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RESEARCH ARTICLE

Contextual effects in interval-duration judgements in vision, audition and touch

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Abstract We examined the effect of temporal context on discrimination of intervals marked by auditory, visual and tactile stimuli. Subjects were asked to compare the duration of the interval immediately preceded by an irrelevant "distractor" stimulus with an interval with no distractor. For short interval durations, the presence of the distractor affected greatly the apparent duration of the test stimulus: short distractors caused the test interval to appear shorter and vice versa. For very short reference durations (≤100 ms), the contextual effects were large, changing perceived duration by up to a factor of two. The effect of distractors reduced steadily for longer reference durations, to zero effect for durations greater than 500 ms. We found similar results for intervals defined by visual flashes, auditory tones and brief finger vibrations, all falling to zero effect at 500 ms. Under appropriate conditions, there were strong cross-modal interactions, particularly from audition to vision. We also measured the *Weber fractions* for duration discrimination and showed that under the conditions of this experiment, Weber fractions decreased steadily with duration, following a square-root law, similarly for all three modalities. The magnitude of the effect of the distractors on apparent duration correlated well with Weber fraction, showing that when duration discrimination was relatively more precise, the context dependency was less. The results were well fit by a simple Bayesian model combining noisy estimates of duration with the action of a resonance-like mechanism that tended to regularize the sound sequence intervals.

Keywords Time perception · Contextual effects · Weber fraction · Bayesian models

Introduction

While there has been great progress in understanding space perception in recent years, perception of time remains poorly understood. Traditional models of timing over the millisecond–second range assume the existence of a dedicated timing mechanism, usually involving an internal clock driven by an oscillator or pacemaker emitting pulses that are counted by an accumulator (Creelman 1962; Treisman 1963; Gibbon 1977). The pulse count provides a linear metric of time, with temporal judgements relying on comparison of pulse counts under various conditions. Dedicated models do not necessarily require a single driving pacemaker: they can be based, for example, on a delay-line principle (Ivry 1996).

An alternative to dedicated models are the recently formulated *intrinsic* models, which assume that timing is an inherent property of neural processing, rather than the result of dedicated mechanisms. The best known of this class of model comes from Buonomano's group (Buonomano and

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Merzenich 1995; Buonomano 2000; Buonomano and Karmarkar 2002; Maass et al. 2002; Buhusi and Meck 2005). These models have enjoyed reasonable success, especially in the sub-second range of interval judgements. One prediction of the models is that temporal judgements of a particular interval will not always be the same, but may depend on the duration of a previous interval. This hypothesis was validated psychophysically and shown to occur only for the sub-second range of intervals (Karmarkar and Buonomano 2007): an interval around 100 ms was judged with less precision when preceded by shorter or longer intervals; however, judgements of intervals around 1,000 ms were unaffected by preceding intervals. This effect was confirmed by Spencer et al. (2009), who went on to show that the irrelevant distractors not only affected precision, but also affected the perceived duration of short-interval sounds. Their new data challenged Karmarker and Buonomano's interpretation in support of state-dependent networks.

That the temporal order of stimuli can affect how they are perceived (including duration) has been known since Fechner's time (see Hellstrom (1985), for an excellent review of "time-order errors"). In particular, many studies (especially in the field of rhythm perception) have reported that irrelevant "distractors" can affect apparent duration of subsequently presented brief stimuli. One clear example is the "time-shrinking" illusion (Nakajima et al. 1991, 1992, 2004). A short distractor interval presented before a short (~200 ms) test interval can shorten the apparent duration of the test, by up to 40 %. They describe this effect as a process of assimilation, which tends to make the two intervals (distractor and test) similar. Similar effects have been reported in other sensory modalities, including vision (Arao et al. 2000) and touch (Van Erp and Spapé 2008). Similarly, McAuley and Jones (2003), Jones and McAuley (2005) showed that sequences of tones influenced the apparent duration of the time interval at the end of the sequence (see also Monahan and Hirsh 1990). Again the effect was consistent with assimilation, short intervals shortening the apparent duration and long intervals lengthening it. The authors explain these results with the concept of "entrainment", a dynamic system of self-sustaining neural oscillations that maintain an internal beat.

The goal of the present series of experiments is to study in more detail contextual effects in time perception and see how they relate to precision of interval discrimination. We measured the effect of neighbouring "distractors" on the apparent duration of stimuli, over a wide range of durations, with a paradigm similar to that of Karmarkar and Buonomano (2007). We show that abutting distractors produce systematic biases in apparent duration, with short distractors causing intervals to seem shorter and vice versa. The effects are maximal at short intervals, gradually diminishing to zero at about 1 s. The data are well explained by a Bayesian model of interval regularization, which takes into account the precision of interval discrimination.

Materials and methods

Subjects

Five subjects (three female, mean age 25 years), of which four were naïve to the goals of the study, participated in the research. All had normal (or corrected-to-normal) vision and normal hearing. Participants gave informed consent to participate in the study according to the guidelines of the University of Florence and ERC project STANIB. The tasks were performed in a dimly lit and sound-attenuated room.

Apparatus and stimuli

Auditory stimuli were generated in Matlab and presented through the two high-quality loudspeakers (Yamaha MSP5) at 65 kHz. They were brief (4 ms) pulses of pure tones, whose frequency varied slightly from presentation to presentation, following a random Gaussian distribution of mean 200 Hz and standard deviation 20 Hz.

Visual stimuli were generated by framestore (Cambridge VSG 2/3) and displayed on a Clinton Monoray monitor, equipped with DP104 type fast phosphor, at a resolution of 640 × 480 pixels and frame rate of 200 Hz. The visual stimuli were black circles of 3° diameter when viewed from 57 cm, presented for one frame (5 ms) on a yellow background of 40 cd/m^2.

Haptic stimuli were created using a mechanical device that delivered a mild pressure to the index finger, gated by a switch under control of the VSG framestore. The duration of the stimuli was 10 ms.

Procedure

Subjects were presented with two pairs of stimuli (visual, auditory or haptic) delimiting an empty interval—a reference pair followed 2 s later by a test pair (Fig. 2a)—and were required to report which interval appeared to be longer (two-alternative-force choice). The reference stimulus was of fixed interval (from 50 to 1,000 ms) that remained constant in any given session. The test was of variable interval around the interval of the reference. The value of the test interval was determined by the adaptive algorithm QUEST (Watson and Pelli 1983), which homed in on the point of subjective equality, to which was added a random jitter (following a Gaussian of standard deviation 15 % of the reference interval). The advantages of this technique are that the data are collected in the most useful intervals

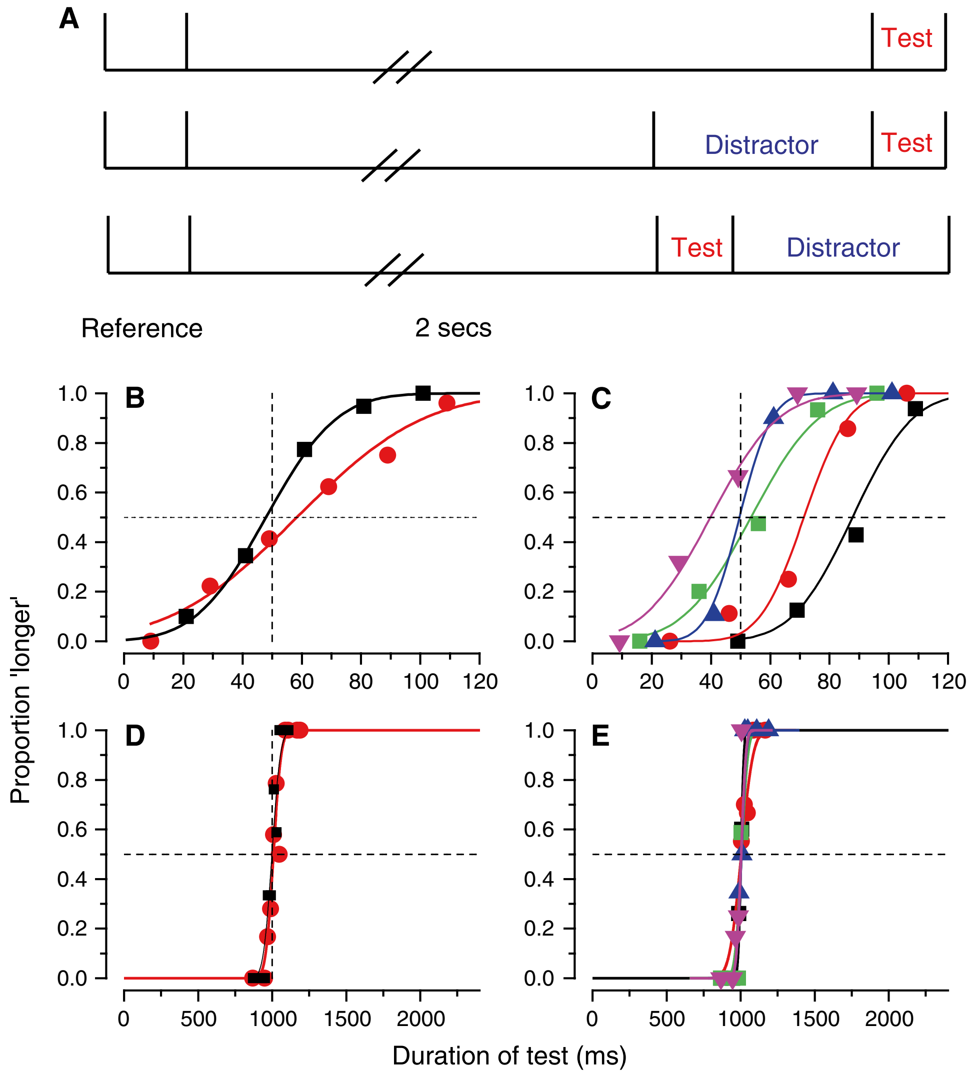

Fig. 1 **a** Schematic of the time course of the stimuli. The *upper timeline* shows the non-distractor condition where two pairs of stimuli were presented, a *reference* pair (of fixed duration within a given session, but varying between session from 50 to 1,000 ms), followed 2 s later by a *test* stimulus pair, whose duration varied around that of the reference within a session. The *second trace* shows the stimulus sequence with a leading distractor (in *blue*), used in most experiments. The *third* shows the sequence used for the final experiment, when the order of test and distractor were inverted. **b–e** Example psychometric functions for base durations of 50 ms (**b**, **c**) and 1,000 ms (**d**, **e**). **b** Proportion of trials in which the probe was judged longer than the test, under conditions of no distractor (*black symbols*) or distractors of variable duration (25, 37.5, 50, 62.5 or 75 ms: all data intermixed). The psychometric function with distractors is broader than that without (34 cf 19 ms). **c** Data of the distractor conditions of **c** plotted separately for each distractor duration (*black* 25 ms, *red* 37 ms, *green* 50 ms, *blue* 632 ms, *magenta* 75 ms). The effect of the distractors is clear to shift the psychometric functions in the direction opposite to distractor duration, without affecting systematically curve width. **d, e** Same as **b** and **c**, except that all times—reference and distractors—are 20 times longer (reference 1,000 ms, distractors 500, 750, 1,000, 1,250 and 1,500 ms). The distractors have no effect at all, either on the width or the PSE of the psychometric functions (colour figure online)

and that the responses distribute evenly between "shorter" and "longer". Data were analysed by fitting Gaussian error functions (like those of Fig. 1), from which the point of subjective equality (PSE) was given by the mean and the precision by the standard deviation.

In the main experiment, the test interval was preceded by a "distractor interval" (also empty), presented immediately prior to the test (Fig. 2a). Within a given session, the duration of the reference was constant, chosen in pseudo-random order from the set: 50, 70, 100, 140, 300, 550 and 1,000 ms. Within each session, the different distractors were interleaved, so its duration varied unpredictably from trial to trial. In all conditions, the distractor intervals were 0.25, 0.5, 1, 1.25 or 1.5 times the reference interval, so the actual values were scaled with the reference. Subjects were asked to ignore the distractor interval and base their

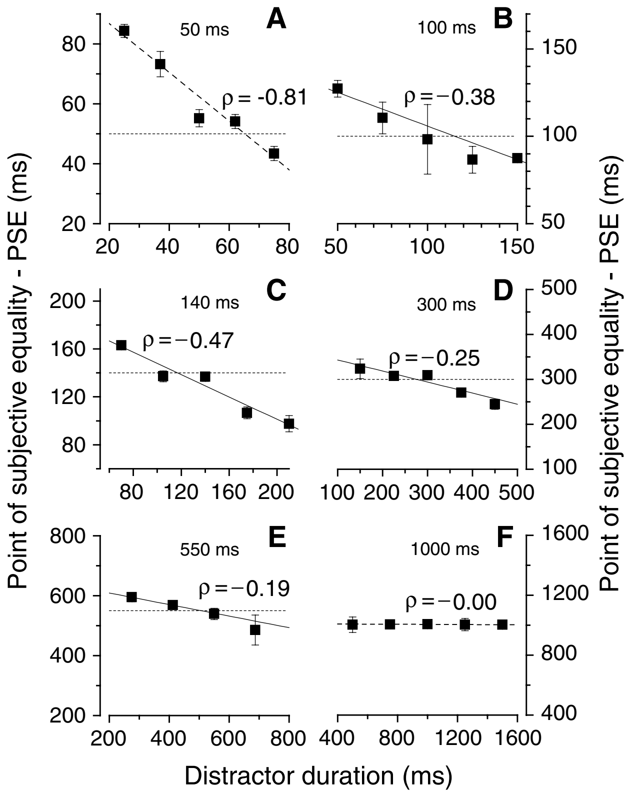

Fig. 2 PSE for judging duration of auditory stimuli, as a function of the duration of the distractor, for base durations of 50, 100, 140, 300, 550 or 1,000 ms (observer EDR). For the short base duration, distractors shifted considerably the PSE: short distractors caused the PSE to lengthen, and vice versa. The slope of the best-fitting line is taken as an index of the effectiveness of the distractors, where a slope of −1 means that the change in PSE is equal to the difference in distractor and reference intervals. As base duration increased, the distractors had progressively no effect, reaching zero for base durations of 1,000 ms. The PSEs are well fit by linear regression (average $R^2 = 0.96$)

judgements comparing the apparent interval of the test with that of the reference. Separate sessions were run for each reference interval with no distractors. We also measured the effect with distractors that followed rather than preceded the test interval, with all other conditions as before.

Results

The effect of distractors

The effect of distractor stimuli on perceived duration was examined for seven different base durations, ranging from 50 to 1,000 ms, for three different modalities. For each base duration, subjects matched a variable test duration to that of the reference, with five different distractor durations (randomly intermingled within each session). Figure 1b–d show example psychometric functions for one subject, for the auditory condition at two base durations (50 and 1,000 ms). The ranges on the abscissae are scaled in proportion to the duration of the reference interval. Figure 1b–d show results for the 50 and 1,000 ms duration, for the no-distractor conditions, and the distractor conditions pooled together (following Karmarkar and Buonomano 2007). The black symbols and fitted Gaussian error functions show results measured without distractor; the red symbols to those measured with the five distractor levels, all grouped together. For the short duration (Fig. 1b), the effect of distractors was detrimental, agreeing with the results of Karmarkar and Buonomano (2007): the psychometric functions became broader in the distractor condition, indicating an increase in discrimination thresholds. Thresholds, defined as the standard deviation of the fitted error function, increased from 19 ± 4 ms in the no-distractor condition to 34 ± 4 ms in the distractor condition. In this instance, there was also a slight change in PSE (the mean of the fitted function) in the distractor condition, but this was not a general trend. For the longer duration (Fig. 1d), the results were quite different. The distractors had virtually no effect, either on PSE or threshold (42 ± 11 ms with distractors compared with 46 ± 10 ms without), again in agreement with the results of Karmarkar and Buonomano (2007).

Figure 2c–e plot the data in a different way, separating out the trials for the different distractors, revealing a pattern that is not apparent when pooled. For short 50 ms reference stimuli (Fig. 2c), the distractors shift the curves, causing systematic changes in PSEs. For example, distractors of 25 ms shift the psychometric function rightwards, implying that the test to be physically longer than the distractor to appear perceptually equal: this means the distractor caused it to appear *shorter*. Longer distractors had the opposite effect, shifting the curves towards left, producing shorter PSEs, implying that they caused the test to appear *longer*. However, when plotted in this way, the thresholds with the different distractors were very similar, on average 14.0 ± 6 (compared with 19 ± 4 without distractors). Thus, the increased threshold when looking at pooled data is a consequence of pooling data of different PSE (or bias).

At base duration of 1,000 ms (Fig. 1e), the distractors had very little effect, with psychometric functions for all distractor conditions were virtually coincident, with no systematic shifts in either direction. Again this is consistent with the fact that thresholds for the pooled data are not degraded by distractors at this base duration. These results suggest that the deterioration in thresholds at low base durations, observed both here and by Karmaker and Buonomano, results from systematic contextual effects, not from a general degradation in time discrimination.

Figure 2 plots for one subject the PSEs of psychometric functions like those of Fig. 1, as a function of distractor duration. As noted above, for the brief test interval of 50 ms, the distractors affected strongly the apparent duration in a systematic fashion, with short durations increasing apparent duration and long distractors decreasing it. The effect in this example was roughly proportional to distractor duration, so the points are well fit by a linear regression ($R^2 = 0.96$) with slope −0.81.

Although the distractors were always spaced with the same proportionality, their effectiveness decreased progressively with base duration. As the base duration increased, the slope of the best-fitting linear regression decreased (Fig. 2a–f), until at 1,000 ms, the distractors had virtually no effect. As data for all conditions were well fit by linear regression (with $R^2 > 0.9$), we took the slope of this best-fitting line as an index of the effectiveness of the distractors; an index of unity means that a distractor that is half the duration of the test will cause the test to seem half as long and vice versa: that is, it makes the test seems to be as long as the distractor. A slope of zero means that the distractor had no effect. The slopes are shown near the best-fitting regression lines, decreasing almost monotonically from 0.81 at base duration of 50 ms to 0.00 at 1,000 ms.

Figure 3a plots indexes of distractor effectiveness as a function of duration of the reference. For all subjects, there was a strong effect of distractors at 50 ms, with indices greater than 0.5. And for all subjects, there was no effect at 1,000 ms. For all subjects, distractor effectiveness decreased steadily with duration to reach zero effect at 1,000 ms (some earlier than others). There was no clear-cut dissociation between short and long durations, but the effect of distractors varied gradually over a wide range. The solid curves show fits of a model described later.

The results presented so far were for brief auditory tones. To test whether the effects are generalizable to other domains, we repeated the experiment for brief visual stimuli (dark discs of 3° diameter), and also for haptic stimuli (brief vibrator pulses). As before, for a range of base durations, we measured psychometric functions for different distractor durations from which we calculated PSE and regressed the obtained PSEs against distractor duration to obtain an index of distractor effectiveness. Figure 3b plots contextual effects for visual stimuli as a function of reference duration, and Fig. 3c contextual effects for haptic stimuli. For all three modalities, the results are similar, although the effects are stronger for visual and haptic stimuli than for auditory stimuli. For all subjects, the effect of the distractors diminished gradually to reach zero at 1,000 ms. The continuous colour-coded curves show best fits of a simple Bayesian model, described later.

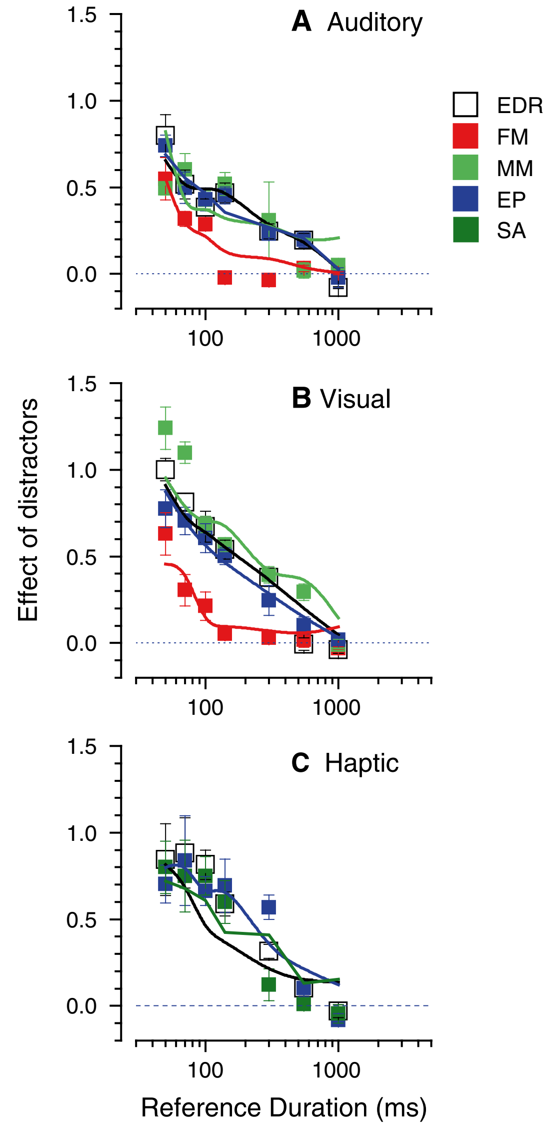

Fig. 3 Effectiveness of distractors as a function of base time, for auditory (**a**), visual (**b**) and haptic (**c**) stimuli. For each subject and each base duration, distractor effectiveness was calculated by regressing the PSE against distractor duration (see Fig. 2). The *error bars* are ±1 standard error of the fit. The *continuous curves* are the best fit of a model that combines a direct estimate of the duration, with a tendency to regularize the sequence (see "Modelling the contextual effects" section). The *various colours* refer to different subjects. RDR—*open black*; FM—*red*; MM—*green*; EP—*blue*; SA—*olive* (colour figure online)

The effect of distractors following the test

We then altered the order of the test and distractor intervals, so the test was first, followed by the distractor (lower timeline of Fig. 1a). The results of three observers, two for audition and vision, and one for all three modalities, are shown in Fig. 4 as a function of base interval (together with the distractor-first results for comparison). The results for the inverted sequence follow a similar pattern to the

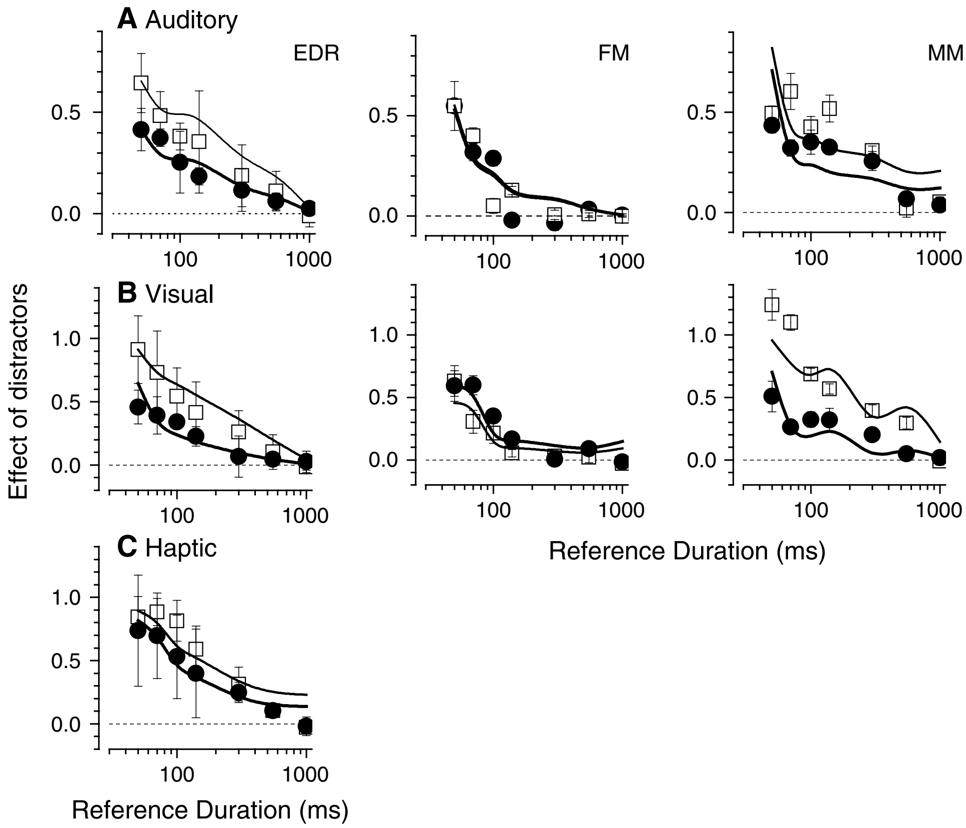

Fig. 4 Effect of trailing distractors (*filled circles*), for subjects EDR, FM and MM. The results with leading distractors are re-plotted from Fig. 3, with *open squares*. The procedure was identical to the other experiments, except that the distractors followed rather than preceded the test (*lower trace* of Fig. 1a). In most cases, trailing distractors were less effective than leading distractors, but were in other ways similar, in producing an assimilation effect at short but not base durations

distractor-first results, except that the effects tended to be weaker. These results show that the context effects are not strictly causal, but can operate in the reversed direction. Again the curves are model fits, discussed later.

Cross-modal distractors

We were interested whether auditory distractors could affect the judgement of visually marked intervals and vice versa. In the first version of this task, we replaced the first flash of the three-flash visual sequence with a noise burst, so the distractor interval was defined by a sound flash, and the test interval by a flash–flash (see cartoon of Fig. 5a). The reference stimulus remained, as before, a pair of flashes. The green symbols show the effect of distractors (calculated as before from the slopes of the PSE/distractor curves) as a function of reference duration for one subject, EDR. At no duration of the reference did the auditory distractor have any effect on the match of the visual stimulus. The red symbols show the results for the complementary experiment, where the distractor marker was a visual flash, and the other two sounds (with the reference defined by two sounds). In neither case did the cross-modal distractor affect the apparent duration of the test.

Figure 5b shows another version of the cross-modal experiment. In this, the centre marker of the three-stimulus presentation was a simultaneous flash and sound. The first was a sound and the last a flash (with a two-flash standard sequence: green symbols), or conversely, the first a flash and the last a sound (with a two-sound standard sequence: red symbols). Thus, both intervals—the distractor and the test—were defined by the same modality, sound or vision. In these conditions, the distractor had a strong effect at short intervals, particularly for the case of auditory distractor and visual test (green symbols). Moreover, the visual distractor affected the auditory test, to about half the extent.

We have data on only one subject, so these results must be considered preliminary. However, they suggest that for the distractors to be effective, the distractor interval needs to be defined by markers of the same modality. Provided that this is the case (with dual stimuli in the central position), the effect is as strong as for single modality stimuli.

Variation of Weber fractions with duration

The experimental procedure allows us to estimate not only the average perceived duration (from the PSE) but also the precision threshold, given by the standard deviation of the curve. Normalizing the threshold by the duration yields the *coefficient of variation*, a close approximation to the *Weber fraction* (threshold normalized by PSE, in this case very close to physical duration). Weber

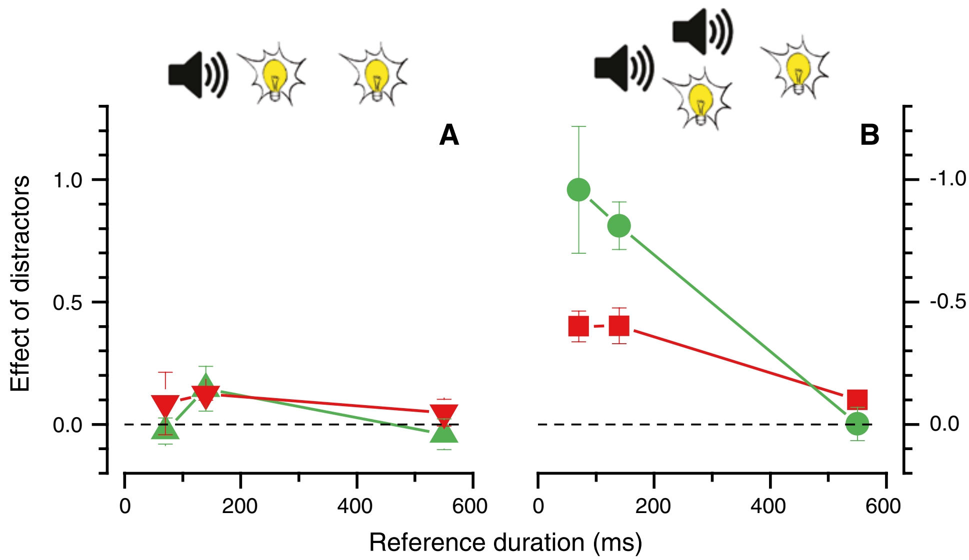

Fig. 5 Cross-modal interactions between vision and audition in subject EDR. **a** an auditory stimulus preceded two visual stimuli (shown in icon: *green symbols*), or a visual stimulus preceded two auditory stimuli (*red symbols*). In both cases, the test interval was demarked by stimuli of the same modality and the distractor a different modality: and in both cases, the distractor had no effect. **b** The stimuli were similar to those of **a**, except the central marker was bimodal, both visual and auditory. Under these circumstances, where both test and distractor intervals are demarked by stimuli of the same modality, the cross-modal distractor did affect the test duration, especially when the distractor was auditory (*green symbols*) (colour figure online)

fractions are often assumed to remain constant with duration (Gibbon 1977; Gibbon et al. 1997). However, inspection of the psychometric functions of Fig. 1 shows that Weber fractions were not constant with duration in our experiment. As the abscissae have been scaled by reference duration (so 50 ms of Fig. 1b–c occupy the same distance as 1,000 ms of Fig. 1d–e), psychometric functions should have the same slope in the two conditions, if their Weber fractions were equal. This is clearly not the case: all curves at short durations are much broader on the normalized scale than those of the long duration, reflecting the larger Weber fractions.

Figure 6 plots Weber fractions as a function of reference duration, for all subjects (colour-coded), for the auditory, visual and haptic modality. Weber fractions were obtained from the condition with no distractors (upper sketch of Fig. 1a). In all cases, the Weber fractions decrease substantially with reference duration. The dashed line above the data shows a square-root dependency (log–log slope of −0.5). The data, on average, tend to follow this trend. The slopes of the best-fitting regressions, given in the caption to Fig. 6, are on average −0.56, consistent with the square-root trend. As the Weber fraction is given by the ratio of threshold to reference duration, the inverse square-root dependency suggests that thresholds increase with the square-root of interval duration (rather than directly with interval duration, necessary for scalar constancy, or constant Weber fraction).

The results show that Weber fraction and context effects both decrease with reference duration. This suggests—qualitatively at least—that large context effects are associated with large Weber fractions. In addition, the contextual effects were greater for the visual and haptic than auditory modalities, and the Weber fractions are also higher in those modalities, consistent with the literature (e.g. Burr et al. 2009a).

Figure 7 plots the index of context dependency against Weber fraction (on log axis), for all subjects and modalities. The positive dependency is clearly apparent, with a linear regression (context effect against log Weber fraction) accounting for 66 % of the variance. This suggests that contextual effects are associated with poor temporal discrimination.

Modelling the contextual effects

That context dependency is strongly associated with Weber fractions suggests a possible causal link: when the Weber fraction is high (poor precision), the contextual effects are weak. It is possible that there exist effects akin to “entrainment” that attempt to regularize the interval sequences, but that this mechanism must compete sensory processes responding to the physical duration of the interval. As the Bayesian approach lends well itself to modelling competing effects, we propose here a simple Bayesian model, similar to that recent model of Sawai et al. (2012). This model

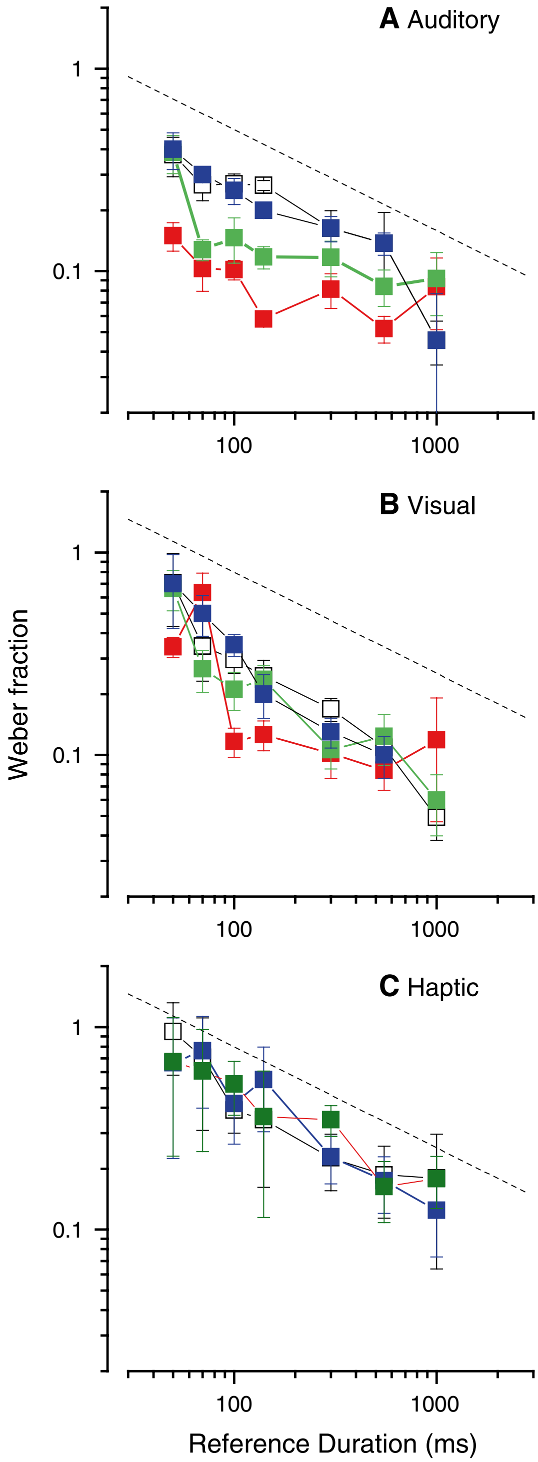

Fig. 6 Weber fractions for the five observers (*symbol colour* as Fig. 3) for the three different modalities. The *dashed black lines* indicate a square-root relationship (slope of 0.5 on log–log coordinates), clearly capturing the general trend (colour figure online)

does not attempt to explain the context dependency itself, but only why context dependency should vary inversely with duration of the reference.

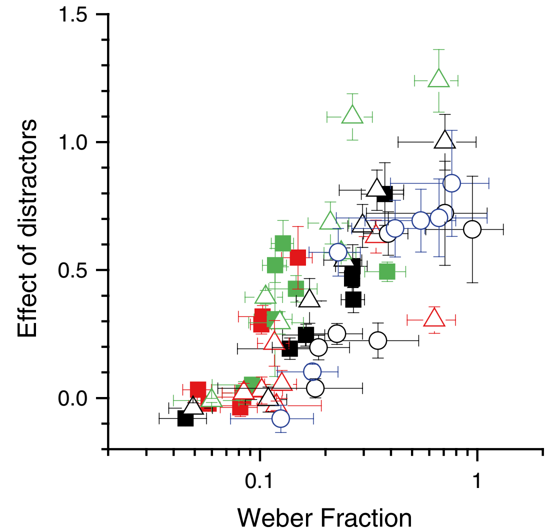

Fig. 7 Scatterplot of the effect of distractors against weber fraction, for the different subjects, modalities and durations. The *symbol colours* refer to different subject, as for Figs. 3 and 6, while *symbol shape* refers to modality: *squares*—audition; *triangles*—vision; haptic—*circles* (colour figure online)

To make the modelling explicit, we suppose that the assimilation results from a mechanism attempting to make the stimulus sequence regular, with equal intervals for distractor and test: this could b a non-linear filter tending towards resonance or any other process leading to duration assimilation or "entrainment" (McAuley and Jones 2003; Jones and McAuley 2005). Figure 8a illustrates how this hypothetical process may work with a specific example: a 100 ms reference, with 50 and 150 ms distractors. In both cases, the regularizing process works to shift the border separating the distractor and test, in one case forwards and in the other, backwards. To compensate for the shift, the test needs either to be increased in duration (for the short distractor) or decreased for the long distractor. If the regularization were complete, the shift would be equal to the difference in test and distractor (as shown in the example), resulting in a distractor index of 1.

We assume that this process is only a tendency, competing with other mechanisms that estimate the physical duration of the interval. The statistically optimal method of combining two sources of information is to multiply the underlying distributions, which for simplicity, we assume to be normal although the argument by no means rests on this assumption:

$$N_{\mathrm{PR}}\left(\mu_{\mathrm{PR}}, \sigma_{\mathrm{PR}}^{2}\right) = N_{\mathrm{P}}\left(\mu_{\mathrm{P}}, \sigma_{\mathrm{P}}^{2}\right) N_{\mathrm{R}}\left(\mu_{\mathrm{R}}, \sigma_{\mathrm{R}}^{2}\right) \quad (1)$$

where the subscripts P and R refer, respectively, to the physical and regularizing estimates of duration.

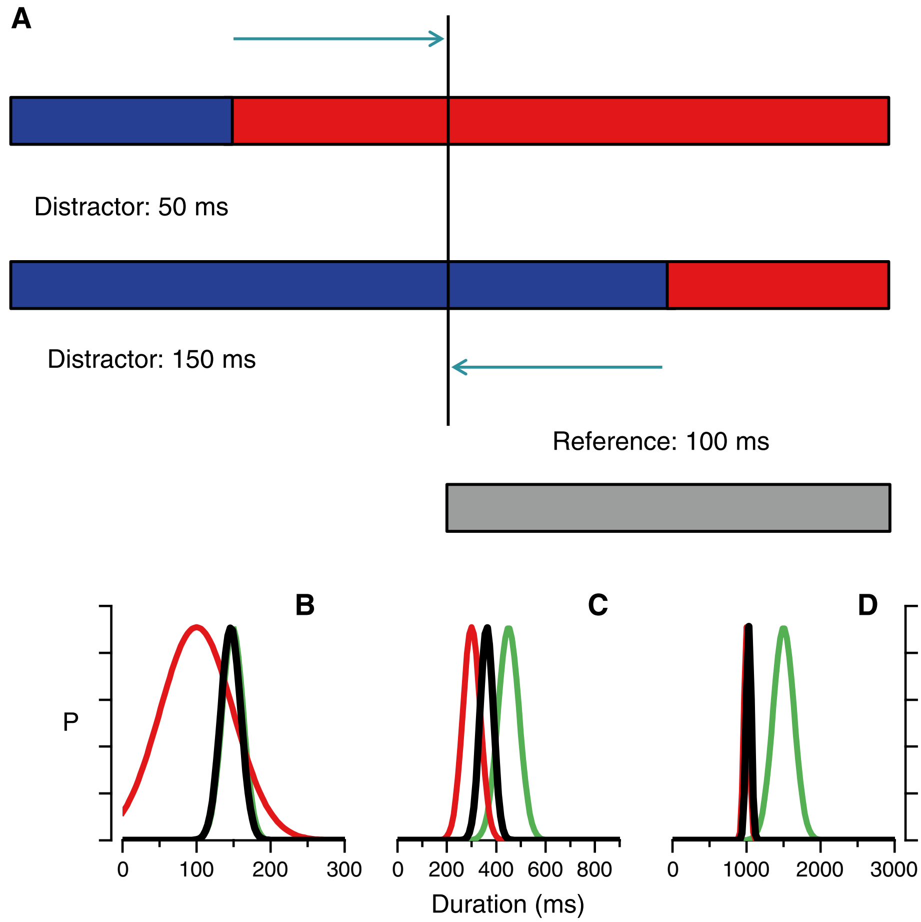

Fig. 8 Illustration of the "Bayesian-like" model used to fit the data of Fig. 3. **a** Illustration of the action of the hypothesized regularization mechanism that acts to move the boundary of distractor and test towards the centre of the full stimulus. **b–d** Illustration of how the Bayesian prior interacts with the likelihood (estimate of duration) within the Bayesian model, for short (150 ms), medium (300 ms) and long (1,000 ms) durations. The prior (in *green*) is assumed to be the same normalized width in the three conditions, corresponding to a Weber fraction of 0.2. The Weber fraction of the likelihood (*red curves*) vary with duration, broad a short durations, narrow at long durations (data from direct measurements shown in Fig. 6). At short durations, the prior dominates in determining the posterior (*black curve*), at long durations the likelihood dominates. At intermediate durations, both contribute to the duration estimate (colour figure online)

Figure 8b–d illustrate the process for three reference intervals, using the data of subject EDR's visual judgements. The red curves show the probability distributions for the physical estimate, green those of the regularizing distribution and black the resultant posterior, the product of the two. The variance of the physical estimate was obtained from the data of Fig. 6. We have no measure of the variance of the supposed regularizing process, but for convenience assume that it is a proportion of the reference interval and allow this parameter to vary freely to optimize the fitting process (one free parameter for each fitted curve of Fig. 3). For the example shown in Fig. 8 (EDR vision), it was estimated to be 0.2 (see Table 1). For a reference of 50 ms, EDR's Weber fraction was 0.7, yielding a broad distribution (Fig. 8b), broader than the regularizing distribution, which dominated on multiplication, predicting contextual effects that are nearly maximum. At 1,000 ms (Fig. 8d), however, the reverse holds: the Weber fraction is far smaller, resulting in a much tighter distribution, which dominates on multiplicative combination. At 300 ms (Fig. 8c), the distributions have similar width, so both contribute to the estimate, which results in an intermediate contextual effect of the distractors.

In practice, it can be demonstrated from Eq. 1 that the optimal method for combining the two independent estimates is to average them with appropriate weights (Ernst and Banks 2002), where the weights are inversely proportional to the presumed underlying noise distribution.

$$D_{PR} = w_P D_P + w_R D_R \tag{2}$$

where D_P and D_R refer, respectively, to the estimates of duration from the physical duration and the regulation tendency, and w to their relevant weighting. The weights are directly proportional to the reliability of the estimates, defined as the inverse of their variance (σ_P^{-2} and σ_P^{-2}), and normalized to sum to unity:

$$w_P = \frac{\sigma_P^{-2}}{\sigma_P^{-2} + \sigma_R^{-2}}, w_R = \frac{\sigma_R^{-2}}{\sigma_P^{-2} + \sigma_R^{-2}} \tag{3}$$

This was the procedure used to fit the effectiveness of the distractors (continuous lines in Fig. 3) from the individual Weber fractions (Fig. 6). The estimates of the root variance of the regularization process and the goodness of the fits are given in Table 1. We also used the model to predict the data with trailing distractors, shown by the curves of Fig. 4.

Table 1 Values of the prior of the regularization process (expressed as a proportion of the duration of the standard) that best fitted the data with a simple Bayesian-like model that combined regularization with a direct estimate of the duration (see "Modelling the contextual effects" section)

Subject	Leading		Trailing		Ratio
	Prior	R^2	Prior	R^2	
Audition					
EDR	0.27	0.90	0.44	0.86	1.63
FM	0.18	0.87	0.19	0.76	1.05
MM	0.18	0.70	0.25	0.31	1.39
NG	0.27	0.97			
Vision					
EDR	0.22	0.93	0.53	0.70	2.41
FM	0.37	0.74	0.28	0.80	0.76
MM	0.14	0.78	0.44	0.51	3.14
EP	0.27	0.97			
Haptic					
EDR	0.34	0.83	0.45	0.91	1.32
EP	0.33	0.82			
SA	0.42	0.77			
Average	0.27	0.84	0.37	0.69	1.67

R^2 shows the proportion of variance explained by the fit (which had one degree of freedom). Where data were available, this was also calculated for the condition where the distractors followed the test. The priors for the trailing condition were consistently and significantly greater than those for the leading condition (paired two-tailed test–test of 8 conditions with both values, $t_7 = 2.54$, $p = 0.48$)

Table 1 shows the parameters of this fit: 7 times out of 8 the prior was larger, implying that the regularization process tended to be weaker under these conditions. The difference was statistically significant: paired two-tailed test–test of 8 conditions with both values ($t_7 = 2.54$, $p = 0.48$).

Discussion

The experiments of this study reveal several facts. Firstly, abutting "distractor" stimuli influences the apparent duration of brief test stimuli in a systematic way: short distractors cause the tests to be perceived as shorter, and vice versa. The effect is stronger for leading distractors, but trailing distractors also produce an effect, generally of weaker magnitude. This occurs only for relatively short base times, less than 500 ms for most observers. Distractors of a different modality have no effect on the perceived duration of the test, unless the whole distractor interval is delimited by two markers of the same modality: then, their effect is as strong or stronger than same-modality distractors. We found a strong and consistent relationship between the magnitude of the distractor effect, and the Weber fraction; strong distractor effects were associated with high Weber fractions (poor temporal resolution).

This study is broadly consistent with the results reported by Karmarkar and Buonomano (2007), that distractors interfere with duration discriminations at short but not long time intervals, and extends them in showing that they do not simply "disrupt" perception at these intervals, but affect them in a systematic way. Furthermore, our results suggest that the distractor effect is associated with poor temporal resolution. Karmarkar and Buonomano (2007) noted that distractors influence temporal discriminations at short intervals, providing support for their model of duration estimation, where neural circuits inherently process time within complex networks driven by time-dependent properties of cell membranes (see Buonomano and Merzenich 1995; Buonomano 2000; Buonomano and Karmarkar 2002; Maass et al. 2002). However, although their model does predict interference between neighbouring stimuli, it cannot predict the retrospective effect we observe when the distractor follows the test, as the model is strictly causal, without memory. The data reveal an asymmetry in the effect of the distractors, with leading distractors more effective than trailing, probably reflecting causality in time. However, although there is an asymmetry in the effectiveness in the distractors, the fact that trailing distractors also have an effect clearly violates pure causality, implicating the involvement of memory, and processes akin to "postdiction" (Eagleman and Sejnowski 2000).

The ubiquity of Weber law has often been considered a hallmark of interval discrimination performance (e.g. Gibbon et al. 1997). However, under the conditions of our study, Weber's law was seriously violated, for auditory, visual and haptic discriminations (Fig. 6). Over the 1.5 log unit range of reference intervals, the Weber fractions varied by about 0.75 log unit. This is consistent with a square-root relationship (shown by dashed lines in Fig. 6), halfway between constant Weber (zero dependency on duration) and constant JND (linear dependency on duration). This square-root relationship is not uncommon for duration discriminations over this range of intervals (e.g. Burr et al. 2009b; Lewis and Miall 2009). And in fact, many studies show that Weber's law is seriously violated at low durations. For instance, in their review championing the scalar property of interval timing, Gibbon et al. (1997) report a meta-analysis of Weber fractions for human and animal studies over the range of milliseconds to hours (their Fig. 3). Although there exist interval ranges where the Weber fractions tend to be quite flat, over the interval of this study (30–1,000 ms), the data tend to follow the inverse square-root law (slope of −45° on their plot), consistent with our results. Similarly, Lewis and Miall (2009) report a progressive decrease in Weber fractions over a very large range of intervals.

Importantly, the magnitude of the distractor effect correlated well with Weber fraction, explaining 66 % of the variance: the lower the Weber fraction, the less effect the distractors had on perceived duration. This suggests that the reason that distractors are effective only at short, but not long durations may be that interval timing is more precise at long durations and therefore less influenced by the distractor intervals. We were able to simulate this effect with a Bayesian-like model that combined the direct estimate of duration with that of a distorting process, which we termed "regularization", and show quantitatively how the influence of a process of this type would be stronger at low than at high intervals, given the different Weber fractions. The fits, with only one degree of freedom (the Weber fraction of this regularization process), accounted on average for 84 % of the variance of the data (see Table 1).

Our results are broadly consistent with the previous literature of contextual temporal effects, such as the "time-shrinking" illusion and "entrainment effects" (Monahan and Hirsh 1990; Nakajima et al. 1991, 2004; McAuley and Jones 2003; Jones and McAuley 2005). All these studies report assimilation-like effects of irrelevant distractors (either singletons or stimulus trains) on apparent duration of test intervals, effects that are stronger at short rather than long intervals. Some differences are apparent. For example, Nakajima and colleagues (Nakajima et al. 1991; Miyauchi and Nakajima 2005) report that short distractors reduce apparent duration more than long distractors increase it (reflected in their description of "time-shrinkage"), whereas we report symmetrical effects both for short and long distractors. Similarly, whereas we show that the effect can occur for both preceding and following distractors, Nakajima et al. (1991) report that only preceding distractors produced the effects. The reasons for these small differences in results are not clear, but affect very little the thrust of the results, or the modelling.

Our results are well described by a Bayesian model in which perceived duration is biased towards perceiving a regular sequence of tones or flashes of equal duration. The regularisation mechanism competes with the measurement of physical duration (the likelihood), which depends on Weber fraction. For this reason, the strength of regularization depends directly on Weber fraction, and like the Weber fraction, it must decrease with base interval duration. Our Bayesian model is similar in principle to one recently proposed by Sawai et al. (2012). They also show a prior that biases perception towards equal intervals can reproduce the "time-shrinking" illusion. Their model has far more parameters than ours, mainly in order to model the strong effect towards "shrinking" rather than expansion (Nakajima et al. 1991), which we did not attempt to model, as we did not replicate this effect. Furthermore, they assume a fixed resolution of 25 ms for the auditory system, rather than incorporating measured estimates of Weber fraction. Nevertheless, the model is very similar in its intention, in showing how Bayes' rule can be applied to combine an estimate of duration with a bias towards regularity.

We prefer not to speculate on the exact nature of the regularization process, but point out that it is very similar in principle to the notion of "entrainment", the idea of an inner rhythm tending towards regularity, usually in the context of music (McAuley and Jones 2003; Jones and McAuley 2005). Our results suggest that entrainment comes into play with very brief sequences of sounds, as few as three. It also suggests that entrainment could work in both directions, where both the leading and the trailing stimuli are modified in the direction of regularity. This idea also finds general agreement with the suggestion of Nakajima et al. (1991, 2004) that the "time-shrinking" illusion results from a tendency to perceive the subjective ratio between adjacent intervals as a simple integer ratio, in this case 1:1. What we add in this paper is a simple mechanism whereby the tendency towards regularity can interact with physical estimates of duration, to result in the systematic pattern of distorted perception that we report here.

By what mechanism do distractors affect apparent duration of short intervals? The effect is reminiscent of the ubiquitous "regression to the mean", observed in all modalities (Hollingworth 1910), recently described for interval judgements within a Bayesian framework (Jazayeri and Shadlen 2010; Cicchini et al. 2012). Regression to the mean, or "central tendency", is the tendency of observers to underestimate long intervals and overestimate short intervals, so all are towards the mean of the intervals observed. The distractors would affect the "mean" in the correct direction for this to be applicable here. However, the effects are far too strong to be considered a regression towards the mean. At short intervals, the test can appear as short as the distractor itself, far exceeding the mean. However, it is conceivable that active mechanisms in temporal processing cause resonation or similar phenomena, which lead to the regularization we suggest.

We prefer at this stage not to delve too deeply into the possible mechanisms driving this phenomenon, given the paucity of our understanding of timing mechanisms in general. Nevertheless, the results reported here should constrain any model of duration perception. The first fact to explain is why distractors influence duration in a positive way (assimilation rather than contrast). The second fact is that the effectiveness of the distractors co-varies with Weber fraction. The third is that both leading and trailing distractors influence the test, but the leading distractors more so. Finally, the cross-modal results are important. A single auditory pulse does not affect a visual judgement and vice versa, suggesting that the distractor effect is operating within the visual or auditory modalities. However, if the distractor is delineated by two markers of the same modality, it affects the apparent duration of the test interval of the other modality. This suggests that either the effect occurs at a level after combination of visual and auditory signals or that there is feedback from this area influencing duration judgements.

This study shows that the perception of event duration, particularly of short intervals, deviates heavily from veridicality, being strongly influenced by distractors. We show how the degree of the influence is well predicted by the precision of the duration judgements (Weber fractions) and show how this explains why the influence is stronger at shorter than at longer intervals. We do not propose a specific mechanism of how the distractors influence so strongly event duration, but do suggest that these results strongly constrain any model of duration perception.

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