Reading English with Japanese in mind: Effects of frequency, phonology, and meaning in different-script bilinguals *

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Keywords: bilingual processing; language non-selective activation; lexical decision; eye movements; Japanese-English bilinguals

Abstract

Previous priming studies suggest that, even for bilinguals of languages with different scripts, non-selective lexical activation arises. This lexical decision eye-tracking study examined contributions of frequency, phonology, and meaning of L1 Japanese words on L2 English word lexical decision processes, using mixed-effects regression modeling. The response times and eye fixation durations of late bilinguals were co-determined by L1 Japanese word frequency and cross-language phonological and semantic similarities, but not by a dichotomous factor encoding cognate status. These effects were not observed for native monolingual readers and were confirmed to be genuine bilingual effects. The results are discussed based on the Bilingual Interactive Activation model (BIA+, Dijkstra & van Heuven, 2002) under the straightforward assumption that English letter units do not project onto Japanese word units.

Keywords: bilingual processing; language non-selective activation; lexical decision; eye movements; Japanese-English bilinguals

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While a business person enjoys a cup of *coffee* to start a day in New York, a university professor in Amsterdam lectures on how *koffie* stimulates neural activity, and a student with drowsy eyes in Tokyo sips on another cup of コーヒー to concentrate on his homework. The words *coffee* /kʰɔfi/, *koffie* /kɔfi/, and コーヒー /koohii/ (with the double vowels representing a moraic long vowel) are examples of cognates. These are word pairs with a significant degree of semantic, orthographic and/or phonological form overlaps across languages, which often reflects a cross-linguistic historical link or lexical borrowing. Although it is easy to see that cross-language semantic similarity motivates orthographic and phonological resemblance (e.g., *coffee* in English, *koffie* in Dutch, and *caffè* in Italian), the orthographic similarities of the words in two languages are not always guaranteed (e.g., *coffee* in English alphabet, コーヒー /koohii/ in Japanese *katakana*, and 咖啡 /kafei/ in Chinese *hanzi*, with phonological resemblance maintained across these languages).

When comparing typologically different languages, lexical borrowings are the main source of cross-language phonological similarity. In Japanese, lexical borrowing is ubiquitous and on-going, and the presence of *katakana* scripts makes it possible for any foreign word to be absorbed into the Japanese lexicon irrespective of whether it is a frequent word across cultures (e.g., ウォーター for water) or a proper noun (e.g., サリー for Sally).[1]

The large number of borrowings in Japanese provides a unique opportunity for investigating how the specific characteristics of English and Japanese writing systems affect bilingual visual word recognition. Resolving this issue is important for the characterization of the human language-processing architecture, which at its most abstract level may arguably be language-independent. We addressed this issue by investigating to what extent *katakana* word knowledge is activated when Japanese-English bilinguals perform a visual lexical decision task on English words.

Visual word processing in bilinguals of languages with different scripts

Some initial studies proposed that reading a word in one language might lead to a restricted activation of words only in that language (the so-called 'language-selective access view', Gerald & Scarborough, 1989; Rodriguez-Fornells, Rotte, Heinze, Nösselt, & Münte, 2002; Scarborough, Gerard, & Cortese, 1984). However, the majority of experimental studies indicate that a presented visual word input leads to activation of word candidates in both languages (the so-called 'language non-selective access view', Dijkstra & van Heuven, 1998, 2002; van Heuven, Dijkstra, & Grainger, 1998).

A direct consequence of such a language non-selective access process is that cognates, due to representations linked in memory, are processed more quickly. Such cognate facilitation effects have been reported for a variety of experimental tasks, such as word association (e.g., van Hell & de Groot, 1998), word naming (de Groot, Borgwaldt, Bos, & van den Eijnden, 2002), picture naming (Hoshino & Kroll, 2008; Kohnert, 2004), sentence reading with eye-tracking (Duyck, van Assche, Drieghe, & Hartsuiker, 2007; van Assche, Duyck, Hartsuiker, & Diependaele, 2009; van Assche, Drieghe, Duyck, Welvaert, & Hartsuiker, 2011), and vocabulary learning (Otwinowska-Kasztelanic, 2009). Collectively, these studies indicate that the lexical processing architecture of bilingual readers utilizes lexical distributional properties of two languages in lexical memory even when processing words in one language.

Perhaps surprisingly, there is a growing amount of data supporting exhaustive cross-language lexical activation even in bilinguals of languages with different scripts. Gollan, Forster, and Frost (1997), Nakayama, Hino, Sears, and Lupker (2012), and Kim and Davis (2003) tested Hebrew-English bilinguals, Japanese-English bilinguals, and Korean-English bilinguals respectively with a masked cross-script priming paradigm and reported that a word in one language still activated its phonologically and semantically related cognate in another orthographically distinct language.

Note that the observation of language non-selective lexical activation in bilinguals of languages with different scripts does not imply that the underlying lexical processing architecture is the same as in bilinguals of languages with identical scripts. In fact, given differences in orthography, there must be some differences in the organization of the first stages of visual word processing for the two languages involved. At more abstract levels, however, the underlying processing mechanisms are likely to be similar. For instance, the role and interaction of lexical-phonological and semantic information sources within the lexicon might be analogous in Japanese-English and Dutch-English bilinguals. The present study aims to clarify what mechanisms remain the same and what must be different to account for lexical processing across languages with different scripts. In this study, we will consider this issue from the theoretical perspective of a localist connectionist framework.

Extending the BIA+ model to languages with different scripts

The Bilingual Interactive Activation (BIA) model (Dijkstra & van Heuven, 1998, 2002; Dijkstra, van Heuven, & Grainger, 1998) is a localist connectionist model that extends the monolingual IA model (McClelland & Rumelhart, 1981) and allows us to conceptualize monolingual and bilingual lexical processes within one theoretical framework. While the original BIA model was limited to orthographic aspects (Dijkstra & van Heuven, 1998; Dijkstra, van Heuven, & Grainger, 1998), it has been extended to account for experimental evidence on cross-language phonological and semantic activation (the BIA+ model, Dijkstra & van Heuven, 2002). In addition, the BIA+ model attempts to account for cross-task variations by incorporating a task/decision system explicitly in the model architecture. Currently, this extra-linguistic system is not expected to immediately affect lexical activation in the word identification system, based on the experimental evidence that bilinguals automatically activate two languages or word representations even when this language non-selective activation is not necessary in the task and, in fact, can even be detrimental (Dijkstra, de Bruijn, Schriefers, & ten Brinke, 2000; Dijkstra & van Hell, 2003; Dijkstra, van Jaarsveld, & ten Brinke, 1998; van Assche et al., 2011; van Hell & Dijkstra, 2002).

For languages sharing the same script (e.g., Dutch-English and French-English bilinguals with Latin alphabets), the BIA and BIA+ models predict that orthographic features of the input word immediately activate orthographic lexical representations in the two languages simultaneously. In contrast, for languages with different scripts, identification of script-specific orthographic features is not expected to facilitate activation of words in both languages. When the BIA+ model is generalized to languages with different scripts, the model may provide different predictions. In the example shown in Figure 1, the input word *interview* activates the corresponding letter nodes I, N, T, E, R, V, W, and these letter features then activate the word node *interview*. For Japanese-English bilinguals, however, it is expected that the feature set coding visuoperceptual features of Latin alphabets does not directly encode Japanese *katakana* script (the dotted line (a) in Figure 1). Consequently, the orthography-driven non-selective lexical access across languages is not expected for bilinguals of languages with different scripts at the earliest processing stages. The section below summarizes our hypotheses for Japanese-English bilinguals with respect to the BIA+ architecture, together with diagnostic variables used to test the predictions in the following lexical decision with eye-tracking experiments (Table 1). Appendix A provides detailed descriptions about the lexical predictors, individual differences, and task effects considered in the present study. The lexical predictors are classified into those specific to Japanese-English bilinguals and those of English target words. All predictors were centered for the regression analyses. In order to study independent contributions of lexical distributional properties, we opted for a residulization procedure to orthogonalize correlated variables, as in Kuperman et al. (2009) and Miwa et al. (2013). Residualized variables are indicated by the suffix *_resid* in Table 1 (see Appendix B for the correlational structure among Japanese and English frequencies, and Appendix C for the details on the residualization procedure).

< Insert Figure 1 and Table 1 around here >

Cross-language phonological similarity. Phonological effects in visual word recognition have been studied predominantly in monolingual word recognition research (Carreiras, Ferrand, Grainger, & Perea, 2005; Ferrand & Grainger, 1992, 1994; Perfetti et al., 1992), but there is growing evidence for cross-language phonological activation in bilingual visual word recognition as well (Brysbaert, 2003; Brysbaert, van Dyck, & van de Poel, 1999; Dijkstra et al., 1999; Duyck, 2005; Duyck, Diependaele, Drieghe, & Brysbaert, 2004; Schwartz, Kroll, & Diaz, 2007).

For Japanese-English bilinguals, if activation of orthographic lexical representations of L1 words is mediated only by the conceptual route, then a cross-language phonological similarity effect should appear late in time. However, phonology-driven sublexical language non-selective access is theoretically still possible (the route to the box (d) in Figure 1). In the latter case, an effect of cross-language phonological similarity may appear early. For Japanese-English bilinguals, it is expected that the activation of English phonemes leads to activation of the corresponding Japanese phonemes and syllabic and/or moraic phonological nodes (e.g., the activation of phonemes /i/, /n/, /t/, /a/, /b/, /j/, and /u/ facilitates the activation of syllabic representations or moraic representations of /i/, /nn/, /ta/, /bju/, and /u/). The phonological similarity between English and Japanese may lead to a larger global activation in the lexicon, just like a greater degree of orthographic similarity between L1 and L2 words matters for languages with the same script. As a diagnostic measure of phonological similarity, we used rated phonological similarity (*PhonologicalSimilarityJPN)*.

Relative word frequencies in two languages. Like the monolingual IA model, the BIA+ model maintains inhibitory connections among words in the target language and also assumes inhibitory connections between orthographic lexical representations in two languages (Figure 1, line c). One consequence of such inhibitory connections is that the magnitude of the expected facilitatory effect of target English word frequency will be smaller when the Japanese translation equivalent has a high frequency of occurrence, because an English lexical orthographic representation and the Japanese lexical orthographic representation become co-active at some point in the course of word identification. It has been reported that word frequencies in two languages interact for interlingual homographs, words that share orthography across languages but not meaning (Dijkstra et al., 2000; Dijkstra, Moscoso del Prado Martin, Schulpen, Baayen, & Schreuder, 2005; Dijkstra, van Jaarsveld, & ten Brinke, 1998; Kerkhofs, Dijkstra, Chwilla, & de Bruijn, 2006).

The BIA+ model predicts that, for Japanese-English bilinguals, any orthographic cross-language inhibition can only occur later in the recognition process, because the Japanese orthographic representation becomes activated only by virtue of cross-language phonological or conceptual mediation. We used log-transformed English word frequency (*FreqHAL*, Balota et al., 2007) and log-transformed Japanese word frequency (*FreqJPN*, Amano & Kondo, 2003) as diagnostic measures of strength of the activation of lexical orthographic representations to test this prediction.

Semantic similarity. Translation equivalents in two languages occasionally have different shades of meaning. For example, unlike the English word *interview* which is used unrestrictedly in various contexts, the use of the Japanese translation equivalent インタビュー is restricted to 'mass media interviews' and not typically used for 'job interviews.' A question relevant to bilingual word processing is whether such cross-language semantic similarity contributes to recognition of L2 words. In an unprimed

English progressive demasking task with Dutch-English bilinguals, Dijkstra et al. (2010) reported a processing advantage for English words with higher semantic similarity. As shown in Figure 1 box (e), we predicted that semantic similarity contributes relatively late in the course of word recognition. We used rated semantic similarity as the diagnostic measure to test this prediction (*SemanticSimilarity*).

Cognate status. While the BIA and BIA+ models account for cognate facilitation effects by means of cross-language orthographic, phonological, and semantic overlaps alone, it has been suggested that a shared morpheme representation for cognate may underlie the effects (Davis et al., 2010; Lalor & Kirsner, 2000; Sánchez-Casas & García-Albea, 2005). The present study considers cognate effects as orchestration of gradient lexical effects but also considers a dichotomized factor coding cognate status to test the special representation view. If cognates have shared morpheme representations, then a factor encoding cognate status should emerge significant on top of relevant numerical predictors encoding orthography, phonology, and semantics. We straightforwardly tested this theoretical prediction with a factor *Cognate* (levels: *Cognate* and *NotCognate*) in a regression model.

Extra-linguistic task/decision processes. Although the BIA and BIA+ models have been frequently discussed together in relation to language non-selective lexical access, the latter is distinguished from the former with respect to its explicit consideration of a non-linguistic system co-determining responses in a given task. The BIA+ model currently assumes that the non-linguistic system does not modulate lexical processes at the earliest processing stages. This assumption will be falsified if response-based data and data sampled from the earliest time frame both reveal the same interactions between lexical and task-related variables. To the best of our knowledge, however, no study has identified such interactions. In order to fully test the BIA+ model, we tracked participants' global and local response criteria by studying extra-linguistic variables: *Trial*, the number of preceding trials, and *PreviousResponseCorrect*, whether the responses in the preceding two trials were correct.

Control variables

We considered objective phonological similarity based on Levenshtein distance (*PhonologicalDistance,* Levenshtein, 1966; Gooskens & Heeringa, 2004; Schepens, Dijkstra, & Grootjen, 2011; Schepens, Dijkstra, Grootjen, & van Heuven, 2013) to check the validity of rated *PhonologicalSimilarity*. A log-transformed Google document frequency measure (*GoogleFreqJPN*) was considered to check the validity of *FreqJPN* because the latter newspaper-based measure contains zero frequencies for some words.

As lexical control predictors, word length (*Length*), orthographic Levenshtein distance (*OLD20*, Yarkoni et al., 2008), log-transformed context diversity (*SUBTLCD*, Brysbaert & New, 2009a, 2009b), and rated *Imageability* were considered.

As task-related variables, we considered *PreviousRT*, inversely transformed RT in the previous trial, and *PreviousFixationDuration* for second fixation duration analyses to account for potential spillover effects from the previous fixation.

To safe-guard against potential individual differences (Kroll & Stewart, 1994; Potter, So, von Eckhardt, & Feldman, 1984), we considered log-transformed participants' months of stay away from Japan (*LengthOfStayCanada*) as a measure of L2 English proficiency.

Methodological concerns and goals of this study

Previous priming studies for bilinguals with different scripts have provided evidence supporting automatic language non-selective lexical activation within an integrated lexicon, as implemented in the BIA model. However, when Japanese-English readers encounter an English word, Japanese lexical orthographic representation is not yet activated due to the orthographic dissimilarity (Figure 1 line a). It is important to note that, in the context of cross-script priming, a lexical orthographic representation of one language (Figure 1 box b) is artificially pre-activated. This raises a concern as to what extent language non-selective activation holds in a task without priming (as exceptions, see Thierry & Wu, 2004, 2007 and Wu & Thierry, 2010 for implicit priming to study bilinguals of languages with different scripts). While masked priming is one of the most popular techniques to test subconscious lexical processes, researchers have not reached a consensus on an interpretation of a masked priming effect (e.g., see Forster, 1998 for a lexical pre-activation account, Kinoshita & Norris, 2010 for a non-lexical account, and Marsolek, 2008 for an antipriming account). At the moment, studies without priming for bilinguals with different scripts are scarce, and we investigate this issue using eye-tracking.

In the present study, combining English lexical decision tasks with eye-tracking, we tested the above predictions of the extended BIA+ model with Japanese-English bilinguals (Experiment 1) and with native monolingual readers (Experiment 2). Lexical decision was chosen, as this has been the most widely used experimental task (Libben & Jarema, 2002). Eye-tracking was used because previous studies employing lexical decision with eye-tracking (Kuperman, Schreuder, Bertram, & Baayen, 2009; Miwa, Libben, Dijkstra, & Baayen, 2013) reported that early and late lexical processes systematically co-determine the initial and late eye-fixation measures respectively.

Although the vast majority of psycholinguistic studies on bilingual processing have analyzed bilingual-specific effects solely in bilingual readers, we also compared late bilinguals and native monolinguals to (1) make sure that effects of interest are genuine bilingual effects, rather than artifacts arising from statistical correlation across languages and general processing mechanisms in reading English words, (2) to confirm that there is a reasonable amount of functional overlap between bilinguals and monolinguals because part of the BIA/BIA+ model architecture, in theory, accounts for monolingual readers' lexical processing, and (3) to explore the 'expert' ability in reading, as native proficiency also provides a benchmark of an expert reader, and the acquisition of such 'expert' ability is often viewed as a goal of late bilinguals.

In order to test the above predictions of the BIA+ model, we opted for a mixed-effects analysis (Baayen, Davidson, & Bates, 2008) without dichotomization of numerical predictors for more power and precision (Baayen & Milin, 2010; Cohen, 1983; MacCallum, Zhang, Preacher, & Rucker, 2002). Mixed-effects modeling allows us to test lexical distributional properties, participants' characteristics, and task-related variables in a single statistical model.

Experiment 1: English lexical decision with Japanese-English bilingual readers

Method

Participants. Nineteen Japanese-English late-bilingual readers (three males, mean age = 25.1, *SD* = 5.8) were recruited at the University of Alberta. The participants had stayed in Canada for 33 months on average (*SD* = 45.6) and acquired English as their second language.

Materials. We sampled 250 words from English Lexicon Project database (Balota et al., 2007) based on the following criteria: (1) the word length was between 6 and 9 letters; (2) the word frequency was greater than 2,000 in the *FreqHAL* frequency distribution; (3) the morphological status was simplex; (4) the part of speech was noun; and (5) the mean accuracy rates in lexical decision and naming were at least 0.9 (see Appendix D for the list of words used in the present study). We also sampled 200 nonwords from the ARC nonword database (Rastle, Harrington, & Coltheart, 2002) to make a total of 450 letter strings. These nonwords were similar to the words: 6 to 9 letters long, with existing onsets, existing bodies, and legal bigrams.

Apparatus. The experiment was designed and controlled by SR Research Experiment Builder software. Words were presented on a 20-inch display. Eye-movements were tracked by an EyeLink II head-mounted eye-tracker (SR Research, Canada) in the pupil-only mode with a sampling rate of 250 Hz. Calibrations were conducted with horizontal three points.

Procedure. In this lexical decision with eye-tracking experiment, participants were asked to decide as quickly and accurately as possible whether the letter strings presented on the computer display were legitimate existing English words or non-existent words (nonwords) by pressing the right and left buttons of a Microsoft SideWinder game-pad respectively. The words were presented following a fixation circle, on which participants were asked to fixate their eyes. This fixation circle served as a drift correct point allowing the researcher to correct for head drifts between trials.

Given our understanding of the optimal viewing position in sentence reading and in isolated word reading (Brysbaert & Nazir, 2005; Farid & Grainger, 1996; O'Regan & Jacobs, 1992; Vitu, O'Regan, & Mittau, 1990), the location of the fixation mark was slightly shifted horizontally so that the first fixations were positioned slightly (1.5 letters) to the left of the word centre. The target words were presented in Courier New 44-point white font on a black background. At a viewing distance of 70 cm from the screen, the visual angle was estimated to be 1.1° for each letter (i.e., for six-letter words and nine-letter words, such as *camera* and *interview*, the visual angles were approximately 6.7º and 10.0º respectively). A testing session contained two breaks, one after every 150 words. Participants saw a summary of their performance regarding accuracy and response speed after the practice trials and at each break point. The experiment took roughly 90 to 120 minutes. The right eye was tracked throughout the experiment.

Results

Response latency. R version 2.13.2 (R Development Core Team, 2011) was used for statistical analyses. We opted for a mixed-effects regression analysis with subjects and items as crossed random effects (Baayen et al., 2008; Bates, Maechler, & Dai, 2007). Response accuracy across subjects ranged from 0.86 to 0.98 (M = 0.92, SD = 0.04). Therefore, no participant was excluded from the analyses. Of all 4,750 trials, those with response latency either shorter than 300 ms or longer than 3,000 ms were excluded (14 trials). Twenty-one words with more than 30% erroneous responses and one word (*heroin*) with a coding error were excluded from the analysis (414 trials, 8.7 % of the remaining data). The following analyses are based on the remaining 228 words, amounting to 4,099 trials after excluding the 5.1% of trials that contained incorrect responses.

A reciprocal transformation (-1000/RT) was applied to the RTs to attenuate skewness in their distribution, based on the appropriate exponent suggested by the Box-Cox power transformation technique (Box & Cox, 1964; Venables & Ripley, 2002). The Box-Cox transformation technique was applied to all dependent measures in the rest of this paper.

We first fitted a simple main effects model with all predictors and also considered all pairwise interactions. We then tested interactions between task-related predictors and lexical predictors. Potentially influential outliers were removed with standardized residuals exceeding 2.5 standard deviation units (1.8% of the data points). Table 2 summarizes the fixed-effects in our final model, and Figure 2 visualizes significant interactions. The reference level for the factor *PreviousResponseCorrect* was *Correct* throughout this study. The random-effect structure of this model consists of random intercepts for item (*SD* = 0.09) and subject (*SD* = 0.16), by-subject random slopes for *Trial* (*SD* = 0.02), and by-subject random contrasts for *PreviousResponseCorrect* (*SD* = 0.06). The standard deviation of the residual error was 0.24. Other random slopes for subjects did not reach significance.

< Insert Table 2 and Figure 2 around here >

Interestingly, all bilingual-specific predictors, except the factor *Cognate*, co-determined the lexical decision response latencies (Figure 2, Panels a, b, and c). Although the effect of *PhonologicalSimilarityJPN* reached significance in a simple main-effects model, a model allowing interactions into the model specification further clarified that *PhonologicalSimilarityJPN* facilitated responses later in the experiment (Panel a). Replacing *PhonologicalSimilarityJPN* with *PhonologicalDistance* successfully replicated this interaction.

The effect of *FreqJPN_resid* emerged in an interaction with target word frequency *FreqHAL* (Panel b). For words with large *FreqJPN_resid* (the solid line in Panel b), the magnitude of the *FreqHAL* effect was attenuated. When *FreqJPN_resid* was replaced with *GoogleFreqJPN_resid*, a virtually identical interaction was obtained.

SemanticSimilarity_resid had a small yet significant facilitatory effect in a simple main-effects model. Upon close inspection, its interaction with the task-related variable, *PreviousRT* (Panel c) indicated that when the response latency in the previous trial was long (the solid line in Panel c), cross-language *SemanticSimilarity_resid* facilitated the response more strongly. This is in line with the finding that the cross-language semantic similarity effect is more likely to be observed when the task induces longer response times (Dijkstra et al., 2010, using a progressive demasking word identification task).

Cognate was not significant in this model, indicating that the variance explained by *Cognate* was absorbed by the numerical predictors.

Several other lexical distributional properties of the English target words co-determined response latencies. The magnitude of the *FreqHAL* facilitatory effect was the greatest for words with large *OLD20_resid*, which are words situated in a sparse orthographic space (the solid line in Figure 2 Panel d). As summarized in Table 2, the context diversity of words gauged by *SUBTLCD_resid* contributed beyond *FreqHAL*, and *Length* inhibited responses.

The effect of *SemanticSimilarity_resid* was not modulated by *LengthOfStayCanada*, suggesting that, for Japanese-English bilinguals, the recruitment of cross-language semantic activation did not vary across readers with different L2 proficiency.

It is likely that the above effects reflected in response times also guide eye-movements. We investigated the time-course of the above effects by studying whether they load onto the early or late fixations, or a combination of both. Japanese-English bilinguals read words with a single fixation only 1% of the time (36% showed two fixations, 39% three fixations, and 17% four fixations, with a mode at three fixations). In order to include as many data points as possible, for all trials with at least two fixations, we analyzed (1) first fixation durations, (2) first subgazes (in this study, sum of all non-final fixations, which were ended by a saccade to a next location) and (3) last fixations as measures of very early processing, relatively early processing, and late processing respectively. The first fixation and subgaze durations were measured from the onset of target word presentation, excluding the fixation duration on the fixation point. It is assumed that the former two measures are less contaminated by conscious lexical decision response strategies than the last fixation, which was ended with a button press.

First fixation duration. For the analysis of the first fixation durations, data points with a first fixation shorter than 100 ms and longer than 850 ms, those before a blink, and those with an incorrect response were excluded from the analysis. A log-transformation was applied to attenuate skewness in the distribution of the fixation durations (*M* = 5.7, *SD* = 0.3, raw median = 280 ms). As in the response time analysis, we tested simple main effects and pairwise interactions, as well as interactions between lexical and task-related predictors. Potentially influential outliers were removed with standardized residuals exceeding 2.5 standard deviation units (2.0% of the data points). The final model is summarized in Table 3. The random-effect structure of this model was comprised of random intercepts for item (*SD* = 0.05) and subject (*SD* = 0.11), and by-subject random slopes for *Trial* (*SD* = 0.01). The standard deviation of the residual error was 0.21.

As expected, the first fixation duration was co-determined by the signature of early bottom-up orthographic processing. The inhibitory effect of *OLD20_resid* indicates that words in sparse orthographic neighbourhood receive a longer first fixation. A word frequency effect (*FreqHAL*) was observed already as well, replicating the early word frequency effects in previous studies (Kuperman et al., 2009; Miwa et al., 2013).

Interestingly, the bilingual-specific predictor *PhonologicalSimilarityJPN* already contributed at this earliest time frame, indicating that a sublexical cross-language phonological decoding route is used (the route to the box d in Figure 1). Unlike its effect in the RT analysis, however, its effect was inhibitory. It should also be noted that the objective *PhonologicalDistance* measure, when replaced with *PhonologicalSimilarityJPN*, did not reach significance.

< Insert Table 3 around here >

First subgaze duration. For the analysis of the first subgaze durations, data points with a first fixation shorter than 100 ms, those before a blink, and those with an incorrect response were excluded from the analysis. A log-transformation was applied to attenuate skewness in the distribution of the subgaze durations (*M* = 6.2, *SD* = 0.4, raw median = 500 ms).

After testing simple main effects, pairwise interactions, and interactions between lexical and task-related predictors, potentially influential outliers were removed with standardized residuals exceeding 2.5 standard deviation units (1.2% of the data points). The final model is summarized in Table 3. The random-effect structure of this model was comprised of random intercepts for item (*SD* = 0.08) and subject (*SD* = 0.16), and by-subject random slopes for *Trial* (*SD* = 0.04). The standard deviation of the residual error was 0.31.

At the first subgaze, some bilingual-specific effects were observed. Whereas at the first fixation, *PhonologicalSimilarityJPN* was inhibitory, at the second fixation, *PhonologicalSimilarityJPN* became facilitatory. It should be noted, however, that *PhonologicalDistance* did not reach significance when replaced with *PhonologicalSimilarityJPN*. Interestingly, the interaction between *FreqHAL* and *FreqJPN_resid* observed in the response time analysis was also observed at the first subgaze in a virtually identical form (Figure 3, Panel a). This interaction was successfully replicated when *FreqJPN_resid* was replaced with *GoogleFreqJPN_resid.*

Cognate and *SemanticSimilarity_resid* were not significant predictors. *LengthOfStayCanada* was not a significant predictor and did not modulate any lexical effects either.

< Insert Figure 3 around here >

Last fixation duration. For the analysis of the last fixation duration, the words excluded in the analysis of response latencies were excluded here as well. Trials with last fixation durations longer than 900 ms, those before or after a blink, and those with an incorrect response were also excluded from the analysis. A square-root transformation was applied to attenuate the skewness in the distribution of the fixation durations (M = 13.5, SD = 4.3, raw median = 188 ms). Unlike the first fixation and subgaze durations, which were terminated by the eye moving to another location in the word, the last fixation duration was terminated by the readers' button-press. Consequently, we expected that the last fixation durations would reflect, in addition to lexical predictors, variables associated with response planning and execution.

We tested simple main effects, pairwise interactions, and tested interactions among task-related predictors and lexical predictors. As in the response time analysis, with all relevant numerical predictors considered in a model, *Cognate* did not contribute significantly to the model fit. The final model for Japanese-English bilinguals' last fixation durations is summarized in Table 3. Figure 4 presents the significant lexical interactions in the model. Potentially influential outliers were removed with standardized residuals exceeding 2.5 standard deviation units (1.6% of the data points). The random-effect structure of the final model was comprised of random intercepts for item (SD = 0.93) and subject (SD = 1.08). The standard deviation of the residual error was 2.97.

Several bilingual-specific effects co-determined the last fixation durations. *FreqJPN_resid* interacted with *PhonologicalSimilarityJPN*. Larger

PhonologialSimilarityJPN shortened the fixation duration for words with high *FreqJPN_resid* (the solid line in Figure 3, Panel b), suggesting that words with higher cross-language *PhonologicalSimilarityJPN* and *FreqJPN* were perceived as more word-like. In a model in which *PhonologicalDistance* was used instead of *PhonologicalSimilarityJPN*, a virtually identical interaction was obtained.

SemanticSimilarity_resid facilitated processing. The magnitude of facilitation was greater for words with low cross-language phonological similarity (Panel c). This interaction may indicate an L1-based response strategy to rely on either phonological similarity or semantic similarity to make a lexical decision response. *SemanticSimilarity_resid* also interacted with *PreviousRT_resid* (Figure 3, Panel d). Recall that the facilitatory effect of *SemanticSimilarity* for words preceded by long *PreviousRT_resid* was also observed in the RT analysis. Since *SemanticSimilarity_resid* was absent both at the first fixation and subgaze, we can conclude that this semantic effect emerges late.

The last fixation seems to be qualitatively different from the first fixation and subgaze, as indicated by the atypical inhibitory effects of *FreqHAL* (20 ms) and *SUBTLCD_resid* (27 ms). The inhibitions from L2 word properties may be due to a response strategy to rely on L1 word properties. Involvement of a conscious response strategy is evident from the significant effect of *PreviousResponseCorrect*. When participants make an error, they usually become aware of it immediately after a button press and try to be cautious in the following trials. *LengthOfStayCanada* did not reach significance and did not modulate lexical effects either.

Discussion

Lexical decision measures emerged, from our analyses, as a composite measure amalgamating processing costs that arise at different stages of information uptake. First fixations reflected early bottom-up processing as witnessed by orthographic neighbourhood density and target word frequency. First subgazes reflected lexical effects in the word identification system not affected by conscious response strategies, followed by last fixations, which were more dedicated to response planning and execution in the task/decision system.

Importantly, using a regression technique, we observed all the expected bilingual-specific effects in the reaction times (i.e., phonological similarity effect, interaction between L1 and L2 frequencies, and semantic similarity effect) simultaneously in a task without priming. Interestingly, these effects also co-determined eye-movements but at different points in time. The early contribution of cross-language phonological similarity indicates that, even for languages with different scripts and in a task without priming, sublexical phonological decoding immediately takes place. It should be noted, however, that its effect was inhibitory in the earliest time frame and facilitatory in later time frames.

It was also notable that the interaction between L1 and L2 frequencies was found relatively early at the first subgaze. Under the assumption that response planning and execution takes approximately 200 ms (Schmidt, 1982), it is likely that the cross-language competition was not due to readers' conscious response strategy but rather part of central lexical processing mechanism, as assumed in the BIA+ model.

The time-course of lexical activation characterized by the relatively early contribution of cross-language phonological similarity, followed by the competition between L1 and L2 words, and then by a cross-language semantic similarity effect is in line with the predictions of the BIA+ model.

In Experiment 2, we tested monolingual readers of English to ensure that the bilingual-specific effects observed in Experiment 1 arose from a bilingual-specific processing mechanism, as well as to explore whether the within-language English lexical distributional properties are utilized similarly by the two groups of readers.

Experiment 2: English lexical decision with English monolingual readers

Method

Participants. Nineteen monolingual English readers (7 males, mean age = 21.6, *SD* = 7.1) were recruited at the University of Alberta. There was no significant difference between the late bilinguals in Experiment 1 and the monolingual readers with respect to age. Monolingual readers were defined here as native readers of English with more than 80% daily exposure to English relative to the amount of exposure to their second languages at the time of the experiment, as reported by the participants. None of the participants had Japanese as their second or third language.

Materials, Apparatus, and Procedure. The same as in Experiment 1.

Results

Response latency. In the analyses of Experiment 2, we excluded the words which elicited higher error rates in Experiment 1 to ensure that the comparisons of Experiment 1 and 2 are based on the same set of words. Response accuracy rate ranged from 0.96 to

1.00 (mean = 0.99, *SD* = 0.01) for English monolingual readers. Therefore, no subject was excluded from the analyses. Of all trials (4,750 data points), data points with response latency shorter than 300 ms or longer than 2,000 ms were excluded (11 data points). The following analyses are based on the same 228 words analyzed in Experiment 1 with correct responses (4,275 data points, after excluding 1.1% of trials with incorrect responses). A reciprocal transformation (-1000/RT) was applied to RTs to attenuate the skew in its distribution.

< Insert Table 4 and Figure 4 around here >

We tested simple main effects, all pairwise interactions, and finally interactions between task effects and lexical effects. Potentially influential outliers were removed with standardized residuals exceeding 2.5 standard deviation units (1.9% of the data points). Table 4 summarizes the coefficients of the final model. The random-effect structure of this model was comprised of random intercepts for item (*SD* = 0.07) and subject (*SD* = 0.14), and by-subject random slopes for *Trial* (*SD* = 0.03) and *PreviousRT_resid* (*SD* = 0.07). The standard deviation of the residual error was 0.25.

There were qualitative differences between bilinguals and monolinguals. A three-way interaction of *FreqHAL* by *OLD20_resid* by *Trial* was observed. As illustrated in Panels a and b of Figure 4, *FreqHAL* provided facilitation without interacting with *OLD20_resid* early in the experiment (*Trial* was dichotomized only for the purpose of visualization in Figure 4). However, later in the experiment, the interaction between *FreqHAL* and *OLD20_resid* emerged (Panel b), as observed for Japanese-English bilinguals in Experiment 1. Unlike Experiment 1, *Imageability_resid* co-determined response times as well, such that the responses were delayed for words with low imageability (Balota et al., 2004).

As reported by previous bilingual processing studies (e.g., Duyck et al., 2008; Gollan et al., 2008), the magnitude of the English word frequency effect was larger for Japanese-English late bilinguals than for English monolingual readers. This larger frequency effect for non-native readers can be interpreted by the negative decelerating functional form of the word frequency effect, indicating that each additional log unit of frequency provides smaller and smaller processing benefits (Baayen, Feldman, & Schreuder, 2006; Baayen & Milin, 2010; Duyck et al., 2008). Overall, monolingual English readers' response latencies (M = 546 ms, SD = 138) in Experiment 2 were faster than those of Japanese-English bilingual readers (M = 733 ms, SD = 248) in Experiment 1 ($p < 0.0001$, mixed-effects model not shown).

A direct quantitative comparison between monolinguals and bilinguals was also conducted by including a factor *FirstLanguage* (levels: *Japanese*, *English*) in a regression model for all data. The results indicate that supposedly bilingual effects (*Freq_HAL* * *FreqJPN_resid*, *PhonologicalSimilarityJPN*, and *SemanticSimilarity_resid*) did not reach significance for the English monolinguals (See Appendix E for the analysis).

First fixation duration. Like bilingual speakers in Experiment 1, native English speakers read words with multiple fixations most of the time (8% showed a single fixation, 66% two fixations, 23% three fixations, and 2% four fixations). We therefore analyzed first fixation durations, first subgaze durations, and last fixation durations, as in Experiment 1.

For the analysis of the first fixation durations, data points with a first fixation shorter than 100 ms and longer than 850 ms, those before a blink, and those with an incorrect response were excluded from the analysis. A log-transformation was applied to attenuate skewness in the distribution of the fixation durations (M = 5.6, SD = 0.2, raw median = 264 ms).

The final model is summarized in Table 5. Potentially influential outliers were removed with as standardized residuals exceeding 2.5 standard deviation units (2.7% of the data points). The random-effect structure of this model was comprised of random intercepts for item (SD = 0.03) and subject (SD = 0.11), and by-subject random slopes for *Trial* (SD = 0.01) and *Length* (SD = 0.03). The standard deviation of the residual error was 0.16. For the expert readers, *Length* was the only significant lexical predictor. An item-wise correlation between the first fixation durations of Japanese-English bilingual readers (Experiment 1) and those of English monolingual readers was significant but weak (r = 0.19, p < 0.01). The first fixations of the monolinguals were not significantly faster than those of the bilinguals.

< Insert Table 5 around here >

First subgaze duration. Data points before a blink and trials with incorrect responses were excluded from the analysis. Trials with a first subgaze duration shorter than 100 ms were excluded, and an inverse transformation was applied to the remaining fixation durations to attenuate skewness in the distribution (M = -3.2, SD = 0.9, raw median = 308).

Table 5 summarizes the final model, after testing all main effects and pairwise interactions. Potentially influential outliers were removed with standardized residuals exceeding 2.5 standard deviation units (1.3% of the data points). The random-effect structure of this model was comprised of random intercepts for item (*SD* = 0.14) and subject (*SD* = 0.4), and by-subject random slopes for *Trial* (*SD* = 0.07). The standard deviation of the residual error was 0.73. The fixed-effect structure of this model was comprised of *Length*, *FreqHAL*, and *SUBTLCD_reisd*, all of which co-determined Japanese-English bilinguals' first subgaze in Experiment 1. A significant item-wise correlation between the first subgaze durations of Japanese-English bilingual readers (Experiment 1) and those of English monolingual readers ($r = 0.46$, $p < 0.01$) also indicates strong commonality in information uptake during the relatively early stage of word recognition. The first subgaze durations of the monolingual readers were significantly faster than those of the bilingual readers ($p < 0.0001$, effect size = 163 ms, mixed effects model not shown).

Last fixation duration. In the same subset of words analyzed above, we excluded last fixation durations longer than 900 ms and applied a square root transformation to attenuate a skew in the distribution (*M* = 14.2, *SD* = 3.7, raw median = 216). Fixations before or after a blink and trials with incorrect responses were excluded from the analysis.

The final model is summarized in Table 5. Potentially influential outliers were removed with standardized residuals exceeding 2.5 standard deviation units (1.9% of the data points). The random-effect structure of this model was comprised of random intercepts for item (*SD* = 0.81) and subject (*SD* = 0.75), and by-subject random slopes for *FirstSubgazeDuration_resid* (*SD* = 0.61). The standard deviation of the residual error was 2.35.

The *Imageability_resid* was the only lexical predictor co-determining monolingual English readers' last fixation durations. Recall that *Imageability_resid* also co-determined the same readers' response times to a comparable magnitude. Interestingly, as in Experiment 1, *PreviousResponseCorrect* was significant here as well. When participants make an error response (which they are usually aware of), their immediately following responses become slower to be more cautious. Such strategic effects do not seem to co-determine early measures and only inflated last fixation durations.

An item-wise correlation between the last fixation durations of Japanese-English bilingual readers (Experiment 1) and those of English monolingual readers was significant but weak ($r = 0.29$, $p < 0.01$). Furthermore, the last fixations of the monolingual English readers were not significantly faster than those of the late bilinguals.

Discussion

Experiment 2 confirmed that there was commonality in processing between monolingual readers and bilingual readers: The lexical processes proceed from visuo-perceptual and sublexical orthographic effects (*Length* and *OLD20_resid*) to orthographic lexical effects (*FreqHAL* and *SUBTLCD_resid*) and then to semantic processes (*Imageability_resid* and *SemanticSimilarity_resid*). Importantly, however, the bilingual-specific effects obtained in Experiment 1 did not reach significance in Experiment 2. This indicates that the bilingual-specific effects of our interest were genuine bilingual effects arising from the theoretical bilingual-specific architecture rather than artifacts arising from processes of target words per se. The functional overlap across different groups of readers and the significance of bilingual-specific effects only for bilingual readers are in line with the BIA and BIA+ models.

However, it should be noted that Experiment 2 also identified differences between monolingual readers and late bilinguals. *Length* was a significant predictor at the first fixation only for the monolinguals. *Imageability_resid* was similarly a significant predictor at the last fixation durations only for the monolinguals.

Finally, the significant effects of *PreviousResponseCorrect* on last fixation durations and response times, but not first fixations and subgaze durations, of both monolingual readers and bilingual readers suggest that this is a language-general phenomenon and that readers' conscious strategy to respond more cautiously affects only the late processes. This, in turn, indicates that the first fixations and subgazes are relatively strategy-free measures for automatic lexical processing. The late involvement of a response strategy and the word identification system's insensitivity to a strategic factor is in line with the prediction of the BIA+ model.

General discussion

This study addressed the question of the time-course of lexical activation in Japanese-English bilinguals. To this end, we combined lexical decision with eye-tracking because eye-tracking, unlike button press responses, affords insight into the time-course of lexical activation. Importantly, because the task did not involve cross-script priming, the lexical orthographic representation (Figure 1, box b) was not artificially activated when the L2 target word was presented, allowing a more natural interpretation of the results. As predicted by the BIA+ model, we observed clear effects of all the bilingual-specific lexical predictors (cross-language phonological similarity, L1 word frequency interacting with L2 word frequency, and cross-language semantic similarity), but not the factor coding cognate status.

First, cross-language phonological similarity facilitated the lexical decision responses of Japanese-English bilinguals. The eye-tracking record clarified that the phonological similarity effect emerged already at the first fixation. Within the framework of BIA+ model, we interpret this effect as sketched in Figure 1 box (d). Once the alphabetic letter representations I, N, T, E, R, V, and W are activated, based on the written input *interview*, the corresponding phonemes (e.g, /ɪ/, /n/, /t/, /ɚ/, /v/, /j/, and /u/) are activated to derive the appropriate lexical phonological representation /ɪntɚvju:/. It is conceivable that, at the same time, the activation of English phonemes can lead to (at least partial) activation of the corresponding Japanese phonemes (e.g., /i/, /n/, /t/, /a/, /b/, /j/, and /u/), eventually leading to the activation of /inntabjuu/.

In this lexical decision study without priming, however, its effect was inhibitory at the first fixation, facilitatory at the first subgaze, and facilitatory also at the moment of the response. Previous studies reported mixed results: Dijkstra et al. (1999) reported an inhibitory effect of phonological similarity, and Lemhöfer and Dijkstra (2004) and Haigh and Jared (2007) reported facilitatory effects. Our results indicate that this is not an either-or problem but that inhibition and facilitation manifest themselves at different points in time. Given that this inhibitory effect was not replicated with the objective *PhonologicalDistance* measure, this may be due to early noise induced by co-activation of English and Japanese sublexical phonology (e.g., vowel-consonant distinction and stress pattern), which are not accounted for by the edit distance coding.

Second, L1 Japanese word frequency co-determined lexical decision responses in an interaction with L2 English word frequency. The English word frequency effect was progressively attenuated as Japanese frequency increased. Importantly, this interaction was also observed at the first subgaze, but not at the first and last fixations. This finding can be understood within the BIA+ model as follows. The BIA+ model posits, as does any interactive activation model, inhibitory links between non-identical orthographic/phonological lexical representations (see Figure 1, line c, and previous studies by Dijkstra et al., 1998; Kerkhofs et al., 2006). Because there are no links projecting from the English letter units to the Japanese *katakana* word representation (i.e., the dashed line a in Figure 1 is not active), the input word *interview* cannot activate the Japanese lexical *katakana* representation インタビュー at an early stage. However, the English input word *interview* activates the alphabetic letter representations I, N, T, E, R, V, and W, and consequently the English orthographic lexical representation INTERVIEW relatively early. The activation of the L1 Japanese lexical orthographic representation of a cognate can and does occur but only via indirect activation mediated by sublexical and lexical phonological representations or via top-down activation from the conceptual representation (see, for instance, the significant cross-language masked priming effect for non-cognate translation equivalents reported by Grainger & Frenck-Mestre, 1998; see also Duñabeitia, Perea, & Carreiras, 2010; Pecher, 2001; Perea, Duñabeitia, & Carreiras, 2008).

Third, a small significant facilitatory contribution of semantic similarity was observed. Its effect arose late, at the last fixation and in the lexical decision responses. In the present study, the semantic similarity effect did not depend on readers' L2 proficiency. This finding is consistent with models assuming a strong form-to-meaning mapping for both L1 and L2, allowing rapid semantic activation in L2 word processing (Duyck & Brysbaert, 2004).

Fourth, regression modeling allowed us to straightforwardly test the view that cognate facilitation arises from special morphological representations for cognates (Davis et al., 2010; Lalor & Kirsner, 2000; Sánchez-Casas & García-Albea, 2005). Although models with the factor *Cognate* as a sole lexical predictor replicated standard cognate facilitation effects (mixed-effect models not shown), it was no longer a significant predictor once the relevant numerical predictors were included in the regression equations. This is in harmony with Voga and Grainger's (2007) conclusion; a cognate facilitation effect is not due to special morphological representations but due to shared form overlaps.

Finally, we observed similarities and differences in how monolingual readers and bilingual readers make lexicality judgments. On one hand, the lexical decision processes of the two groups were comparable, as indicated by the lexical distributional properties that similarly co-determined the lexical decision responses of the two groups of readers (e.g., the interaction between *FreqHAL* and *OLD20_resid*). Such a functional overlap between groups is consistent with the general architecture of the BIA+ model, which is a generalization of the monolingual interactive activation model.

On the other hand, the lexical decision process of the bilingual readers diverged from that of the monolingual readers, beyond the bilingual-specific lexical effects mentioned above. At the first fixation, *Length* was significant only for monolinguals, and

OLD_resid was significant only for bilinguals. This may reflect the fact that the perceptual span of proficient readers is wider than that of less proficient readers (Rayner, Slattery & Belanger, 2010). At the last fixation, *Imageability_resid* was significant only for monolinguals.

Japanese-English bilinguals and English monolinguals also fine-tuned response criteria differently throughout the experiment. The former group apparently adjusted the response threshold for lexical decision with respect to a Japanese word property (i.e., cross-language phonological similarity). Because the co-activation of an L1 Japanese word via its phonological overlap with the L2 English word is a reasonable criterion for a 'yes' response in lexical decision, participants may have fine-tuned their response criteria so that phonology received progressively more weight to optimize responses. Monolinguals, on the other hand, fine-tuned response criteria with respect to word frequency and orthographic density (*FreqHAL* * *OLD20_resid*).

The above qualitative differences between bilinguals and monolinguals reconfirm that bilinguals are not two monolinguals in one mind (Grosjean, 1989). Although the BIA+ is, in terms of processing architecture, an extension of the monolingual IA model, such an architectural extension leads to qualitative and quantitative differences in processing (see Appendix E for a direct quantitative comparison between the two groups with the factor *FirstLanguage* in a mixed-effects analysis). The processing differences observed in this study do not provide evidence against the BIA+ model but point out specific areas that are yet to be clarified, namely the visuo-perceptual level and the decision/response level. The observed processing differences between bilinguals and monolinguals also imply that it requires caution to use monolinguals as experimental controls when studying bilingual processing.

Future research should further investigate potential consequences of quality of lexical predictors, experimental manipulations, and individual differences. As for quality of lexical predictors, human-rated phonological and semantic similarity measures were used in this study, as in many previous studies. The objective edit distance measure replicated the late effect, but not early effects, of the rated phonological similarity successfully. It is important to clarify what constitute rated measures and how to code them objectively (see Appendix F for a comparison of various rated measures). As for experimental manipulations, the font size chosen in this study was relatively larger than that in normal reading. While readers made multiple fixations most of the time in this study, an analysis of eight native English speakers' reading with smaller font (visual angle per letter = 0.4˚) revealed that multiple eye-movements were still used 73% of the time (median = 2, range = 1:5), indicating that isolated word reading itself triggers a task-specific pattern of eye-movements. As for individual differences, we only considered a potential effect of readers' length of stay in Canada, and we leave it to future research to disentangle contributions of various related measures. In addition, although monolinguals and bilinguals tested in this study were comparable in terms of age, more rigorous assessment should be conducted in the future (e.g., matched intelligence and reading speed). It is likely that the by-subject random intercepts, which capture between subject variability that cannot be traced back to the predictors included in the present study, comprise variation that can be explained by more refined measures characterizing the individual subjects, including measures for their reading skills.

In conclusion, without using a priming technique, the present study tapped into the time-course of lexical activation by observing eye movements to test various predictions of the BIA+ model. The bilingual-specific lexical processes that are characterized by early cross-language phonological similarity to an interaction between L1 and L2 frequency, and then to late cross-language semantic similarity is in line with the BIA+ model. The absence of a significant contribution of a factor coding cognate status indicates that a cognate facilitation effect can be sufficiently captured by numerical predictors coding form and meaning in two languages. The localist connectionist framework of the BIA+ model, which thus far has been guided by research on bilinguals with the same script, can be modified to account for lexical processes of Japanese-English bilinguals, under the straightforward assumption that English letter units do not project onto Japanese word units.

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Footnotes

[1] Nakayama, Kiryu, and Yamaguchi (2007) studied a corpus covering 10 years of a nation-wide newspaper from 1994 to 2003 and counted 8,226 katakana words that appeared at least 100 times. Typical Japanese dictionaries of katakana words list many more katakana words (e.g., the Concise Dictionary of Katakana Words lists as many as 56,300 words, Sanseido Henshujo, 1994).

Table 1

Lexical predictors, individual differences, and task effects considered in this study. The range and mean are presented for their original values before residualization and centralization procedures. The superscripts represent a transformation method used for the given predictor. The values for *Individual* and *Task* variables are those in Experiment 1 (Japanese-English bilinguals).

Type	Predictor	Range (Mean, SD) / Levels
Japanese-English	PhonologicalSimilarityJPN	2.3: 6.0 (*M* = 4.5, *SD* = 0.7)
	PhonologicalDistance	0.2 : 1.2 (*M* = 0.6, *SD* = 0.2)
	FreqJPN_resid $^{(log)}$	0 : 10.4 (*M* = 4.2, *SD* = 2.8)
	GoogleFreqJPN_resid $^{(log)}$	5.7 : 19.2 (*M* = 14.8, *SD* = 2.4)
	SemanticSimilarity_resid	3.0 : 7.0 (*M*= 5.9, *SD* = 0.7)
	Cognates	Levels: Cognate, NotCognate
English	Length	6 : 9 (*M* = 6.8, *SD* = 0.9)
	OLD20_resid	1.3 : 4.5 (*M* = 2.4, *SD* = 0.5)
	FreqHAL $^{(log)}$	7.8 : 13.0 (*M* = 9.6, *SD* = 1.1)
	SUBTLCD_resid $^{(log)}$	0 : 4.4 (*M* = 1.9, *SD* = 0.9)
	Imageability_resid	1.6 : 7.0 (*M* = 4.7 , *SD* = 1.7)
Individual	LengthOfStayCanada $^{(log)}$	1.1 : 5.3 (*M* = 2.7, *SD* = 1.3)
Task	PreviousResponseCorrect	Levels: Correct, Error
	Trial $^{(/100)}$	0.11 : 4.60 (*M* = 2.4, *SD* = 1.3)
	PreviousRT_resid $^{(-1000/)}$	-1.1 : 1.1 (*M* = 0, *SD* = 0.4)
	FirstSubgazeDuration_resid $^{(log)}$	4.9 : 7.8 (*M* = 6.2, *SD* = 0.4)

Table 2

Estimate, standard error, t-value, p-value, and effect size of the fixed effects in the model of Japanese-English bilingual readers' lexical decision response times. *Task* = task-related predictors, *Engl* = English target word properties, *Jpn-Engl* = Japanese-English-bilingual-specific predictors. The effect sizes refer to the magnitude of an effect calculated as the difference between the model's prediction for the minimum and the maximum back-transformed values of a given predictor

Response time	Type	Estimate	Std.Error	t-value	p-value	Effect size
(Intercept)		-1.477	0.037	-40.43	< 0.0001	
PreviousResponseCorrect (Error)	Task	0.064	0.017	3.69	0.0002	28
Trial	Task	-0.085	0.006	-13.52	< 0.0001	-168
PreviousRT_resid	Task	0.135	0.012	10.81	< 0.0001	124
Length	Engl	0.046	0.008	5.79	< 0.0001	65
OLD20_resid	Engl	0.033	0.021	1.55	0.1209	26
FreqHAL	Engl	-0.064	0.007	-9.61	< 0.0001	-129
SUBTLCD_resid	Engl	-0.061	0.010	-5.89	< 0.0001	-111
FreqJPN_resid	Jpn-Engl	-0.003	0.003	-1.04	0.2966	-14
PhonologicalSimilarityJPN	Jpn-Engl	-0.027	0.011	-2.38	0.0172	-44
SemanticSimilarity_resid	Jpn-Engl	-0.037	0.011	-3.27	0.0011	-70
PreviousRT_resid * Trial	Jpn-Engl * Task	-0.028	0.009	-2.98	0.0029	No plotted
PhonologicalSimilarityJPN * Trial	Jpn-Engl * Task	-0.011	0.005	-2.28	0.0227	Figure 2 (a)
FreqHAL * FreqJPN_resid	Engl * Jpn-Engl	0.006	0.003	2.53	0.0116	Figure 2 (b)
SemanticSimilarity_resid * PreviousRT_resid	Jpn-Engl * Task	-0.046	0.016	-2.92	0.0035	Figure 2 (c)
FreqHAL * OLD20_resid	Engl * Engl	-0.050	0.018	-2.74	0.0061	Figure 2 (d)

Table 3

Estimate, standard error, t-value, p-value, and effect size (ms) of the fixed effects in the model of Japanese-English bilingual readers' first fixation durations, first subgaze durations, and last fixation durations.

First fixation duration	Type	Estimate	Std.Error	t-value	p-value	Effect size
(Intercept)		5.673	0.025	228.98	< 0.0001	
Trial	Task	0.011	0.004	2.93	0.0034	14
PreviousRT_resid	Task	-0.025	0.011	-2.23	0.0256	-16
OLD20_resid	Engl	0.040	0.013	2.99	0.0028	23
FreqHAL	Engl	-0.009	0.004	-1.99	0.0467	-13
PhonologicalSimilarityJPN	Jpn-Engl	0.017	0.007	2.29	0.0219	18

First subgaze duration	Type	Estimate	Std.Error	t-value	p-value	Effect size
(Intercept)		6.207	0.037	168.46	< 0.0001	
Trial	Task	-0.087	0.009	-9.41	< 0.0001	-181
PreviousRT_resid	Task	0.129	0.017	7.78	< 0.0001	134
Length	Engl	0.084	0.008	10.42	< 0.0001	130
OLD_resid	Engl	0.039	0.021	1.83	0.0680	34
FreqHAL	Engl	-0.064	0.007	-9.54	< 0.0001	-143
SUBTLCD_resid	Engl	-0.068	0.010	-6.61	< 0.0001	-132
PhonologicalSimilarityJPN	Jpn-Engl	-0.029	0.012	-2.45	0.0142	-49
FreqJPN_resid	Jpn-Engl	-0.004	0.003	-1.22	0.2215	-18
FreqHAL * FreqJPN_resid	Engl * Jpn-Engl	0.005	0.003	2.04	0.0412	Figure 3 (a)

Last fixation duration	Type	Estimate	Std.Error	t-value	p-value	Effect size
(Intercept)		13.282	0.261	50.89	< 0.0001	
FirstSubgazeDuration_resid	Task	-8.006	0.145	-55.23	< 0.0001	-507
PreviousResponseCorrect (Error)	Task	0.667	0.130	5.14	< 0.0001	18
Trial	Task	0.257	0.038	6.83	< 0.0001	31
PreviousRT_resid	Task	-1.643	0.158	-10.41	< 0.0001	-104
FreqHAL	Engl	0.142	0.071	1.98	0.0474	20
SUBTLCD_resid	Engl	0.240	0.112	2.14	0.0323	27
FreqJPN_resid	Jpn-Engl	0.016	0.031	0.52	0.6042	0
PhonologicalSimilarityJPN	Jpn-Engl	0.006	0.124	0.05	0.9609	5
SemanticSimilarity_resid	Jpn-Engl	-0.497	0.121	-4.10	< 0.0001	-49
PhonologicalSimilarityJPN * FreqJPN_resid	Jpn-Engl * Jpn-Engl	-0.104	0.044	-2.38	0.0175	Figure 3 (b)

SemanticSimilarity_resid * PhonologicalSimilarityJPN	Jpn-Engl * Jpn-Engl	0.444	0.199	2.23	0.0259	Figure 3 (c)
SemanticSimilarity_resid * PreviousRT_resid	Jpn-Engl * Task	-0.616	0.199	-3.09	0.0020	Figure 3 (d)

Table 4

Estimate, standard error, t-value, p-value, and effect size (ms) of the fixed effects in the model of English monolingual readers' lexical decision response times.

Response time	Type	Estimate	Std.Error	t-value	p-value	Effect size
(Intercept)		-1.946	0.034	-57.59	< 0.0001	
PreviousResponseCorrect (Error)	Task	0.085	0.018	4.82	< 0.0001	23
Trial	Task	-0.040	0.008	-5.25	< 0.0001	-46
PreviousRT_resid	Task	0.181	0.020	9.27	< 0.0001	124
Length	Engl	0.029	0.007	4.33	< 0.0001	23
OLD20_resid	Engl	0.005	0.018	0.30	0.7668	3
FreqHAL	Engl	-0.030	0.006	-5.22	< 0.0001	-37
SUBTLCD_resid	Engl	-0.058	0.009	-6.64	< 0.0001	-63
Imageability_resid	Engl	-0.011	0.004	-2.81	0.0049	-19
OLD20_resid * Trial	Engl * Task	0.019	0.009	2.07	0.0381	Figure 4 (a, b)
FreqHAL * Trial	Engl * Task	-0.002	0.003	-0.71	0.4749	Figure 4 (a, b)
FreqHAL * OLD20_resid	Engl * Engl	-0.027	0.016	-1.70	0.0886	Figure 4 (a, b)
FreqHAL * OLD20_resid * Trial	Engl * Engl * Task	-0.016	0.008	-2.08	0.0380	Figure 4 (a, b)

Table 5

Estimate, standard error, t-value, p-value, and effect size (ms) of the fixed effects in the model of English monolingual readers' first fixation durations, first subgaze durations, and last fixation durations.

First fixation duration	Type	Estimate	Std.Error	t-values	p-value	Effect size
(Intercept)		5.620	0.024	230.71	< 0.0001	
Trial	Task	0.007	0.004	1.69	0.0911	8
PreviousRT_resid	Task	-0.023	0.008	-2.98	0.0029	-17
Length	Engl	0.031	0.007	4.51	< 0.0001	26

First subgaze duration	Type	Estimate	Std.Error	t-value	p-value	Effect size
(Intercept)		-3.157	0.094	-33.5	< 0.0001	
Trial	Task	-0.047	0.018	-2.61	0.0092	-11
PreviousRT_resid	Task	0.136	0.036	3.82	0.0001	32
Length	Engl	0.223	0.017	13.48	< 0.0001	76
FreqHAL	Engl	-0.043	0.014	-3.12	0.0018	-20
SUBTLCD_resid	Engl	-0.089	0.021	-4.18	< 0.0001	-34
Length * Trial	Engl *Task	-0.024	0.010	-2.44	0.0147	Not plotted

Last fixation duration	Type	Estimate	Std.Error	t-value	p-value	Effect size
(Intercept)		14.164	0.184	76.91	0.0000	
FirstSubgazeDuration_resid	Task	-3.050	0.063	-48.11	0.0000	-1080
PreviousResponseCorrect (Error)	Task	0.530	0.167	3.17	0.0015	15
Imageability_resid	Engl	-0.100	0.042	-2.36	0.0181	-19

Figure Captions

Figure 1. A bilingual interactive activation (BIA+) architecture applied for Japanese-English bilinguals' processing of an L2 English word. Arrows represent facilitatory links and circular connectors represent inhibitory links.

Figure 2. Lexical interactions in the mixed-effects model for Japanese-English bilinguals' lexical decision response times. Different lines represent quantiles, and the rug in the x-axes represents the pattern of distribution.

Figure 3. Lexical interactions in the mixed-effects model for Japanese-English bilinguals' first subgaze and last fixation durations in trials with two fixations. Different lines represent quantiles, and the rug in the x-axes represents the pattern of distribution.

Figure 4. Lexical interactions in the mixed-effects model for monolingual English readers' response latencies. Different lines represent quantiles, and the rug in the x-axes represents the pattern of distribution.

Figure 1

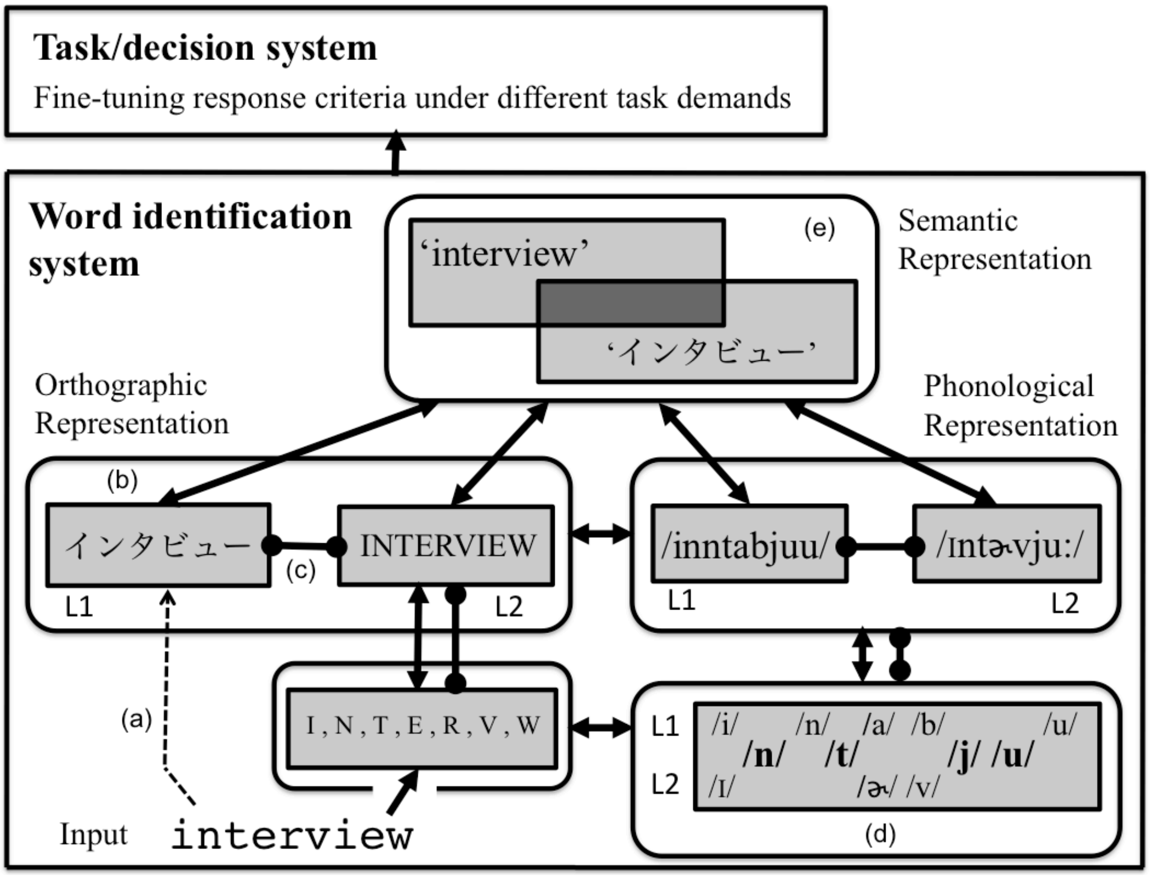
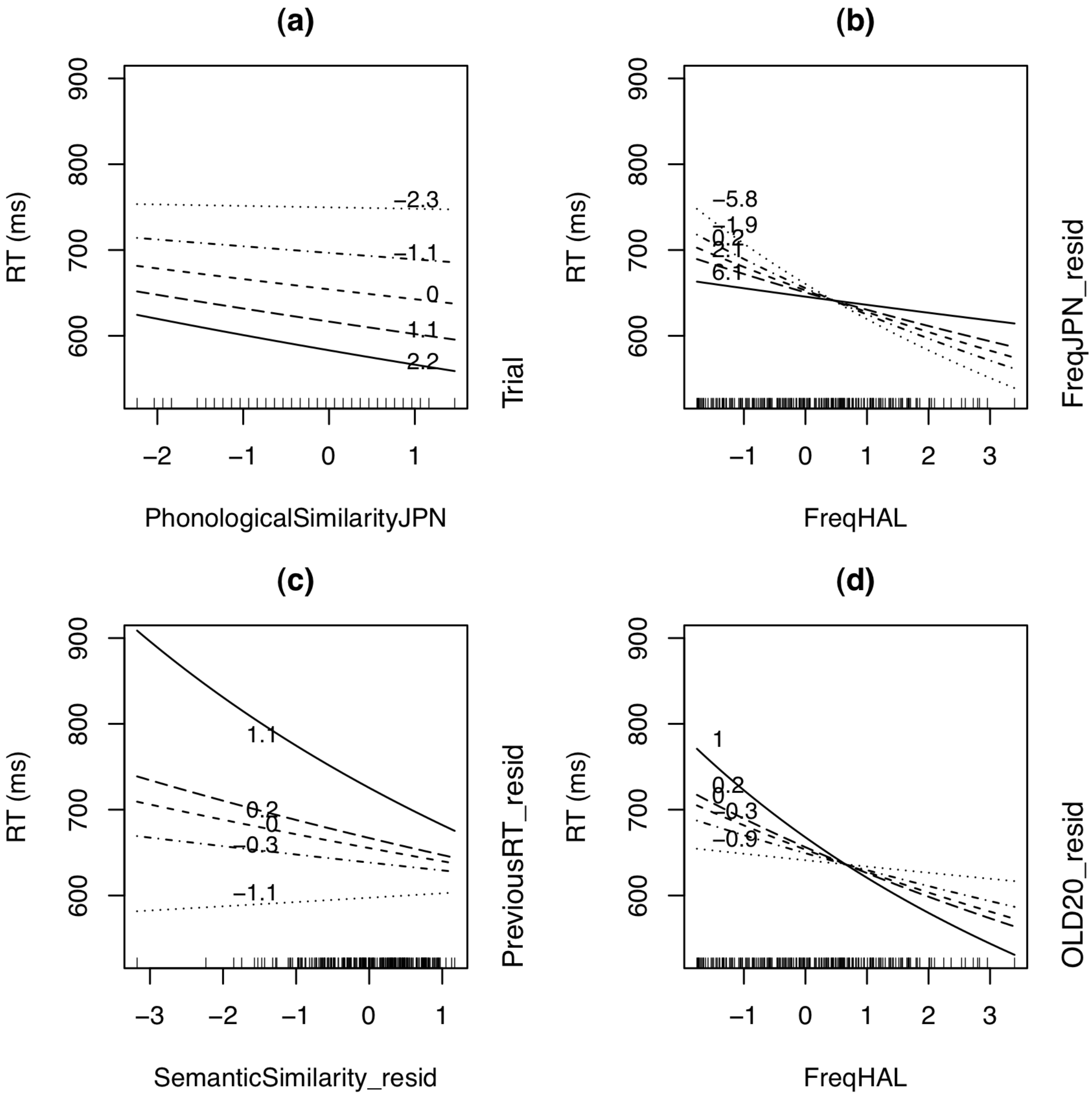

Task/decision system
Fine-tuning response criteria under different task demands
Word identification system
'interview'
(e)
'インタビュー'
Semantic Representation
Orthographic Representation
Phonological Representation
(b)
インタビュー
(c)
INTERVIEW
L1
L2
/inntabjuu/
/ɪntəvju:/
L1
L2
(a)
I , N , T , E , R , V , W
L1
L2
/i/ /n/ /a/ /b/ /u/
/n/ /t/ /j/ /u/
/I/ /ə/ /v/
(d)
Input
interview

Figure 2



(a)
RT (ms)
900
800
700
600
−2.3
−1.1
0
1.1
2.2
Trial
−2
−1
0
1
PhonologicalSimilarityJPN
(b)
RT (ms)
−5.8
−1.9
0.2
2.1
6.1
FreqJPN_resid
−1
0
1
2
3
FreqHAL
(c)
RT (ms)
1.1
0.2
0
−0.3
−1.1
PreviousRT_resid
−3
−2
−1
0
1
SemanticSimilarity_resid
(d)
RT (ms)
1
0.2
0
−0.3
−0.9
OLD20_resid
FreqHAL

Figure 3

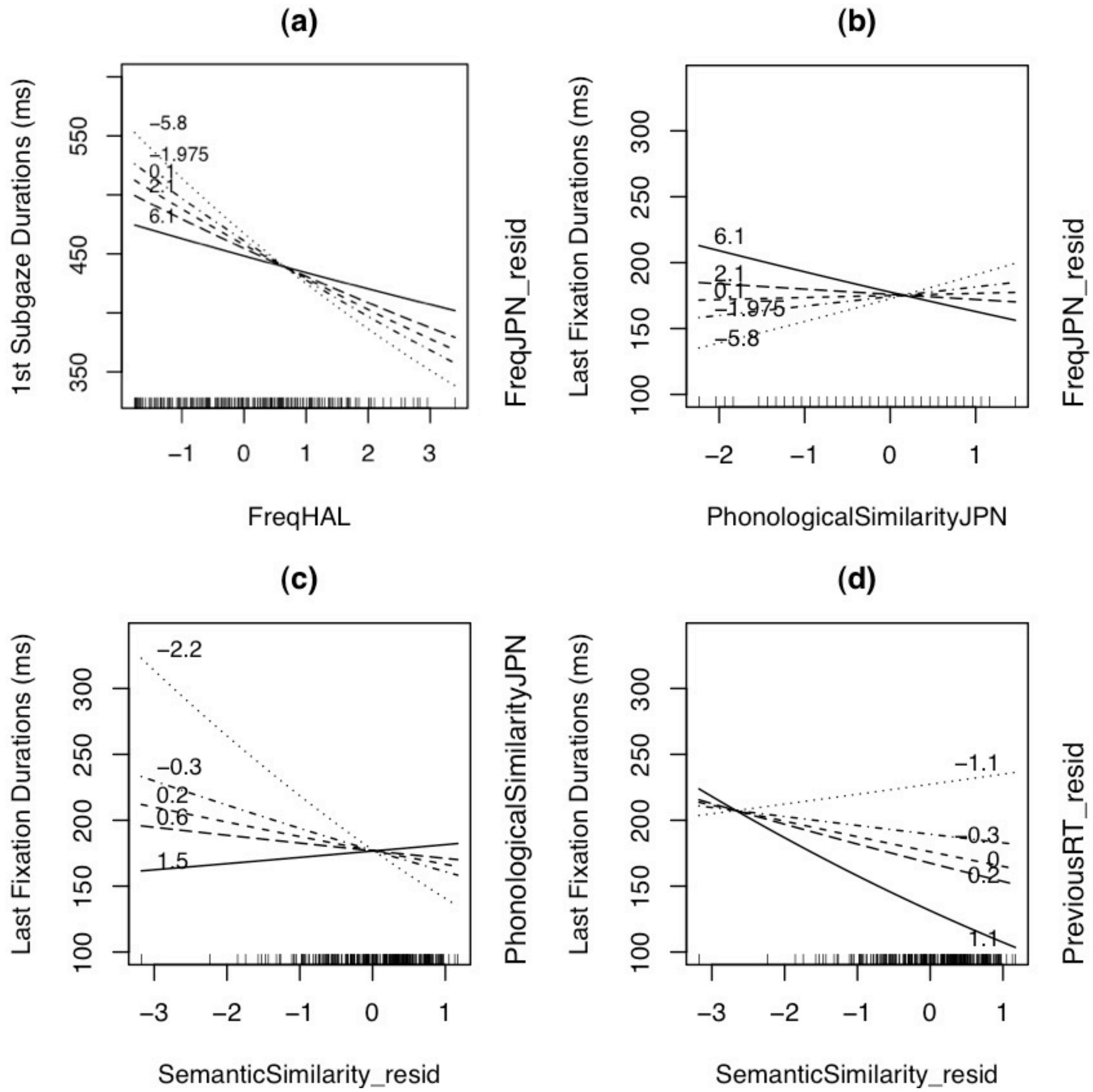
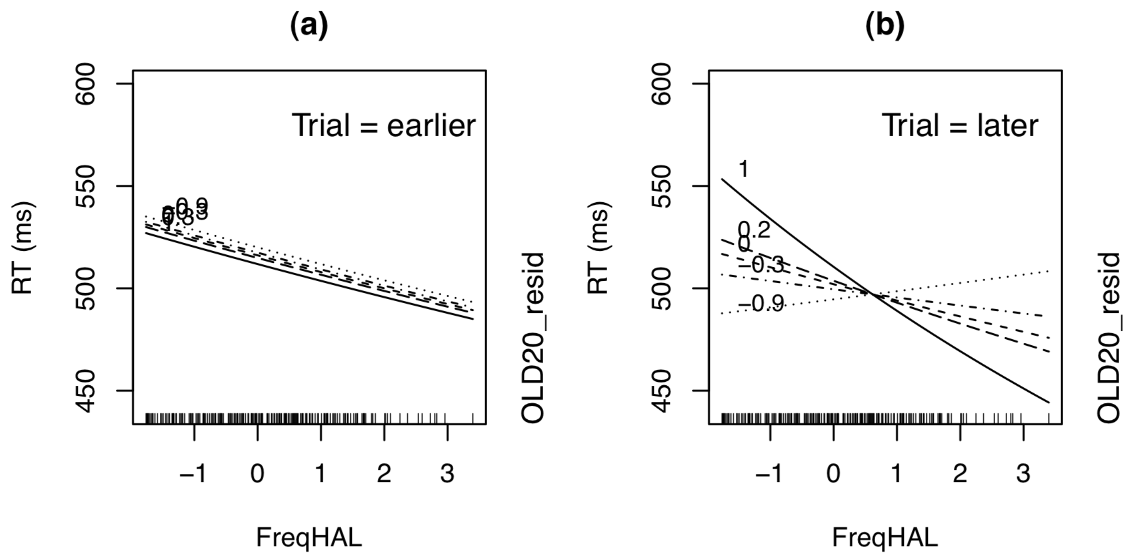

(a)
1st Subgaze Durations (ms)
FreqHAL
FreqJPN_resid
−5.8
−1.975
0.1
2.1
6.1
(b)
Last Fixation Durations (ms)
PhonologicalSimilarityJPN
FreqJPN_resid
6.1
2.1
0.1
−1.975
−5.8
(c)
Last Fixation Durations (ms)
SemanticSimilarity_resid
PhonologicalSimilarityJPN
−2.2
−0.3
0.2
0.6
1.5
(d)
Last Fixation Durations (ms)
SemanticSimilarity_resid
PreviousRT_resid
−1.1
−0.3
0
0.2
1.1

Figure 4


(a)
Trial = earlier
RT (ms)
600
550
500
450
OLD20_resid
−1
0
1
2
3
FreqHAL
(b)
Trial = later
1
0.2
0
−0.3
−0.9
RT (ms)
600
550
500
450
OLD20_resid
−1
0
1
2
3
FreqHAL