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Statistical learning of novel graphotactic constraints in children and adults

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A B S T R A C T

The current study explored statistical learning processes in the acquisition of orthographic knowledge in school-aged children and skilled adults. Learning of novel graphotactic constraints on the position and context of letter distributions was induced by means of a two-phase learning task adapted from Onishi, Chambers, and Fisher (*Cognition, 83* (2002) B13–B23). Following incidental exposure to pattern-embedding stimuli in Phase 1, participants' learning generalization was tested in Phase 2 with legality judgments about novel conforming/nonconforming word-like strings. Test phase performance was above chance, suggesting that both types of constraints were reliably learned even after relatively brief exposure. As hypothesized, signal detection theory d' analyses confirmed that learning permissible letter positions ($d' = 0.97$) was easier than permissible neighboring letter contexts ($d' = 0.19$). Adults were more accurate than children in all but a strict analysis of the contextual constraints condition. Consistent with the statistical learning perspective in literacy, our results suggest that statistical learning mechanisms contribute to children's and adults' acquisition of knowledge about graphotactic constraints similar to those existing in their orthography.

Introduction

There is a growing interest in the development of orthographic knowledge and its contribution to skilled reading and spelling. Learners become sensitive to general properties of their orthography—such as the frequency of occurrence of individual letters/letter sequences in print and the legality of different

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spelling patterns—earlier in development than previously thought (Cassar & Treiman, 1997; Ferreiro & Teberosky, 1982; Hayes, Treiman, & Kessler, 2006; Pacton, Perruchet, Fayol, & Cleeremans, 2001; Treiman, 1993). These findings challenge the long-standing view that phonological information is the only resource available to beginning spellers (e.g., Frith, 1985; Gentry, 1982). In fact, recent theorizing about spelling development emphasizes children's early insights into different types and levels of linguistic knowledge, including knowledge of simple orthographic and morphological conventions (Bourassa & Treiman, 2001; Deacon, Conrad, & Pacton, 2008; Treiman & Bourassa, 2000).

Given that some information about legal orthographic forms is available to children before or as soon as literacy instruction begins, learning must be, to some extent, incidental in nature. Accordingly, several authors have argued that children's pattern extraction skills rely on *implicit* or *statistical learning* processes (e.g., Kessler, 2009; Pollo, Treiman, & Kessler, 2007; Steffler, 2001; Treiman & Kessler, 2013). However, to date little empirical research has examined the mechanisms underpinning the acquisition of orthographic knowledge. In addition, very few studies have investigated learning of written language patterns under well-controlled experimental conditions that allow manipulations of the type and amount of exposure given to the observer. The goal of the current study was to address these issues in a group of 7-year-old children in comparison with a group of adults. We investigated whether sensitivity to novel orthographic constraints similar to those found in the English orthography develop under incidental learning conditions. Can learning be induced in 7-year-old typically developing children as well as in adults?

Development of sensitivity to orthographic structure

In a seminal study, Treiman (1993) examined a large corpus of naturalistic writing data produced by a group of U.S. English-speaking children at the end of kindergarten or beginning of first grade. Detailed analyses of children's misspellings revealed that even the youngest children actively used some knowledge of simple spelling conventions in their own written productions such as the positional constraint that "*ck* is not a legal onset in English." Moreover, first graders rarely committed errors that were inconsistent with orthographic constraints on permissible letter doublets (e.g., *xx, kk*). Treiman further demonstrated that school beginners are sensitive to statistical probabilities in print using simple nonword judgment tasks (e.g., which one looks more like a real word, *moyl* or *moil*?).

The standard version of Treiman's (1993) now widely used orthographic constraints task presents participants with the oral pronunciation of a nonword and asks them to choose between two alternative written spellings, only one of which conforms to patterns in their orthography. In one of the first experimental studies of this sort, Cassar and Treiman (1997) assessed children's sensitivity to untaught constraints on allowable consonant and vowel doublets. Children in Grade 1 preferred nonwords containing doublets in allowable positions (in word middles but not at word beginnings) as well as permissible doublets (e.g., *ll* but not *xx*). Sensitivity to both types of constraints increased as a function of age. However, more complex relationships, such as the influence of phonological context on consonantal doubling, had an effect only in more advanced spellers' choices (Grade 6). Cassar and Treiman's key findings have been replicated in languages other than English (French: Pacton et al., 2001; Finnish: Lehtonen & Bryant, 2005). Moreover, Pacton and colleagues (2001) demonstrated in recognition (judgment) and fragment completion production tasks that French-speaking first graders have a preference for nonwords embedding letters that are frequently doubled in their orthography (e.g., *illaro* > *ivvaro*) and for stimuli containing a doublet in a medial legal position over stimuli where the doublet is illegally situated (e.g., *nnulor* < *nullor*).

Young learners' appreciation of positional constraints, such as those governing the legality of consonantal doublets, provides them with but one cue to correct spelling. Taking surrounding context into account further facilitates the process of translating speech to print, in inconsistent orthographies (Alegria & Mousty, 1996; Kessler & Treiman, 2001; Pacton, Fayol, & Perruchet, 2002, 2005) as well as more consistent orthographies (Caravolas & Mikulajová, 2008). Several studies in English have demonstrated that children and adults take advantage of contextually conditioned phonological patterns in order to spell consonants (Treiman & Kessler, 2006; Treiman, Kessler, & Bick, 2002) or vowels (Hayes et al., 2006; see also Varnhagen, Boechler, & Steffler, 1999), although this type of knowledge seems to develop gradually, with some contexts being learned more quickly than others. For example, Treiman and Kessler (2006) showed that, by the seventh-grade level, children's spellings reliably reflected knowledge of contingencies between vowel–coda contexts such as that the vowel /ɛ/ is commonly spelled as *ea* when followed by the coda /d/—as in *head*—but not when followed by other coda consonants—as in *hen* or *help* (phonemes are denoted with International Phonetic Alphabet symbols throughout this article; see International Phonetic Association, 1999; International Phonetic Association., 1996). In the same study, already by the fourth-grade level, children's spellings demonstrated reliable sensitivity to onset–vowel contexts such as that the vowel /ɝ/ is usually spelled as *or* when preceded by the onset /w/—as in *word*—whereas several other spellings follow when the onset consonants are different such as /b/ or /k/ in *bird, curd,* and the like.

In the above studies, surrounding context was manipulated through phonological cues. Of particular relevance to the current study, Hayes and colleagues (2006) investigated children's sensitivity to contextually conditioned spelling patterns that were better described in graphotactic rather than phonological terms. For example, an alternative account for the doubling effects in the context of long versus short vowel phonemes in Cassar and Treiman's (1997) study is that consonantal coda spellings are extended (e.g., doubled) when preceded by vowels spelled with a single letter (e.g., *Jeff*), whereas two-letter vowel spellings do not require an extension of the following consonant (e.g., *deaf*). Thus, the number of letters in the vowel spelling determines the number of letters in the coda spelling—a graphotactic effect that is independent of phonological information. Accordingly, children as young as Grade 2 preferred spellings with coda doublets (e.g., *ff*) when preceded by single vowel spellings, but they preferred single-letter codas (e.g., *f*) when preceded by two-letter vowels. Even more interesting, children's spelling of nonwords reflected the graphotactic rule (number of letters) rather than the phonological rule (length of phoneme).

In summary, the above literature provides evidence of children's sensitivity to certain statistical properties of the orthography during the early stages of literacy acquisition. Undoubtedly, children's initial orthographic sensitivity is limited and imperfect (Pacton et al., 2002). Thus, an aim of the current study was to investigate differences in the learning of simpler, positionally conditioned spelling patterns versus more complex, contextually conditioned spelling patterns. Studies such as those by Treiman and colleagues demonstrate that knowledge of more complex, context-based patterns in the orthography takes longer to develop and reaches adult-like levels only in more advanced spellers (Hayes et al., 2006; Treiman & Kessler, 2006). However, none of these studies directly assesses how learning of simple or more complex orthographic patterns occurs. Are relatively few exposures to word-like strings embedding the constraints sufficient for learning? Does pattern complexity moderate learnability? We sought to address these questions by inducing learning under brief incidental conditions.

Our approach was informed by recent research on statistical learning in different knowledge domains. Previous studies have shown that longer exposure enhances incidental learning of statistical regularities in a wide range of domains (e.g., sequential structure learning: Gaillard, Destrebecqz, Michiels, & Cleeremans, 2009; harmonic pattern learning: Jonaitis & Saffran, 2009). For example, Saffran, Newport, Aslin, Tunick, and Barrueco (1997) showed that longer exposure to a continuous speech stream has a direct benefit to participants' ability to parse syllable words. Using the same task, Evans, Saffran, and Robe-Torres (2009) replicated this finding in a subgroup of children with specific language impairment. We also investigated the effects of amount of exposure on incidental graphotactic learning in the current study.

Statistical learning perspective

Although children are typically taught to spell through systematic explicit instruction in the classroom, learning to spell is not likely to be determined by explicit processes alone (Steffler, 2001), perhaps especially in highly inconsistent orthographies. Spelling rules that are more probabilistic in nature are hard to teach overtly or to state even for proficient linguists (Kessler, 2009). Thus, a plausible hypothesis is that spellers also acquire pattern knowledge in an implicit fashion, that is, through simple exposure to print or reading experience. According to this view, children are powerful statistical learners who observe and internalize the relative frequency with which letters or letter combinations occur and co-occur (Pollo et al., 2007). Sensitivity to the distributional information in written language is thought to rely on the same general learning mechanisms operating in different domains (e.g., Fiser & Aslin, 2001, 2002) and sensory modalities (e.g., Conway & Christiansen, 2002) by abstracting statistically defined patterns from the input. Statistical or distributional learning underlies learning of a wide range of contingencies, ranging from basic statistics (e.g., frequency of occurrence of individual elements) to more complicated patterns such as conditional probability statistics (Romberg & Saffran, 2010).

Statistical learning processes in the acquisition of natural languages have been abundantly demonstrated to operate from early infancy onward (for a review, see Romberg & Saffran, 2010). For example, infants (8-month-olds), children, and adults are able to compute and exploit syllabic transitional information in order to postulate word boundaries (e.g., Saffran, Aslin, & Newport, 1996; Saffran et al., 1997). Probabilistic distributional learning has been shown to subserve other aspects of language acquisition such as the ability to discover rudimentary syntax (e.g., Thompson & Newport, 2007) and grammatical category membership (e.g., Mintz, 2002). Another line of research with direct relevance to the current work has explored implicit learning at the level of phonotactic constraints. Onishi and colleagues have demonstrated that infants and adults become sensitive to novel phonotactic constraints embedded within aurally presented speech sequences without engaging deliberate effort (Chambers, Onishi, & Fisher, 2003, 2010; Onishi, Chambers, & Fisher, 2002). For example, in a syllable repetition task, adult participants were faster at repeating novel syllables containing phoneme sequences that conformed to the experimental constraints on position (e.g., syllables with permissible phonemes in the onset position) than novel sequences that violated them (e.g., syllables with phonemes not permissible in the onset position). In addition, repetition reaction times were faster for novel syllables embedding newly learned contingencies between vowel phonemes and onset/coda phonemes (context-based learning). Positional learning effects were also observed in a follow-up study with preverbal infants. Following brief habituation with critical syllables similar to those presented to adults, infants demonstrated a novelty preference, with longer listening times for new items violating rather than conforming to the positional constraints.

Aim of the current study

Using an adaptation of Onishi and colleagues' (2002) paradigm, we investigated whether statistical learning is involved in the acquisition of orthographic knowledge, in particular knowledge about graphotactic constraints of varying complexity (position vs. context). We sought to induce learning similar to that demonstrated for aurally presented phonotactically constrained syllables but through the visual domain. Therefore, we used visual presentation of graphotactically constrained syllables. All of the stimuli contained onset, vowel, and coda graphemes as well as bigrams (consonant–vowel [CV] and VC) that are permissible in the English orthography (and phonology), and all conformed to English graphotactic rules for permissible onset, vowel, and coda spellings for monosyllabic words (although a small number of tokens occurred less often in the English orthography). Thus, participants were not exposed to unnatural graphotactic patterns; rather, they were merely exposed to novel sequences and relative distributions of these. A between-participant design was used to investigate whether learning can be induced under the same brief experimental conditions in typically developing 7-year-olds and adults. Similar to previous studies (e.g., Gaillard et al., 2009), we manipulated the amount of exposure (short vs. long) to enhance learning.

Method

Participants

Adults

A sample of 113 undergraduate students (29 male and 84 female) with normal or corrected-to-normal visual acuity was recruited through a psychology department participant panel and received course credit or small monetary compensation for their participation. All were monolingual native English speakers and reported no history of dyslexia or other learning difficulties. They were randomly

Table 1
Proportion of items endorsed as legal (and *SD*) as a function of stimulus type in each experimental condition and the resulting signal detection theory measures.

Condition	*n* (m;f)	Endorsement rates[a]		SDT measures	
		Illegal (FA)	Legal unseen (hits)	*d*′	Bias *c*
Adults					
PC short	27 (8;19)	.32 (0.13)	.70 (0.13)	1.05 (0.57)	−0.03 (0.26)
CC short	28 (7;21)	.46 (0.12)	.58 (0.11)	0.30 (0.37)	−0.06 (0.25)
PC long	29 (6;23)	.26 (0.18)	.65 (0.18)	1.23 (0.80)	0.17 (0.45)
CC long	29 (8;21)	.49 (0.16)	.57 (0.13)	0.21 (0.37)	−0.07 (0.35)
Children					
PC short	33 (18;15)	.32 (0.16)	.55 (0.18)	0.66 (0.62)	0.18 (0.41)
CC short	34 (13;21)	.40 (0.20)	.44 (0.17)	0.14 (0.47)	0.24 (0.51)
PC long	34 (20;14)	.28 (0.17)	.59 (0.19)	0.95 (0.77)	0.22 (0.43)
CC long	31 (16;15)	.44 (0.18)	.48 (0.15)	0.12 (0.45)	0.11 (0.39)

Note. m, males; f, females; FA, false alarms; SDT, signal detection theory; PC, positional constraints; CC, contextual constraints.
[a] Adults' endorsement rates were calculated based on all test phase data. Children's endorsement rates were calculated based on test phase data above 300 ms.

assigned to one of four experimental conditions formed by crossing the factors of type of constraint (positional vs. contextual) and exposure duration (short vs. long). Sample sizes are reported in Table 1. All participants gave their informed consent before the start of the experimental session.

Children

A group of 137 typically developing children at the end of Year 2 (comparable in age to Grade 1 in the United States) or beginning of Year 3 (71 male and 66 female) was drawn from a longitudinal cohort of 189 monolingual English-speaking children (Caravolas et al., 2012). The sample was recruited from seven classrooms of six primary schools in three counties of Northern England. At the time of the current study, the mean age of the cohort was 88.75 months (7;5 [years;months], range = 6;10–8;2). Children were randomly assigned to one of four groups (sample sizes are reported in Table 1). Across schools, all children had received the same amount of formal tuition (almost 3 years) and were being taught by similar literacy methods, according to governmental guidelines. An earlier study (Caravolas et al., 2012) revealed no significant effects of school on the cohort's development of literacy skills.

Material

Following the paradigm of Onishi and colleagues (2002), 32 Ci_Cj syllable frames were created by using two sets of four consonants (Set 1: d, m, l, and f; Set 2: t, n, p, and s) and combining them with each other. For half of the syllable frames, Set 1 consonants were used as onsets and Set 2 consonants were used as codas (C_1_C_2 sequences; e.g., d_t); for the other half, the reverse was true (C_2_C_1 sequences; e.g., t_d). Two vowels (o and e) were used to fill in the syllable frames, giving rise to a total of 64 CVC sequences that were arranged in two pairs of lists (i.e., four lists in total), and the presentation of these was counterbalanced within experimental groups[1] (see further details in Appendixes A and B). Although the particular sets of consonants and vowels were chosen in an effort to keep the number of real words to a minimum, 7 of the 32 C_1VC_2 sequences and 1 of the 32 C_2VC_1 sequences were unavoidably real words. To investigate whether the pattern of results obtained in the main analyses was unduly influenced by the presence of real words, follow-up analyses that excluded real words were also carried out.

For positional constraints learning in one of the counterbalanced list conditions, one C_1VC_2 list (e.g., List 1, $n = 16$) served as exposure items, whereas the other list (e.g., List 2, $n = 16$) served as legal

[1] Unfortunately, due to experimenter error, Lists 3 and 4 were not used as exposure and legal unseen materials with adult participants; only Lists 1 and 2 were used for this purpose and were counterbalanced within experimental groups (List 3 served as illegal test items).

unseen items at test. C_2VC_1 items (List 3, *n* = 16) served as illegal items during the test phase. Items presented during the exposure phase were governed by the following underlying zero-order statistics. First, in Lists 1 and 2, for example, there was an equal probability of appearance of any C_1 letter as an onset [e.g., *p*(d) = .25], whereas the probability of appearance of any C_2 letter as an onset [e.g., *p*(t)] was zero. Second, there was an equal probability of appearance of any C_2 letter as a coda [e.g., *p*(t) = .25], whereas the probability of appearance of any C_1 letter as a coda [e.g., *p*(d)] was zero.

For contextual constraints learning, stimuli were recombined into two pairs of lists in order to introduce a first-order constraint according to which consonant position depended on the vowel type. For one of the counterbalanced list conditions, stimuli with Set 1 consonants as onsets and Set 2 consonants as codas were combined with the vowel *o* (e.g., fos), whereas sequences with Set 2 consonants as onsets and Set 1 consonants as codas were combined with the vowel *e* (e.g., tem). Item assignment to list was counterbalanced across participants in a manner similar to positional constraints learning. For example, in one condition, List 1 items (*n* = 16) served as exposure items, List 2 items (*n* = 16) were presented at test as legal unseen items, and List 3 items (*n* = 16) were presented at test as illegal items. In contrast to the positional constraints variant, the probability of appearance of any C_1 or C_2 letter as an onset or coda was equated for contextual constraints learning [e.g., *p*(d) = *p*(t) = .125]. Exposure items were governed by the following underlying first-order statistics. In Lists 1 and 2, for example, there was an equal joint probability of occurrence of the vowel *o* and any C_1 letter as an onset [e.g., *p*(d, o) = *p*(m, o) = .125] and the vowel *e* and any C_2 letter as an onset [e.g., *p*(t, e) = *p*(n, e) = .125]. Consequently, the joint probability of appearance of any C_1 letter and *e* [e.g., *p*(d, e)] or any C_2 letter and *o* [e.g., *p*(m, o)] was zero. A corollary of this design was that VC contingencies [e.g., *p*(o, t), *p*(o, n)] also occurred with a probability of .125. Thus, participants could benefit from first-order contingencies in the CV or VC portions of the stimuli. Importantly, these first-order contingencies occurred systematically and with the same statistical probabilities throughout the task.

Apparatus

Computer experiments were run on a Windows XP-based PC with a 15-inch CRT color monitor. Stimulus displays were generated using E-Prime software (Schneider, Eschman, & Zuccolotto, 2002). Presentation of stimuli and millisecond accurate response registration was achieved by means of the same software package. Participants responded on a standard QWERTY keyboard.

Procedure

Adult participants were tested individually in a quiet room. Children were tested in a quiet area of the school. All participants were first administered the learning task followed by several cognitive ability measures, which are not considered here. Testing took place in a single 30-min session.

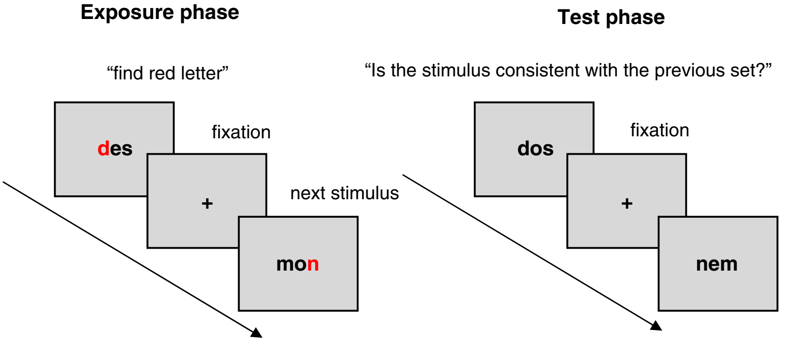

Fig. 1. The incidental graphotactic learning task: Trial outline.

The incidental graphotactic learning task consisted of an exposure phase, a distraction phase, and a test phase (Fig. 1).

The exposure phase was presented as a (red) color detection task. At the beginning of each exposure trial, a three-letter string with one red letter and two white letters was displayed at the center of a black background. Participants were asked to respond to the position of the red letter as fast and accurately as possible by pressing one of three keys. Strings remained on the screen until a response was collected, and no feedback was given. Stimulus presentation was followed by a fixation point (white cross) appearing centrally for 500 ms. A practice block of 6 trials was given in order to familiarize participants with the task. This was followed by either three or six exposure blocks of 48 trials each, for a total of 144 trials (9 repetitions/string) or 288 trials (18 repetitions/string) in the short exposure and long exposure conditions, respectively. The target red letter appeared in each of the first, mid, or final position one third of the time.

After all exposure strings were presented, participants moved to a short distraction phase involving simple calculations. Adult participants completed 10 simple arithmetic facts (single-digit additions of the form $x + y = z$, with $1 \leqslant x, y \leqslant 8$, and $2 \leqslant z \leqslant 9$) by pressing on the number key corresponding to the correct result. No feedback was given during this phase of the experiment.

Completion of the distraction phase was followed by a "surprise" test phase. Participants were informed that the strings they had previously seen were constructed according to a set of rules. They were told that during this phase they would see a set of new strings and were to decide whether each one of them "goes well" with the rules. To avoid extreme biases in the test responses, participants were informed that half of the stimuli were compatible with the previously seen set of strings, whereas the other half were not. Similar to the instructions given in implicit learning experiments, participants were encouraged to trust their "gut feeling" whenever they were not sure about their response. Each CVC test item appeared in the center of the screen, and participants pressed one button if they judged it to conform to the previously seen items and another button if they judged it to be contrary to the previously seen items. Items remained on the screen until a response was given. Half of them (legal unseen items, $n = 16$) embedded the same graphotactic constraints as the relevant exposure strings, whereas the other half (illegal items, $n = 16$) violated them. Each trial was followed by a fixation point (white cross) appearing at the center of the computer screen for 500 ms. All stimuli appeared in a single block, and no feedback was given.

Task procedures were almost identical for child and adult participants, with the following adjustments made for children. First, a brief "pre-exposure" phase was introduced, during which children were asked to practice a few button sequences in order to become familiar with the keyboard. Second, the number of practice trials was increased from 6 to 12. Third, the distraction phase consisted of two simple counting tasks (from 1 to 10 [forward] and from 20 to 1 [backward]) and five single-digit oral calculations of the form $x + y = z$, with $1 \leqslant x, y \leqslant 3$, and $2 \leqslant z \leqslant 6$.

Results

Data analyses

All analyses on reaction times (RTs) and accuracy-based measures were conducted separately. RTs below 300 ms and above 5000 ms were identified as extreme outliers and were removed from all RT analyses. RTs more than 2.5 standard deviations above or below the same participant's mean RT were trimmed to the values at 2.5 standard deviations from the mean. A criterion of $\alpha = .05$ was used for all statistical tests, and the reported p values are two-tailed unless stated otherwise. Post hoc tests were interpreted using the Bonferroni adjustment for multiple comparisons. Greenhouse–Geisser estimates were used to interpret significance for data that did not meet the requirement of sphericity (Mauchley's test of sphericity). Measures of effect size—eta squared (η^2) for analyses of variance (ANOVAs) and Cohen's d for one-sample t tests—are reported.

Visual examination of the data identified some participant outliers. One child participant who was at chance during the exposure phase of the task and four additional child participants who repeatedly pressed the same button during the test phase were dropped from all analyses. No adult participant was dropped from the analyses.

Exposure phase

It was anticipated that if the color letter detection task was a valid and effective means of exposing participants to the letter strings, children and adults would perform with high accuracy and RTs would remain relatively stable or would decrease over blocks of trials; thus, both measures would demonstrate participants' sustained attention on the task. Inspection of all groups' accuracy scores confirmed that performance in the color detection task was at ceiling in every block (>.98 for adults and >.93 for children). Thus, accuracy data were not further analyzed. To investigate whether RTs were also stable, mean correct RTs (Fig. 2) were subjected to two separate mixed ANOVAs with block (short exposure [1, 2, 3] vs. long exposure [1, 2, 3, 4, 5, 6]) as a within-participant variable and group (adults vs. children) and type of constraint (positional vs. contextual) as between-participant variables. Performance did not vary as a function of counterbalanced list; therefore, we omit further discussion of this factor.

Reaction times were stable in the short exposure variants, confirmed by a nonsignificant main effect of block, $F(1.89, 222.98) = 1.95$, $p > .05$, $\eta^2 = .02$. Block did not interact with group, $F(1.89, 222.98) = 1.57$, $p > .05$, $\eta^2 = .01$, or type of constraint, $F(1.89, 222.98) = 2.11$, $p > .05$, $\eta^2 = .02$. The three-way interaction among block, group, and type of constraint was not significant, $F(1.89, 222.98) = 2.32$, $p > .05$, $\eta^2 = .02$. The main effect of type of constraint and the group by type of constraint interaction were not significant, $Fs(1, 118) < 1$. The analyses revealed an unsurprising significant main effect of group, $F(1, 118) = 353.99$, $p < .001$, $\eta^2 = .75$, with longer RTs for children ($M = 1152.72$, $SE = 25.44$) than for adults ($M = 439.88$, $SE = 28.08$).

Some instability in RTs was observed in the long exposure variants, confirmed by a significant main effect of block, $F(3.92, 466.63) = 7.21$, $p < .001$, $\eta^2 = .05$, order 5 trend, $F(1, 119) = 30.12$, $p < .001$. The interaction between group and block was significant, $F(3.92, 466.63) = 7.00$, $p < .001$, $\eta^2 = .05$, due to a much smaller level of variation in adults, $F(3.52, 200.59) = 4.09$, $p = .005$, $\eta^2 = .07$, cubic trend, $F(1, 57) = 6.30$, $p = .015$, than in children, $F(3.91, 250.12) = 8.07$, $p < .001$, $\eta^2 = .11$, order 5 trend, $F(1, 64) = 38.39$, $p < .001$. For example, post hoc paired t tests revealed a significant difference between Block 2 and Block 3 RTs in children, $t(64) = 4.29$, $p < .001$, $d = 0.53$, but not in adults, $t(57) = 0.37$, $p > .05$, $d = 0.05$. Critically, neither group showed a systematic increase in RT; this, in combination with

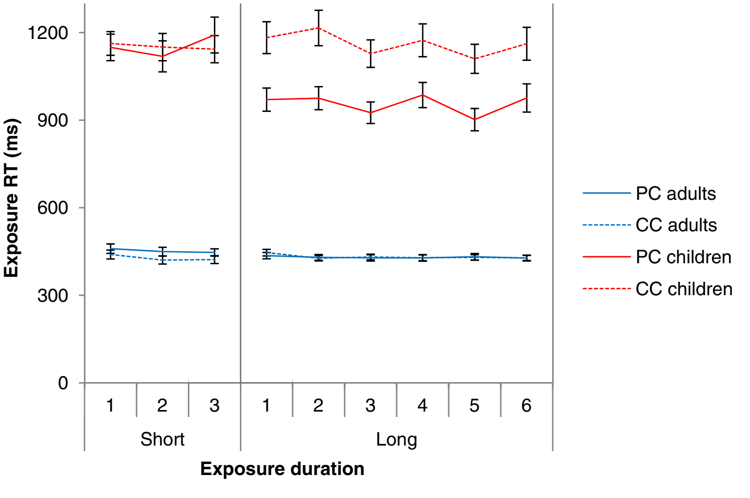

Fig. 2. Exposure RTs. Mean correct exposure RTs (ms) as a function of block are plotted separately for each experimental group in the short and long exposure variants of the incidental graphotactic learning task. Error bars represent standard errors of the means. PC, positional constraints; CC, contextual constraints.

the very high accuracy scores, suggested that both participant groups remained focused on task. The main effect of type of constraint was significant, $F(1, 119) = 9.17$, $p = .003$, $\eta^2 = .02$, and was qualified by a significant group by type of constraint interaction, $F(1,119) = 8.90$, $p = .003$, $\eta^2 = .02$. This revealed that child participants were faster during the exposure phase of the positional constraints variant ($M = 956.43$, $SE = 39.13$) than of the contextual constraints variant ($M = 1162.40$, $SE = 50.98$), $t(63) = 3.24$, $p = .002$, $d = 0.80$, whereas no significant difference was found for adults, $t(56) = 0.11$, $p > .05$, $d = 0.03$. Importantly, because the task demands for the color detection task were identical across conditions, we interpret this significant difference to reflect random child group differences in RTs. As in the previous analyses, there was a large significant main effect of group, $F(1, 119) = 335.25$, $p < .001$, $\eta^2 = .71$, such that children ($M = 1058.95$, $SE = 23.56$) were slower than adults ($M = 431.05$, $SE = 24.92$). No other interaction was significant (all *F*s < 1). In sum, adults performed the letter detection task more quickly and with greater stability in RT; however, all groups perfomed with very high levels of accuracy, confirming their sustained attention during the exposure manipulation.

Test phase

Discrimination accuracy data were subjected to signal detection theory analyses (Snodgrass & Corwin, 1988). To correct for any response bias toward rejecting or accepting items, a measure of legality sensitivity (d' prime) was calculated for each experimental condition. The d' scores were computed by calculating the difference between the z-transformed proportion of correct "yes" responses to legal sequences (hit rate) and the z-transformed proportion of incorrect "yes" responses to illegal sequences (false alarm rate). Hit or false alarm rates of 0 were replaced with $1/2n$, where n corresponds to the number of signal or noise trials, respectively. Rates of 1 were replaced with $1 - 1/2n$ (Macmillan & Kaplan, 1985). Response bias was measured with the criterion score c, calculated by averaging z(hits) and z(false alarms) and multiplying the result by -1.

Discrimination accuracy

All data points were included for analyses of discrimination accuracy in adults' data. For children's data, responses associated with latencies below 300 ms were identified as extreme outliers and discarded from these analyses. These represented only a small number of cases in each experimental condition (all $ns \leqslant 6$), and removing them did not affect the qualitative pattern of results. Table 1 presents the mean proportion of items endorsed as legal as a function of stimulus type (illegal vs. legal unseen items) for adults and children in each experimental condition. The resulting signal detection theory measures (d' and c values) are shown in the right panel of Table 1.

Mean d' values were subjected to a three-way ANOVA with group (adults vs. children), type of constraint (positional vs. contextual), and exposure duration (short vs. long) as between-participant variables. Performance did not vary across lists, and this factor was not considered further. A significant main effect of group was obtained, $F(1, 237) = 9.53$, $p = .002$, $\eta^2 = .03$, with higher d' values for adults ($M = 0.70$, $SE = 0.05$) than for children ($M = 0.47$, $SE = 0.05$), and a highly significant main effect of type of constraint, $F(1, 237) = 110.96$, $p < .001$, $\eta^2 = .30$, with higher d' values for positional constraints learning ($M = 0.97$, $SE = 0.05$) than for contextual constraints learning ($M = 0.19$, $SE = 0.05$). Group and type of constraint did not interact, $F(1, 237) = 2.09$, $p > .05$, $\eta^2 = .01$. There was no effect of exposure duration, $F(1, 237) = 1.56$, $p > .05$, $\eta^2 = .00$, exposure duration by group interaction, $F(1, 237) < 1$, or group by type of constraint by exposure duration interaction, $F(1, 237) < 1$. The type of constraint by exposure duration interaction was marginally significant, $F(1, 237) = 4.01$, $p = .046$, $\eta^2 = .01$, due to a trend for better performance in the positional constraints long exposure condition ($M = 1.08$, $SE = 0.10$) than in the positional constraints short exposure condition ($M = 0.83$, $SE = 0.08$); however, the Bonferroni-corrected post hoc t test failed to reach significance, $t(121) = 1.93$, $p = .056$, $d = 0.35$. For contextual constraints learning, there was no difference between learning following short exposure ($M = 0.21$, $SE = 0.05$) and learning following long exposure ($M = 0.16$, $SE = 0.05$), $t(120) = 0.65$, $p > .05$, $d = 0.12$.

One-sample t tests confirmed that children and adults correctly classified strings above chance in the positional constraints condition (children: $d' = 0.81$, $SE = 0.09$, $d = 1.14$; adults: $d' = 1.14$, $SE = 0.09$,

$d = 1.63$), $t_1(66) = 9.32$, $p < .001$, and $t_2(55) = 12.21$, $p < .001$, as well as in the contextual constraints condition (children: $d' = 0.13$, $SE = 0.06$, $d = 0.29$; adults: $d' = 0.25$, $SE = 0.05$, $d = 0.68$), $t_1(64) = 2.35$, $p = .022$, and $t_2(56) = 5.15$, $p < .001$; effect sizes were large for positional constraints learning and small to medium for contextual constraints learning.

Mean *c* values were also subjected to a three-way ANOVA with group (adults vs. children), type of constraint (positional vs. contextual), and duration of exposure (short vs. long) as between-participant variables. There was a main effect of group, $F(1, 237) = 13.24$, $p < .001$, $\eta^2 = .05$, such that children ($M = 0.19$, $SE = 0.04$) showed a larger response bias than adults ($M = 0.00$, $SE = 0.04$), but no effect of type of constraint, $F(1, 237) = 2.34$, $p > .05$, $\eta^2 = .01$, or exposure duration, $F(1, 237) < 1$. There was no interaction of group by type of constraint, $F(1, 237) = 1.06$, $p > .05$, $\eta^2 = .00$, group by exposure duration, $F(1, 237) = 2.02$, $p > .05$, $\eta^2 = .01$, type of constraint by exposure duration, $F(1, 237) = 3.58$, $p > .05$, $\eta^2 = .01$, or type of constraint by exposure duration by group, $F(1, 237) < 1$. The detection theory analysis revealed a significant bias on the part of child participants toward responding "no", $t(131) = 4.99$, $p < .001$, $d = 0.43$, and no response bias among adult participants, $t(112) = 0.13$, $p > .05$, $d = 0.01$.

Latency analyses

Trimmed data accounted for less than 6% of all correct test phase responses in adult participants and less than 7% in child participants. Adults' and children's mean correct trimmed RTs for legal and illegal items are shown separately for each experimental condition in Table 2. These were subjected to a three-way ANOVA with group (children vs. adults), type of constraint (positional vs. contextual), and duration of exposure (short vs. long) as between-participant variables and legality (legal vs. illegal) as a within-participant variable.

All four main effects were significant—legality, $F(1, 236) = 7.17$, $p = .008$, $\eta^2 = .03$; group, $F(1, 236) = 5.75$, $p = .017$, $\eta^2 = .02$; type of constraint, $F(1, 236) = 6.24$, $p = .013$, $\eta^2 = .02$; and exposure duration, $F(1, 236) = 12.83$, $p < .001$, $\eta^2 = .05$—with longer RTs for test strings in the short exposure condition ($M = 1598.49$, $SE = 46.02$) than in the long exposure condition ($M = 1366.54$, $SE = 45.55$). The group by legality interaction (Fig. 3) was significant, $F(1, 236) = 8.37$, $p = .004$, $\eta^2 = .03$, such that an RT advantage for legal over illegal items was found for adults, $t(112) = 4.75$, $p < .001$, $d = 0.45$, but not for children, $t(130) = 0.16$, $p > .05$, $d = 0.01$. Furthermore, the interaction indicated that children's RTs for legal items were significantly slower than adults' legal RTs, $t(242) = 3.12$, $p = .002$, $d = 0.40$. RTs for illegal items did not differ between groups, $t(242) = 1.20$, $p > .05$, $d = 0.15$.

A group by type of constraint interaction was also observed, $F(1, 236) = 5.43$, $p = .021$, $\eta^2 = .02$. As illustrated in Fig. 4, adults responded significantly faster to items in the positional constraints learning condition than in the contextual constraints learning condition, $t(105.02) = 3.27$, $p = .001$, $d = 0.62$, whereas children's RTs did not differ reliably across conditions, $t(129) = 0.39$, $p > .05$, $d = 0.07$. What is more, adults were faster than children in the positional constraints learning condition, $t(121) = 3.30$, $p = .001$, $d = 0.60$, but not in the contextual constraints learning condition, $t(119) = 0.13$, $p > .05$, $d = 0.02$.

Table 2
Adults' and children's mean correct test RTs [ms (*SD*s)] for legal and illegal items in all experimental conditions.

Condition	Illegal	Legal unseen
Adults		
PC short	1508.49 (470.91)	1303.04 (459.46)
CC short	1701.55 (546.64)	1591.24 (534.99)
PC long	1154.37 (431.19)	1028.42 (323.57)
CC long	1547.31 (596.77)	1404.74 (668.11)
Children		
PC short	1600.39 (568.04)	1695.62 (854.11)
CC short	1710.58 (581.17)	1677.01 (579.09)
PC long	1452.75 (386.30)	1470.09 (486.50)
CC long	1465.54 (528.89)	1409.13 (433.72)

Note. PC, positional constraints; CC, contextual constraints.

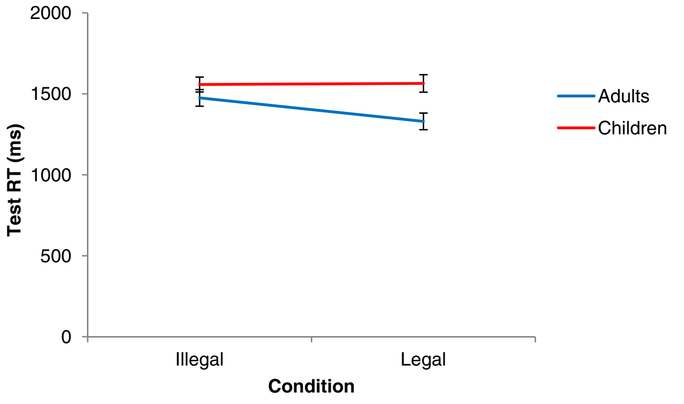

Fig. 3. Group × Legality interaction. Mean correct RTs (ms) as a function of legality are plotted separately for adult and child participants. Error bars represent standard errors of the means.

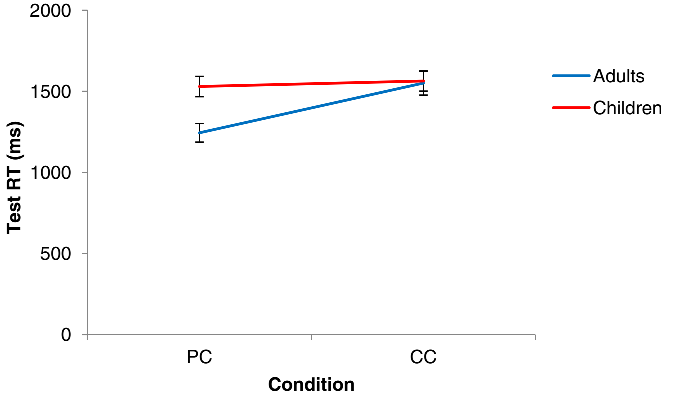

Fig. 4. Group × Type of Constraint interaction. Mean correct RTs (ms) as a function of type of constraint (PC, positional constraints; CC, contextual constraints) are plotted separately for adult and child participants. Error bars represent standard errors of the means.

The three-way interaction among group, type of constraint, and legality was not significant, $F(1, 236) = 1.79$, $p > .05$, $\eta^2 = .01$, and neither was any other interaction (all Fs < 1).

Analyses excluding lexical items

As mentioned in the Method section, although we aimed to construct lists of nonword stimuli for the learning tasks, this was not fully possible while observing the other critical stimulus constraints; consequently, a few real words were included in some of the learning lists. To explore whether the patterns of results obtained in the previous analyses were driven by the presence of real words (zero to four words per list), all lexical items were removed and the discrimination accuracy, bias *c*, and RT analyses were repeated. These follow-up analyses yielded the same patterns of results in all but one outcome on discrimination accuracy. For considerations of space, we report the analyses of bias *c* and of RTs in Appendix C, and here we report only the ANOVA on discrimination accuracy. Table 3 shows the proportion of items endorsed as legal in each experimental condition and the resulting signal

Table 3
Proportion of items endorsed as legal (and *SD*) as a function of stimulus type in each experimental condition and the resulting signal detection theory measures (analyses excluding lexical items).

Condition	*n* (m;f)	Endorsement rates[a]		SDT measures	
		Illegal (FA)	Legal unseen (hits)	d'	Bias c
Adults					
PC short	27 (8;19)	.32 (0.13)	.73 (0.15)	1.18 (0.63)	−0.09 (0.28)
CC short	28 (7;21)	.48 (0.12)	.58 (0.12)	0.26 (0.41)	−0.08 (0.26)
PC long	29 (6;23)	.24 (0.18)	.66 (0.21)	1.31 (0.82)	0.17 (0.51)
CC long	29 (8;21)	.51 (0.17)	.58 (0.14)	0.18 (0.39)	−0.10 (0.36)
Children					
PC short	33 (18;15)	.33 (0.17)	.56 (0.19)	0.70 (0.64)	0.17 (0.47)
CC short	34 (13;21)	.40 (0.20)	.43 (0.17)	0.12 (0.49)	0.26 (0.53)
PC long	34 (20;14)	.28 (0.17)	.61 (0.20)	0.99 (0.81)	0.19 (0.46)
CC long	31 (16;15)	.44 (0.19)	.48 (0.17)	0.14 (0.48)	0.12 (0.43)

Note. m, males; f, females; FA, false alarms; SDT, signal detection theory; PC, positional constraints; CC, contextual constraints.
[a] Adults' endorsement rates were calculated based on all test phase data. Children's endorsement rates were calculated based on test phase data above 300 ms.

detection theory measures. As in the previous analyses, neither the main effect of list nor any interaction with this factor was significant (all *p*s > .05).

A comparison of the data in Tables 1 and 3 reveals broadly similar patterns between groups and conditions, suggesting that, overall, the inclusion of real words did not qualitatively affect performance, as was also confirmed by bias *c* and latency analyses (see Appendix C). However, a small systematic difference can be observed in the corresponding d' values, such that exclusion of real words seemed to strengthen legality sensitivity for positional constraint learning while reducing it slightly for contextual constraints learning (except between the contextual constraints long condition estimates in children). Statistical analysis confirmed these observations, replicating the previous finding of a main effect of group, $F(1, 237) = 9.82$, $p = .002$, $\eta^2 = .03$, and type of constraint, $F(1, 237) = 123.57$, $p < .001$, $\eta^2 = .33$. However, a small significant group by type of constraint interaction, $F(1, 237) = 3.97$, $p = .047$, $\eta^2 = .01$, was now obtained. Simple effect analyses demonstrated that, relative to children ($M = 0.84$, $SE = 0.09$), adults ($M = 1.25$, $SE = 0.10$) showed an advantage in discrimination accuracy for positional constraints learning, $t(121) = 3.00$, $p = .003$, $d = 0.54$, but not for contextual constraints learning, $t(120) = 1.10$, $p > .05$, $d = 0.20$ (adults: $M = 0.22$, $SE = 0.05$; children: $M = 0.13$, $SE = 0.06$). As in the main analyses, discrimination accuracy was not significantly affected by duration of exposure, $F(1, 237) = 1.48$, $p > .05$, $\eta^2 = .00$, and exposure duration did not interact with group, $F(1, 237) < 1$. In addition, the group by type of constraint by exposure interaction was not significant, $F(1, 237) < 1$. Furthermore, the marginal trend for type of constraint by exposure interaction was no longer significant, $F(1, 237) = 2.21$, $p > .05$, $\eta^2 = .01$.

One-sample *t* tests confirmed that both child and adult participants correctly classified strings above chance in the positional constraints condition, $t_1(66) = 9.34$, $p < .001$, and $t_2(55) = 12.70$, $p < .001$, as well as in the contextual constraints condition: child participants, $t(64) = 2.19$, $p = .033$; adult participants, $t(56) = 4.16$, $p < .001$. The corresponding effect sizes were large for both children's and adults' performance in positional constraints learning ($d_1 = 1.14$ and $d_2 = 1.70$) and were small to medium in contextual constraints learning ($d_1 = 0.27$ and $d_2 = 0.55$).

To summarize, analyses based exclusively on responses to nonword stimuli yielded essentially the same pattern of results as those in which some real words were included. The one discrepant result revealed that adults' performance worsened somewhat in contextual learning when real words were removed, bringing their sensitivity scores in line with those of children.

Discussion

The current study sought to determine whether statistical learning could be induced in the acquisition of novel orthographic constraints in 7-year-old children and skilled adult learners. Previous research has demonstrated that brief listening experience (Chambers et al., 2003, 2010; Endress &

Mehler, 2010; Onishi et al., 2002) or speaking experience (e.g., Dell, Reed, Adams, & Meyer, 2000; Taylor & Houghton, 2005) is sufficient to induce similar learning of phonotactic constraints in the auditory domain. Our study focused on two types of graphotactic constraints that vary in complexity and are similar to those often operating in the English orthography. Positional learning was induced by introducing zero-order constraints on the allowable position of letters (e.g., Set 1 consonants can appear only in word-initial positions), whereas first-order contextual constraints learning was investigated by relating information about letter identity to its neighboring vowel (e.g., Set 1 consonants are always followed by the vowel *o*).

Our results are easily summarized in three main findings. First, following a successful exposure phase, where ceiling performance by all participant groups suggested sustained attention throughout this phase, we obtained clear evidence of learning. Even though the red letter exposure task did not explicitly invite participants to learn the statistical patterns embedded in the word-like strings, adults and 7-year-old children demonstrated above-chance discrimination (d') ability at test. Learning was further demonstrated on RTs among adults (but not among children) who made correct classification responses more quickly for pattern-conforming test strings than for nonconforming test strings, a result that was consistent with Onishi and colleagues' RT demonstration in the phonotactic domain (Chambers et al., 2010; Onishi et al., 2002).

Second, as anticipated, children and adults learned the zero-order positional constraints more easily than the first-order contextual constraints, suggesting that incidental learning is moderated by the complexity of the orthographic pattern being learned and that learners are sensitive to these stimulus characteristics from at least 7 years of age. Our findings on pattern complexity are consistent with the developmental trends observed in the acquisition of orthographic patterns in tasks more closely related to reading and spelling. Simple positional patterns (e.g., deterministic patterns on the allowable position of doublets) are readily acquired and used by children who have just begun to receive formal literacy instruction (Cassar & Treiman, 1997; Pacton et al., 2002; Treiman, 1993). Context-based effects, such as the coda-to-vowel contingency studied by Treiman and Kessler (2006), take significantly longer to acquire and reliably influence children's choice of alternative spellings only after several years of schooling (Caravolas, Kessler, Hulme, & Snowling, 2005). We return to this issue below to consider in more detail how the complexity effect might have arisen in the current study and how it might operate in statistical learning more generally.

Third, we observed a developmental effect in learning performance. This manifested in children responding more slowly and somewhat more variably than adults during the exposure task and attaining slower discrimination RTs in the positional constraints condition. Moreover, in our main discrimination accuracy analysis, adults learned the informational structure embedded in the exposure strings more easily than 7-year-olds. However, although both groups learned both positional and contextual patterns reliably to above-chance levels, the adults' advantage over children proved to be less robust for learning contextual constraints; they were not faster than children in responding to test items, nor were they more accurate in this condition when lexical items were removed from the discrimination analyses (see "Analyses excluding lexical items" in Results section). The latter result tentatively suggests that adults especially benefited from the presence of real words in learning the contextual constraints. However, bearing in mind that our tasks included a relatively small pool of items and that removing items from analyses reduces statistical power, these latter analyses need to be considered with appropriate caution. Regardless, the fluctuations in d' scores across the full and follow-up analyses and the RT and discrimination accuracy results for positional constraints learning are not fully consistent with the claim that implicit learning is age invariant, as proposed by Reber (1993). Instead, our results suggest that although such learning is certainly possible across the age span (as demonstrated by above-chance d' scores across groups and conditions), *what* and *how much* can be learned seems to be influenced by the amount of experience accrued in the learning domain, a variable typically correlated with age, and by the complexity of the patterns being abstracted. Age-related differences in the domain of orthographic learning have been reported in numerous previous studies (e.g., Cassar & Treiman, 1997; Hayes et al., 2006; Treiman & Kessler, 2006) and are consistent with the view that pattern sensitivity strengthens over time as children progress toward more complete orthographic learning (Hayes et al., 2006).

Contrary to expectations, our manipulation of length of exposure did not have a robust influence on learning. Doubling the amount of exposure led to faster correct responses at test; however, it did not result in higher discrimination ability. This result is inconsistent with the idea that graphotactic sensitivity strengthens as a result of spellers' increasing exposure to print (Hayes et al., 2006). Furthermore, it conflicts with previous studies showing that longer exposure enhances incidental learning of statistical regularities in a wide range of domains (e.g., Gaillard et al., 2009) and learner groups (e.g., Saffran et al., 1997). There are a number of possible reasons for our participants' failure to benefit from the longer exposure. Perhaps the difference in length between our short versus long exposure condition was not sufficiently large to induce better learning in the long exposure condition. Alternatively, learning of complex patterns may be enhanced through the experience of a wider variety of tokens that embed the constraint or by the "little but often" approach of spaced learning (e.g., Karpicke & Roediger, 2007). More research is needed to determine the appropriate threshold in the frequency of repetitions and/or variety of tokens that would lead to reliable improvements in learning. It will also be of interest to examine specific characteristics of the learning process involved in this paradigm (e.g., its time course and durability). At a minimum, clearly, learning of statistically predictable patterns and contingencies develops quickly after only a few exposures to critical stimuli (e.g., Dell et al., 2000; Onishi et al., 2002).

We now turn to considerations of what factors other than the complexity of the constraints (zero order vs. first order) may have influenced performance in our study. For example, task demands in our exposure manipulation may have inadvertently preferentially enhanced positional constraints learning. Although the exposure task was designed to ensure that participants attended to the whole string in seeking the red letter, and the letters flanking the red target will have been in their effective visual field (e.g., Rayner, Well, & Pollatsek, 1980), attending explicitly to the red letter may have augmented learning of the single-letter positional constraints while attenuating that of the two-letter contextual constraints (see Pacton & Perruchet, 2008). In addition, it is possible that differences in the number of exposures to specific tokens in the positional constraints condition (four repetitions per token, e.g., *d* as onset) versus the contextual constraints condition (two repetitions per token, e.g., *do* body), a necessary consequence of controlling for the length of the overall exposure phase, may have biased learning in favor of positional constraints. However, mitigating against this hypothesis and corroborating the suggestion that pattern complexity is a factor in statistical learning are the results of the long exposure manipulation. Here, we found that even when doubling the number of exposures in the contextual constraints condition, such that participants saw each target instance (e.g., *do*) four times, performance did not reach the levels obtained in the positional constraints short exposure condition even though the number of instance exposures was now equated and the number of set-wide exposures was doubled.

We further considered what type of knowledge is formed during the incidental graphotactic learning task. Although our task involved only a visual presentation of letter strings, it cannot be ruled out that participants did not also covertly verbalize these and coincidentally also learned the constraints as phonotactic. However, because verbalization was neither encouraged nor necessary here, learning of graphotactics presumably took precedence over any additional incidental phonotactic learning. The extent to which graphotactic and phonotactic constraints are separable during orthographic learning is an interesting question for future research.

We also explored which stimulus characteristics may have provided additional, redundant cues and possibly contributed to learners' above-chance performance. In our stimuli, both positional and contextual constraints were effectively anchored at the outer edges of the letter strings, where learning may have benefited from their perceptual salience. Statistical learning might have been less pronounced if both types of constraints were embedded in less salient positions of longer letter strings. Along these lines, it would be interesting to tease apart the relative importance of the complexity versus the position of the pattern being learned, for example, in embedding the positionally constrained patterns in medial positions and the contextually constrained patterns in edge positions. Another source of redundancy arose from the design of the stimuli, in which the embedded constraints were reflected in more than one position across strings. For example, our positional constraints learning condition embedded constraints on single consonants occurring in the onset *and* coda positions. Therefore, it is possible that participants were learning the constraints from both units. Similarly, in the contextual learning experiments, we presented CV as well as VC first-order contingencies equally often, and it is possible that participants relied on the constraints embedded in either the body or rime or both units for their learning. In addition, learning about the legal positions of single letters may have also benefited from the presence of statistical information governing the legal position of body/rime units that occurred, although only half as often, in the positional constraints learning experiment. Spoken and written language are replete with redundancy of cues, and in this sense our stimuli may have provided a very favorable learning context. Teasing apart correlated cues in the input and evaluating their respective importance for learning is an exciting challenge for this area of orthographic learning.

Numerous studies in the artificial grammar learning literature have demonstrated that task performance is driven by several often highly correlated test item properties (Perruchet & Pacteau, 1990; Servan-Schreiber & Anderson, 1990). Sensitivity to chunk information (i.e., co-occurring elements in a string) usually quantified in terms of a string's associative chunk strength (e.g., Knowlton & Squire, 1994), chunk novelty (e.g., Meulemans & Van der Linden, 1997), and similarity to specific training items (e.g., Brooks & Vokey, 1991) are only a few alternatives to Reber's (1969) original claim that people form an abstract representation of rules. On this issue, we suggest that above-chance classification performance on our task does not necessarily entail forming an abstract representation of our graphotactic rules for the position and context of letter distributions (see Pacton et al., 2002, for a similar point). For example, positional learning may be interpreted as sensitivity to locations where not only single letters (as in our conceptualization) but also digraphs occur, whereas in the contextual constraints learning condition it can be argued that participants learned allowable letter co-occurrences (as in our conceptualization) or digraph constraints. The current study was not designed to disentangle the roles of all the metrics mentioned above, and investigations of their relative importance for learning are the focus of our ongoing research.

To conclude, the current study offers novel insights into the learning mechanisms contributing to the acquisition of graphotactic knowledge. Although several authors have argued that statistical learning processes underlie children's appreciation of simple orthographic conventions, to our knowledge this is the first study to directly assess how this learning can occur. Our data clearly demonstrate that a few minutes of incidental exposure is sufficient to induce learning of graphotactic patterns that are commonly found in most alphabetic orthographies. Learning effects in our study appear to emerge as early as the end of Year 2 (comparable in age to Grade 1 in the United States), a finding that contrasts with the widely held position that orthographic knowledge is unavailable to beginner spellers (e.g., Frith, 1985; Gentry, 1982). Admittedly, our study is limited by the artificial nature of our deterministic underlying statistics. For example, the significant contextual learning effect does not necessarily imply that children in Year 2 are already sensitive to the probabilistic, context-based patterns they encounter in real language settings. However, above-chance learning of some simple effects of context provides support to the notion that statistical learning processes operate early on and may provide the mechanism for the acquisition of at least some patterns in written language.

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Appendix A. Stimuli used in positional constraints variant of incidental graphotactic learning task.

The list was counterbalanced across participants, such that half of them were presented with List 1/2 items during exposure and List 2/1 and 3 items at test and the other half were presented with List 3/4 items during exposure and List 4/3 and 1 items at test.

List 1	List 2	List 3	List 4
dot	det	tod	ted
den	don	tom	tem
dop	dep	tel	tol
des	dos	tef	tof
met	mot	ned	nod
mon	men	nem	nom
mep	mop	nol	nel
mos	mes	nof	nef
lot	let	pod	ped
len	lon	pom	pem
lop	lep	pel	pol
les	los	pef	pof
fet	fot	sed	sod
fon	fen	sem	som
fep	fop	sol	sel
fos	fes	sof	sef

Note. In one of the four counterbalanced list conditions, *den* was presented during exposure, *don* as a legal unseen test item, and *tom* as an illegal test item.

Appendix B. Stimuli used in contextual constraints variant of incidental graphotactic learning task

List was counterbalanced across participants, such that half of them were presented with List 1/2 items during exposure and List 2/1 and 3 items at test and the other half were presented with List 3/4 items during exposure and List 4/3 and 1 items at test.

List 1	List 2	List 3	List 4
dot	don	det	den
dop	dos	dep	des
mon	mot	men	met
mos	mop	mes	mep
lot	lon	len	let
lop	los	les	lep
fon	fot	fet	fen
fos	fop	fep	fes
tem	ted	tod	tom
tef	tel	tol	tof
ned	nem	nom	nod
nel	nef	nof	nol
pem	ped	pom	pod
pef	pel	pof	pol
sed	sem	sod	som
sel	sef	sol	sof

Note. In one of the four counterbalanced list conditions, *dop* was presented during exposure, *dos* as a legal unseen test item, and *dep* as an illegal test item.

Appendix C. Analyses of bias *c* and latency excluding lexical items

Removing real words from the lists resulted in lists comprising 12, 13, 15, and 16 stimuli for Lists 1, 2, 3, and 4, respectively, for positional constraints learning and comprising 14, 15, 15, and 12 stimuli for Lists 1, 2, 3, and 4, respectively, for contextual constraints learning. To clarify, bearing in mind that participants were only ever tested on two lists (see Appendixes A and B), participants who were exposed to List 1 in positional constraints learning, for example, were tested on 13 legal nonword items (List 2) and 15 illegal nonword items (List 3).

The mean *c* values (Table 3) were subjected to a three-way ANOVA, as was done for the full data set. We observed a significant main effect of group, $F(1, 237) = 14.74$, $p < .001$, $\eta^2 = .06$, such that children ($M = 0.19$, $SE = 0.04$) tended to show a larger response bias than adults ($M = -0.03$, $SE = 0.04$). As in the previous analysis, there was no significant effect of type of constraint, $F(1, 237) = 1.26$, $p > .05$, $\eta^2 = .00$, or exposure duration, $F(1, 237) < 1$, on mean *c* values. There was no interaction of group by type of constraint, $F(1, 237) = 1.74$, $p > .05$, $\eta^2 = .01$, group by exposure duration, $F(1, 237) = 2.78$, $p > .05$, $\eta^2 = .01$, or type of constraint by exposure by group, $F(1, 237) < 1$. The ANOVA revealed a significant interaction of type of constraint by exposure, $F(1, 237) = 4.01$, $p = .046$, $\eta^2 = .02$. The difference between *c* values for positional constraints learning ($M = 0.18$, $SE = 0.05$) and those for contextual constraints learning following long exposure ($M = 0.01$, $SE = 0.06$), $t(121) = 2.09$, $p = .039$, $d = 0.38$, was not significant after Bonferroni correction. There was no difference between *c* values in the two conditions following short exposure, $t(120) = 0.66$, $p > .05$, $d = 0.12$. In summary, consistent with the analysis that included lexical items, *c* values differed only as a function of group. One-sample *t* tests replicated the finding of children's bias, but not adults' bias, toward responding "no", $t_1(131) = 4.55$, $p < .001$, $d = 0.40$, and $t_2(112) = 0.68$, $p > .05$, $d = 0.06$.

Adults and children's RTs as a function of legality are shown separately for each experimental condition in Table C1.

Table C1. Adults' and children's mean correct test RTs [ms (*SD*s)] for legal and illegal items in all experimental conditions (analyses excluding lexical items)

Condition	Illegal	Legal unseen
Adults		
PC short	1508.03 (470.70)	1314.76 (452.94)
CC short	1751.10 (587.17)	1592.10 (548.49)
PC long	1157.96 (433.37)	1046.08 (329.18)
CC long	1557.96 (621.61)	1417.76 (677.28)
Children		
PC short	1616.70 (577.45)	1723.38 (867.82)
CC short	1712.68 (577.83)	1717.36 (607.11)
PC long	1418.53 (379.97)	1426.10 (467.07)
CC long	1477.01 (553.11)	1430.11 (436.99)

Note. PC, positional constraints; CC, contextual constraints.

Replicating the previous set of analyses, we obtained a significant main effect of legality, $F(1, 236) = 6.18$, $p = .014$, $\eta^2 = .02$, a significant main effect of group, $F(1, 236) = 4.98$, $p = .027$, $\eta^2 = .02$, a significant main effect of type of constraint, $F(1, 236) = 7.52$, $p = .007$, $\eta^2 = .03$, and a significant main effect of exposure duration, $F(1, 236) = 14.47$, $p < .001$, $\eta^2 = .05$, with faster RTs for items in the long exposure variants ($M = 1366.44$, $SE = 46.33$) than in the short exposure variants ($M = 1617.01$, $SE = 46.81$). As in the analyses carried out on all test phase items, the group by legality interaction was significant, $F(1, 236) = 9.98$, $p = .002$, $\eta^2 = .04$, showing that adults were faster with legal items ($M = 1340.96$, $SE = 51.69$) than with illegal items ($M = 1491.23$, $SE = 53.63$), $t(112) = 4.57$, $p < .001$, $d = 0.43$, but children were not (legal RTs: $M = 1575.31$, $SE = 55.06$; illegal RTs: $M = 1556.39$, $SE = 46.65$), $t_2(130) = 0.47$, $p > .05$, $d = 0.04$. Children's RTs for legal items were significantly slower than those of adults, $t(242) = 3.07$, $p = .002$, $d = 0.40$, whereas no group difference was observed with regard to RTs for illegal items, $t(242) = 0.92$, $p > .05$, $d = 0.12$. The significant group by type of constraint interaction also was replicated, $F(1, 236) = 4.68$, $p = .032$, $\eta^2 = .02$, showing adults' faster RTs to items in the positional constraints condition ($M = 1260.98$, $SE = 58.33$) versus the contextual constraints condition ($M = 1569.88$, $SE = 75.86$), $t_1(104.58) = 3.23$, $p = .002$, $d = 0.61$, but not children's (positional learning: $M = 1522.93$, $SE = 64.07$; contextual learning: $M = 1580.80$, $SE = 64.37$), $t_2(129) = 0.64$, $p > .05$, $d = 0.11$. Children were slower than adults in the positional constraints learning condition, $t(121) = 2.97$, $p = .004$, $d = 0.54$, but not in the contextual constraints learning condition, $t(119) = 0.11$, $p > .05$, $d = 0.02$. None of the remaining interactions was significant ($ps > .05$).

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