_{R,1}(a_1{}^n) + E_{L,1}(a_1{}^n) - \int_{-\infty}^{\infty}\left(\sigma_{R,1}x + \mu_{R,1}\right)^n F_o\left\{\frac{\sigma_{R,1}}{\sigma_{L,1}}\,x - \frac{\mu_{L,1} - \mu_{R,1}}{\sigma_{L,1}}\right\}\cdot f_o(x)\,dx - \int_{-\infty}^{+\infty}\left(\sigma_{L,1}x + \mu_{L,1}\right)^n F_o\left\{\frac{\sigma_{L,1}}{\sigma_{R,1}}\,x + \frac{\mu_{L,1} - \mu_{R,1}}{\sigma_{R,1}}\right\}\cdot f_o(x)\,dx \quad (94)$$

where the subscripts on the expectation operator indicate which density function the expectation is taken over. These noncentral moments have been examined numerically for the special case where $\sigma_{L,1} = \sigma_{R,1} = \sigma_1$. In this case the final two integrals in equation 94 depend only on the value $d_1' = (\mu_{L,1} - \mu_{R,1})/\sigma_1$. The numerical evaluation of these integrals revealed several important relationships (at least for n=0,1,2,3). First,

$$\int_{-\infty}^{+\infty} x^n F_o(x - d_1')f_o(x)\,dx = \begin{cases} \displaystyle\int_{-\infty}^{+\infty} x^n F_o(x + d_1')f_o(x)\,dx & \text{if n is odd} \\ 1 - \displaystyle\int_{-\infty}^{+\infty} x^n F_o(x + d_1')f_o(x)\,dx & \text{if n is even} \end{cases} \quad (95)$$

All of the odd values of n yield "bell-shaped," symmetric (about the origin) functions. For n = 1, this function is given by

$$\int_{-\infty}^{+\infty} xF_o(x - d_1')f_o(x)dx = \frac{1}{2\sqrt{\pi}} \cdot e^{-d_1'^2} \tag{96}$$

that is, one-half of a normal density function with mean zero and variance one-half. These results (equations 95 and 96) appear to be exact within in the limits of error for the numerical method. No mathematical proof is offered here to substantiate this claim, however. A third important relationship can be demonstrated for the case where n = 0. Here,

$$\begin{aligned}
&\frac{d}{dd_1'}\int_{-\infty}^{+\infty} F_o(x + d_1')f_o(x)dx \\
&\qquad = \int_{-\infty}^{+\infty} f_o(x + d_1')f_o(x)dx \\
&\qquad = \frac{1}{2\sqrt{\pi}}\cdot\exp\left[-\tfrac{1}{4}d_1'^2\right]\cdot\int_{-\infty}^{+\infty}\frac{1}{\sqrt{\pi}}\cdot\exp -\tfrac{1}{2}\left[\frac{x+\frac{1}{2}d_1'}{\frac{1}{2}\sqrt{2}}\right]^2 dx \\
&\qquad = \frac{1}{2\sqrt{\pi}}\cdot\exp\left[-\tfrac{1}{4}d_1'^2\right]
\end{aligned} \tag{97a}$$

that is, the integral of the first line must be a cumulative distribution function with mean zero and variance two. In other words,

$$\int_{-\infty}^{+\infty} F_o(x + d_1'^2)f_o(x)dx = F_o(d_1'/\sqrt{2}) \tag{97b}$$

Applying these relationships (equations 95, 96 and 97b) to equation 94 leads to the characterization of a_1 with the mean and variance

$$\begin{aligned} E(a_1) = {} & \left[1 - F_o(d'_1/\sqrt{2})\right]\cdot\mu_{L,1} + F_o(d'_1/\sqrt{2})\cdot\mu_{R,1} \\ & -\left[\frac{1}{\sqrt{\pi}}\, e^{-d'^{2}_1}\right]\cdot\sigma_1 \end{aligned} \tag{98a}$$

$$\begin{aligned} \mathrm{VAR}(a_1) &= E(a_1{}^2) - E(a_1)^2 \\ &= \sigma_1{}^2\cdot\Bigg\{1 - \left(\frac{1}{\sqrt{\pi}}\, e^{-d'^{2}_1}\right)^2 \\ &\qquad + F_o(d'_1/\sqrt{2})\cdot\left[1 - F_o(d'_1/\sqrt{2})\right]\cdot d'^{2}_1 \\ &\qquad + \left[\frac{1}{\sqrt{\pi}}\, e^{-d'^{2}_1}\right]\cdot\left[1 - 2F_o(d'_1/\sqrt{2})\right]\cdot d'_1\Bigg\} \end{aligned} \tag{98b}$$

A similar line of argument leads to the characterization of a_2 with mean and variance

$$\begin{aligned} E(a_2) = {} & \left[1 - F_o(d'_2/\sqrt{2})\right]\cdot\mu_{L,2} + F_o(d'_2/\sqrt{2})\cdot\mu_{R,2} \\ & -\left[\frac{1}{\sqrt{\pi}}\, e^{-d'^{2}_2}\right]\cdot\sigma_2 \end{aligned} \tag{99a}$$

$$\begin{aligned} \mathrm{VAR}(a_2) = {} & \sigma_2{}^2\cdot\Bigg\{1 - \left[\frac{1}{\sqrt{\pi}}\, e^{-d'^{2}_2}\right]^2 \\ &\qquad + F_o(d'_2/\sqrt{2})\cdot\left[1 - F_o(d'_2/\sqrt{2})\right]\cdot d'^{2}_2 \\ &\qquad + \left[\frac{1}{\sqrt{\pi}}\, e^{-d'^{2}_2}\right]\cdot\left[1 - 2F_o(d'_2/\sqrt{2})\right]\cdot d'_2\Bigg\} \end{aligned} \tag{99b}$$

Equations 98a, 98b, 99a and 99b could be applied to equations 90 and 92 to characterize the binocular temporal order judgment performance. These equations are omitted here because they follow directly and because they cannot be related compactly.

To simplify these equations, it may be assumed that the performances by the two monocular systems are identical, that is, that $f_{L,1}(a_1) = f_{R,1}(a_1)$ and that $f_{L,2}(a_2) = f_{R,2}(a_2)$. In this case $d_1' = d_2' = 0.0$. Then the mean and variance of the first arrival of the S1 evoked signal, a_1, and of the S2 evoked signal, a_2, can be expressed as

$$E(a_1) = \mu_{L,1} - 0.56419\,\sigma_{L,1} \qquad (100)$$

$$E(a_2) = \mu_{L,2} - 0.56419\,\sigma_{L,2} \qquad (101)$$

and

$$VAR(a_1) = 0.68169\ \sigma_{L,1}^{2} \qquad (102)$$

$$VAR(a_2) = 0.68169\ \sigma_{L,2}^{2} \qquad (103)$$

and the binocular temporal order judgment can be characterized as it was in the text in equations 22 and 23 where the constants are $\sigma_0^2 = 0.31831\ VAR(b_1) + VAR(b_2)$ and $\mu_0 = E(b_1) - E(b_2)$ and were assumed to be negligible.

APPENDIX E

THE HIGH-THRESHOLD BINOCULAR SUMMATION MODEL

APPENDIX E

THE HIGH-THRESHOLD BINOCULAR SUMMATION MODEL

In the binocular temporal order judgment, the total binocular latency, Y_1, for the S1 evoked signal from the onset of S1 to the arrival of the combined response at the decision center would be given by

$$Y_1 = \max\,(a_{l_1}, a_{r_1}) + b_1 \qquad (104)$$

where max is the maxima operator, that is, $\max(a_{l_r}, a_{r_1})$ is equal to the larger of a_{l_1} and a_{r_1}. Compare this to equation 89 for the low-threshold version of this model. A similar equation would hold for the total binocular latency for the signal evoked by S2. This leads to a binocular performance characterized by the mean, μ_B, given by

$$\mu_B = E\left[\max(a_{l_1}, a_{r_1})\right] - E\left[\max(a_{l_2}, a_{r_2})\right] + \mu_o \qquad (105)$$

and variance $\sigma_B{}^2$ given by

$$\sigma_B{}^2 = \mathrm{VAR}\left[\max(a_{l_1}, a_{r_1})\right] + \mathrm{VAR}\left[\max(a_{l_2}, a_{r_2})\right] + \sigma_o{}^2 \qquad (106)$$

where $_{o}$ and σ_o^2 are as defined at the end of Appendix D in regard to equations 22 and 23. The latencies, $\max(a_{l_1}, a_{r_1})$ and $\max(a_{l_2}, a_{r_2})$ hereafter referred to as a_1 and a_2, have probability density functions g_1 and g_2, respectively, given by

$$g_1(a_1) = F_{L,1}(a_1)f_{R,1}(a_1) + F_{R,1}(a_1)f_{L,1}(a_1) \tag{107}$$

and

$$g_2(a_2) = F_{L,2}(a_2)f_{R,2}(a_2) + F_{R,2}(a_2)f_{L,2}(a_2) \tag{108}$$

where the terms are all defined as they were in regard to equations 92 and 93.

The means and variances of a_1 and a_2 for the special cases where $\sigma_{L,1} = \sigma_{R,1} = \sigma_1$ and where $\sigma_{L,2} = \sigma_{R,2} = \sigma_2$ may be calculated using the relations put forth in Appendix D (equations 95 through 97b). In these cases,

$$E(a_1) = F_o(d_1'/\sqrt{2})\cdot\mu_{L,1} + \left[1 - F_o(d_1'/\sqrt{2})\right]\cdot\mu_{R,1} + \frac{1}{\sqrt{\pi}}\cdot e^{-d_1'^2}\cdot\sigma_1 \tag{109}$$

$$E(a_2) = F_o(d_2'/\sqrt{2})\cdot\mu_{R,2} + \left[1 - F_o(d_2'/\sqrt{2})\right]\cdot\mu_{R,2} + \frac{1}{\sqrt{\pi}}\cdot e^{-d_2'^2}\cdot\sigma_2 \tag{110}$$

The variances, $VAR(a_1)$ and $VAR(a_2)$, are identical to those given for the low-threshold version of this model (see Appendix D, equations 98b and 99b). Equations 98b, 99b, 109 and 110 may be applied to equations 105 and 106 to yield characterizations of the binocular performance. These are omitted here to conserve space.

These may be simplified if, as was done for the low-threshold model, it is assumed that the performance by the two monocular systems is identical, that is, that $f_{L,1}(a_1) = f_{R,1}(a_1)$ and that $f_{L,2}(a_2) = f_{R,2}(a_2)$. Then the means and variances of the last arrivals of the S1 evoked signals, a_1, and of the last arrivals of the S2 evoked signals, a_2, are equal to

$$E(a_1) = \mu_{L,1} + 0.56419\, \sigma_{L,1} \tag{111}$$

$$E(a_2) = \mu_{L,2} + 0.56419\, \sigma_{L,2} \tag{112}$$

and

$$VAR(a_1) = 0.68169\, \sigma_{L,1}^2 \tag{113}$$

$$VAR(a_2) = 0.68169\, \sigma_{L,2}^2 \tag{114}$$

So the combined temporal order judgment can be characterized as it was in equations 23 and 24.

APPENDIX F

THE PSYCHOMETRIC FUNCTIONS FROM THE PILOT EXPERIMENT

APPENDIX F

THE PSYCHOMETRIC FUNCTIONS FROM THE PILOT EXPERIMENT

Results for Observer GEN (1) at 1.1 log unit above threshold in the pilot experiment.

Viewing Condition	RIGHT EYE		LEFT EYE		BOTH EYES	
d (ms)	n	P(S1)	n	P(S1)	n	P(S1)
-50			1	0.000		
-45			2	0.000		
-40			3	0.000		
-35	1	0.000	5	0.000		
-30	8	0.125	12	0.000	1	0.000
-25	30	0.067	33	0.091	18	0.000
-20	71	0.211	70	0.129	33	0.121
-15	113	0.221	109	0.220	77	0.117
-10	90	0.300	109	0.330	128	0.250
-5	66	0.349	72	0.431	85	0.341
0	53	0.566	60	0.533	81	0.333
5	60	0.583	75	0.547	98	0.551
10	94	0.628	117	0.709	106	0.708
15	88	0.727	112	0.768	90	0.778
20	63	0.905	58	0.897	43	0.814
25	20	0.850	21	0.905	11	0.909
30	6	1.000	14	0.929	3	1.000
35	1	1.000	6	1.000		
40			4	1.000		
45			3	1.000		
50			2	1.000		
PROBIT ANALYSIS						
Correlation		0.9735		0.9944		0.9886
Mean		0.4855		-0.2321		2.8262
Standard Deviation		20.7538		18.7175		16.5138

Results for Observer GEN (2) at 1.1 log unit above threshold in the pilot experiment.

Viewing Condition	RIGHT EYE		LEFT EYE		BOTH EYES	
d (ms)	n	P(S1)	n	P(S1)	n	P(S1)
-40			1	0.000		
-35	2	0.000	6	0.167		
-30	17	0.059	18	0.167	2	0.000
-25	39	0.180	44	0.114	19	0.053
-20	83	0.205	80	0.238	70	0.114
-15	142	0.247	120	0.167	106	0.245
-10	125	0.448	113	0.372	120	0.208
-5	63	0.540	61	0.426	98	0.316
0	57	0.597	64	0.453	110	0.500
5	55	0.673	56	0.804	103	0.553
10	96	0.677	107	0.701	126	0.683
15	130	0.800	123	0.691	106	0.793
20	87	0.828	88	0.852	59	0.797
25	47	0.957	40	0.900	27	0.926
30	11	0.909	15	0.867	5	1.000
35	2	1.000	5	0.000		
40			1	0.000		
PROBIT ANALYSIS						
Correlation		0.9790		0.9529		0.9897
Mean		-3.4383		-0.7769		1.6710
Standard Deviation		21.6791		22.0642		18.1906

Results for observer GEN (3) at 0.7 log unit above threshold in the pilot experiment.

Viewing Condition	RIGHT EYE		LEFT EYE		BOTH EYES	
d (ms)	n	P(S1)	n	P(S1)	n	P(S1)
-50	26	0.000	30	0.000	30	0.000
-45	50	0.040	52	0.039	61	0.000
-40	66	0.106	62	0.081	61	0.033
-35	69	0.145	73	0.055	77	0.026
-30	66	0.106	63	0.175	84	0.083
-25	69	0.130	81	0.111	81	0.124
-20	81	0.247	81	0.222	77	0.156
-15	72	0.181	92	0.359	100	0.180
-10	119	0.395	117	0.325	113	0.204
-5	131	0.435	140	0.414	131	0.366
5	138	0.652	138	0.507	130	0.446
10	108	0.722	104	0.625	116	0.603
15	86	0.791	83	0.747	79	0.760
20	76	0.895	75	0.907	66	0.894
25	64	0.891	61	0.869	67	0.866
30	63	0.921	71	0.916	76	0.947
35	79	0.987	68	0.971	52	0.981
40	72	1.000	61	0.967	61	1.000
45	70	1.000	54	0.982	47	1.000
50	26	1.000	19	1.000	20	1.000
PROBIT ANALYSIS						
Correlation		0.9817		0.9885		0.9868
Mean		-3.7302		-0.9580		2.3287
Standard Deviation		21.7739		23.3338		19.3102

Results for observer GEN (4) at 0.4 log unit above threshold in the pilot experiment.

Viewing Condition	RIGHT EYE		LEFT EYE		BOTH EYES	
d (ms)	n	P(S1)	n	P(S1)	n	P(S1)
-50	1	0.000				
-40	14	0.071	7	0.000	7	0.000
-30	51	0.196	44	0.136	42	0.024
-20	76	0.329	99	0.232	63	0.286
-10	48	0.438	93	0.441	54	0.333
0	33	0.455	78	0.551	45	0.667
10	44	0.568	88	0.705	48	0.583
20	54	0.796	95	0.800	71	0.718
30	46	0.913	58	0.931	45	0.889
40	15	1.000	11	1.000	5	1.000
50	1	1.000				
PROBIT ANALYSIS						
Correlation		0.9789		0.9954		0.9331
Mean		-3.6392		-3.3786		0.2937
Standard Deviation		29.1746		24.3450		25.9014

Results for observer GEN (5) at 0.4 log unit above threshold in the pilot experiment.

Viewing Condition	RIGHT EYE		LEFT EYE		BOTH EYES	
d (ms)	n	P(S1)	n	P(S1)	n	P(S1)
-50			1	0.000		
-48			2	0.000		
-46						
-44			2	0.000		
-42			3	0.000		
-40	3	0.000	4	0.250		
-38			2	0.000		
-36			13	0.077		
-34			13	0.385		
-32	8	0.125	18	0.056	3	0.000
-30	8	0.000	25	0.120	2	0.500
-28	5	0.200	46	0.304	5	0.000
-26	10	0.200	33	0.303	6	0.333
-24	14	0.071	42	0.310	7	0.143
-22	29	0.276	36	0.222	12	0.250
-20	45	0.222	35	0.314	18	0.111
-18	52	0.289	23	0.391	34	0.147
-16	58	0.259	18	0.389	41	0.366
-14	47	0.234	15	0.133	39	0.256
-12	42	0.357	11	0.546	46	0.261
-10	29	0.310	7	0.286	27	0.370
-8	25	0.320	3	0.667	21	0.429
-6	11	0.364	2	0.500	23	0.174
-4	8	0.375	1	1.000	17	0.353
-2	1	0.000	1	0.000	8	0.625
2	3	0.000	1	1.000	4	0.500
4	4	0.750			12	0.583
6	15	0.467	1	1.000	13	0.769
8	21	0.714	2	0.500	9	0.778
10	23	0.652	5	0.600	30	0.500
12	36	0.528	7	0.571	40	0.650
14	68	0.691	14	0.500	38	0.763
16	43	0.814	21	0.619	36	0.694
18	33	0.697	26	0.731	37	0.811
20	36	0.778	41	0.683	16	0.750
22	32	0.813	33	0.667	7	0.857
24	29	0.828	46	0.804	9	0.667
26	11	0.818	38	0.737	5	0.800

Results for observer GEN (5) at 0.4 log unit above threshold in the pilot experiment (continued).

28	9	0.778	16	0.750	3	0.667
30	4	0.750	34	0.794	4	0.500
32	2	0.500	21	0.857	3	1.000
34	1	1.000	5	1.000		
36			13	0.923		
38			7	0.857		
40			4	1.000		
42			3	1.000		
44			2	1.000		
46			2	1.000		
48			1	1.000		
50						
52			2	1.000		

PROBIT ANALYSIS			
Correlation	0.9032	0.8954	0.7813
Mean	0.9283	-1.1043	1.4684
Standard Deviation	27.5932	36.1494	26.0600

Results for observer MWH (1) at 0.7 log unit above threshold in the pilot experiment.

Viewing Conditions	RIGHT EYE		LEFT EYE		BOTH EYES	
d (ms)	n	P(S1)	n	P(S1)	n	P(S1)
-50	14	0.000	14	0.143	16	0.063
-45	20	0.000	20	0.000	28	0.036
-40	27	0.148	28	0.143	28	0.071
-35	32	0.094	32	0.125	34	0.235
-30	48	0.292	48	0.167	44	0.296
-25	64	0.250	64	0.313	72	0.222
-20	60	0.483	60	0.317	84	0.286
-15	66	0.394	66	0.394	75	0.320
-10	112	0.411	112	0.402	114	0.412
-5	72	0.514	72	0.528	76	0.500
5	74	0.703	74	0.568	76	0.592
10	88	0.636	88	0.636	96	0.656
15	83	0.735	84	0.714	80	0.763
20	58	0.810	58	0.690	68	0.941
25	74	0.892	74	0.878	62	0.887
30	64	0.938	64	0.922	30	1.000
35	44	0.909	44	0.955	36	0.972
40	34	0.912	34	0.971	36	1.000
45	28	0.893	28	0.964	14	1.000
50	4	1.000	4	1.000	8	1.000

PROBIT ANALYSIS	RIGHT EYE	LEFT EYE	BOTH EYES
Correlations	0.9679	0.9783	0.9731
Mean	-6.4813	-4.3211	-5.8500
Standard Deviation	30.2240	29.4552	25.2288

Results for observer JLZ at 1.0 log above threshold in the pilot experiment

Viewing Condition	RIGHT EYE		LEFT EYE		BOTH EYES	
d (ms)	n	P(S1)	n	P(S1)	n	P(S1)
-50	4	0.000	4	0.250	4	0.250
-45	12	0.333	12	0.167	12	0.083
-40	16	0.313	16	0.250	16	0.063
-35	20	0.200	20	0.100	20	0.100
-30	28	0.214	27	0.296	28	0.214
-25	35	0.257	36	0.250	36	0.306
-20	40	0.175	40	0.350	40	0.325
-15	44	0.341	44	0.500	44	0.386
-10	48	0.271	47	0.681	48	0.375
-5	47	0.489	48	0.479	47	0.404
5	55	0.527	56	0.607	56	0.429
10	48	0.563	48	0.708	47	0.702
15	44	0.546	44	0.796	44	0.614
20	40	0.500	40	0.625	40	0.700
25	36	0.833	36	0.944	35	0.857
30	28	0.786	28	0.893	28	0.786
35	20	0.750	20	0.850	20	0.950
40	16	0.688	16	0.938	16	0.938
45	12	0.833	12	0.917	12	1.000
50	4	0.750	4	1.000	4	1.000
55	8	1.000	8	1.000	8	1.000
PROBIT ANALYSIS						
Correlation		0.8974		0.9386		0.9456
Mean		4.4130		-9.3469		-0.9856
Standard Deviation		45.2682		35.4393		32.4260

APPENDIX G

THE PSYCHOMETRIC FUNCTIONS FROM EXPERIMENT 1

APPENDIX G

THE PSYCHOMETRIC FUNCTIONS FROM EXPERIMENT 1

Results for observer GEN (6) at 1.0 log unit above threshold in experiment 1.

Viewing Condition	RIGHT EYE		LEFT EYE		BOTH EYES	
d (ms)	n	P(S1)	n	P(S1)	n	P(S1)
-50	27	0.037	31	0.065	36	0.000
-45	25	0.040	26	0.154	29	0.034
-40	79	0.114	71	0.169	91	0.011
-35	37	0.135	30	0.300	37	0.027
-30	94	0.128	81	0.346	89	0.067
-25	51	0.137	44	0.295	36	0.111
-20	103	0.282	108	0.370	97	0.134
-15	51	0.333	42	0.476	49	0.224
-10	138	0.362	131	0.580	131	0.313
-5	95	0.442	80	0.525	96	0.365
5	65	0.462	76	0.737	92	0.565
10	141	0.546	114	0.754	160	0.606
15	55	0.618	58	0.845	55	0.782
20	92	0.761	96	0.885	89	0.798
25	40	0.725	43	0.930	43	0.837
30	81	0.802	89	0.876	98	0.878
35	49	0.857	39	0.923	29	0.931
40	77	0.909	87	0.977	85	0.894
45	31	0.871	31	0.935	34	1.000
50	36	0.944	37	0.973	40	0.975
PROBIT ANALYSIS						
Correlation		0.9909		0.9848		0.9903
Mean		2.4885		-12.7174		2.9427
Standard Deviation		31.9589		30.7947		22.5474

Results for observer GEN (7) at 1.4 log units above threshold in experiment 1.

Viewing Condition	RIGHT EYE		LEFT EYE		BOTH EYES	
d (ms)	n	P(S1)	n	P(S1)	n	P(S1)
-50	20	0.050	20	0.000	17	0.000
-45	27	0.000	30	0.067	36	0.000
-40	33	0.000	35	0.086	41	0.024
-35	35	0.057	38	0.158	49	0.020
-30	43	0.070	41	0.122	42	0.095
-25	36	0.056	34	0.294	43	0.047
-20	43	0.163	45	0.400	54	0.185
-15	50	0.140	42	0.333	53	0.321
-10	64	0.344	51	0.412	74	0.351
-5	91	0.375	99	0.505	79	0.506
5	71	0.465	94	0.745	102	0.627
10	55	0.636	52	0.692	57	0.632
15	66	0.667	47	0.830	58	0.724
20	55	0.745	40	0.875	50	0.860
25	42	0.810	41	0.951	43	0.930
30	38	0.868	41	0.927	27	0.926
35	35	0.971	39	0.974	41	0.951
40	39	0.923	31	0.968	43	0.953
45	28	0.964	33	1.000	41	1.000
50	10	1.000	18	1.000	23	1.000
PROBIT ANALYSIS						
Correlation		0.9785		0.9860		0.9859
Mean		3.9757		-7.4819		-1.0948
Standard Deviation		22.6896		23.6684		20.4813

Results for observer GEN (8) at 0.6 log unit above threshold in experiment 1.

Viewing Condition	RIGHT EYE		LEFT EYE		BOTH EYES	
d (ms)	n	P(S1)	n	P(S1)	n	P(S1)
-50	10	0.000	16	0.375	6	0.000
-45	35	0.029	29	0.345	34	0.088
-40	40	0.050	33	0.182	36	0.028
-35	42	0.024	36	0.250	31	0.097
-30	37	0.189	41	0.122	45	0.133
-25	51	0.157	41	0.415	48	0.146
-20	43	0.186	50	0.500	45	0.178
-15	40	0.200	41	0.463	59	0.254
-10	63	0.333	62	0.548	67	0.358
-5	79	0.367	66	0.606	77	0.468
5	83	0.530	72	0.722	95	0.611
10	72	0.597	57	0.719	62	0.726
15	67	0.746	55	0.782	56	0.732
20	47	0.702	52	0.923	44	0.727
25	50	0.760	48	0.854	27	0.741
30	35	0.829	39	0.923	41	0.927
35	38	0.895	44	0.909	44	0.955
40	45	0.867	52	0.904	38	0.921
45	29	0.897	34	0.971	40	0.975
50	22	0.955	13	1.000	15	0.867

PROBIT ANALYSIS			
Correlation	0.9856	0.9404	0.9399
Mean	0.3620	-1.4879	-0.6152
Standard Deviation	27.5245	35.8110	25.8056

Results for observer MWH (2) at 0.6 log unit above threshold in experiment 1.

Viewing Condition	RIGHT EYE		LEFT EYE		BOTH EYES	
d (ms)	n	P(S1)	n	P(S1)	n	P(S1)
-50	16	0.188	15	0.267	18	0.111
-45	34	0.235	48	0.188	38	0.079
-40	39	0.359	42	0.048	35	0.029
-35	41	0.146	47	0.298	50	0.200
-30	39	0.359	39	0.256	47	0.191
-25	51	0.314	54	0.389	56	0.304
-20	46	0.348	53	0.283	46	0.348
-15	65	0.385	52	0.481	64	0.516
-10	62	0.581	83	0.590	70	0.557
-5	69	0.623	101	0.564	85	0.600
5	100	0.700	100	0.630	96	0.667
10	62	0.629	78	0.731	88	0.761
15	39	0.821	52	0.750	60	0.800
20	51	0.784	53	0.943	57	0.877
25	45	0.933	50	0.820	68	0.868
30	64	0.812	38	0.895	41	0.805
35	32	0.875	59	0.932	56	0.875
40	47	0.851	44	0.909	56	0.929
45	42	0.857	32	0.938	39	0.949
50	21	0.905	16	1.000	14	1.000

PROBIT ANALYSIS			
Correlation	0.9456	0.9404	0.9675
Mean	-11.7728	-11.8902	-10.3420
Standard Deviation	42.3869	34.4628	31.4517

Results for observer MWH (3) at 1.0 log unit above threshold in experiment 1.

Viewing Condition	RIGHT EYE		LEFT EYE		BOTH EYES	
d (ms)	n	P(S1)	n	P(S1)	n	P(S1)
-50	18	0.056	18	0.222	13	0.077
-45	35	0.057	26	0.154	37	0.216
-40	36	0.222	45	0.111	41	0.146
-35	32	0.094	41	0.171	36	0.083
-30	46	0.239	33	0.273	38	0.132
-25	48	0.250	44	0.341	54	0.185
-20	49	0.327	56	0.393	51	0.255
-15	58	0.276	57	0.368	52	0.385
-10	62	0.435	66	0.515	65	0.492
-5	86	0.535	75	0.467	79	0.354
5	91	0.593	67	0.642	80	0.625
10	62	0.629	55	0.727	66	0.652
15	50	0.660	54	0.741	67	0.761
20	54	0.778	49	0.796	54	0.685
25	53	0.736	55	0.945	37	0.784
30	40	0.775	32	0.875	42	0.857
35	30	0.933	36	0.944	36	0.861
40	59	0.847	55	1.000	37	0.946
45	33	0.970	32	0.906	26	0.885
50	21	0.905	25	0.920	20	0.950
PROBIT ANALYSIS						
Correlation		0.9726		0.9661		0.9745
Mean		-2.2304		-8.6058		-2.0808
Standard Deviation		34.7854		31.4625		32.6717

Results for observer MWH (4) at 1.4 log units above threshold in experiment 1.

Viewing Condition	RIGHT EYE		LEFT EYE		BOTH EYES	
d (ms)	n	P(S1)	n	P(S1)	n	P(S1)
-50	9	0.000	6	0.000	14	0.000
-45	21	0.095	29	0.069	27	0.037
-40	41	0.024	44	0.182	40	0.125
-35	38	0.105	33	0.273	37	0.108
-30	38	0.158	33	0.273	38	0.079
-25	44	0.364	29	0.310	49	0.265
-20	48	0.312	42	0.452	51	0.373
-15	45	0.600	59	0.475	56	0.393
-10	67	0.507	48	0.562	66	0.515
-5	83	0.554	79	0.620	73	0.493
5	90	0.722	94	0.734	70	0.671
10	55	0.673	50	0.800	61	0.820
15	58	0.828	38	0.737	52	0.788
20	55	0.782	42	0.833	42	0.976
25	36	0.861	44	0.977	47	0.851
30	37	0.973	30	0.933	47	0.957
35	35	0.943	27	0.963	38	0.895
40	38	0.974	35	0.886	36	1.000
45	30	0.967	31	1.000	28	0.964
50	18	0.944	18	1.000	13	1.000
PROBIT ANALYSIS						
Correlation		0.9634		0.9599		0.9617
Mean		-8.3743		-12.9297		-7.6249
Standard Deviation		27.0161		29.2920		24.4061

APPENDIX H

THE ANALYSES OF VARIANCE

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THE ANALYSES OF VARIANCE

The analysis of variance performed on the standard deviations in experiment 1.

SOURCE	df	SS	mSS	F	
Observer (O)	1	120.922	120.922	7.693	
Intensity (I)	2	222.695	111.347	7.084	$p < 0.05$
Eye (E)	2	90.713	45.356	2.886	
O X I	2	3.261	1.261	0.104	
O X E	2	28.577	14.288	0.909	
I X E	4	14.480	3.620	0.230	
O X I X E	4	62.876	15.719		
Total	17	543.522			

The analysis of variance performed on the means in experiment 1.

SOURCE	df	SS	mSS	F	
Observer (O)	1	217.452	217.452	49.208	$p < 0.0025$
Intensity (I)	2	22.903	11.452	2.592	
Eye (E)	2	157.889	78.944	17.865	$p < 0.02$
O X I	2	63.645	31.823	7.201	$p < 0.05$
O X E	2	24.481	12.240	2.770	
I X E	4	71.660	17.915	4.054	
O X I X E	4	17.676	4.419		
Total	17	575.706			

The analysis of variance performed on the reaction time standard deviations in experiment 2.

SOURCE	df	SS	mSS	F	
Observer (O)	1	3,723.664	3,723.664	87.663	$p < 0.001$
Intensity (I)	2	607.469	303.735	7.151	$p < 0.05$
Stimulus (S)	1	417.051	417.051	9.818	$p < 0.05$
Eye (E)	2	330.292	165.146	3.888	
O X I	2	43.027	21.513	0.507	
O X S	1	49.588	49.588	1.167	
O X E	2	114.850	57.425	1.352	
I X S	2	119.307	59.654	1.404	
I X E	4	78.715	19.679	0.463	
S X E	2	105.223	52.611	1.239	
O X I X S	2	12.014	6.007	0.141	
O X I X E	4	95.146	23.786	0.560	
O X S X E	2	223.262	111.631	2.628	
I X S X E	4	241.181	60.295	1.420	
O X I X S X E	4	169.908	42.477		
Total	35	6,330.697			

The analysis of variance performed on the reaction time mean values in experiment 2.

SOURCE	df	SS	mSS	F	
Observer (O)	1	12,285.099	12,285.099	1,640.419	$p < 0.000005$
Intensity (I)	2	9,408.118	4,704.059	628.129	$p < 0.00005$
Stimulus (S)	1	3,354.900	3,354.900	447.977	$p < 0.00005$
Eye (E)	2	4,439.325	2,219.662	296.390	$p < 0.00005$
O X I	2	1,757.160	878.580	117.316	$p < 0.0005$
O X S	1	380.036	380.036	50.746	$p < 0.005$
O X E	2	216.661	108.330	14.465	$p < 0.05$
I X S	2	97.849	48.925	6.533	
I X E	4	457.123	114.281	15.260	$p < 0.05$
S X E	2	318.964	159.482	21.296	$p < 0.01$
O X I X S	2	14.253	7.126	0.952	
O X I X E	4	187.917	46.979	6.273	
O X S X E	2	165.020	82.510	11.017	$p < 0.05$
I X S X E	4	73.864	18.466	2.466	
O X I X S X E	4	29.955	7.489		
Total	35	33,186.244			