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LOYOLA UNIVERSITY CHICAGO

YOU HAD ME AT HELEN:  
THE NAME LETTER EFFECT IN JUDGMENTS OF HUMOR

A DISSERTATION SUBMITTED TO  
THE FACULTY OF THE GRADUATE SCHOOL  
IN CANDIDACY FOR THE DEGREE OF  
DOCTOR OF PHILOSOPHY

PROGRAM IN APPLIED SOCIAL PSYCHOLOGY

BY

JENNA FINWALL RYAN

CHICAGO, ILLINOIS

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This is dedicated to Jack, Jack, Jr., Jakobi, Jan, Jane, Janel, Jayne,  
Jeff, Jeramie, Jimmy, Joe, John, Josie, Joyce, Judy, and Julia.

Humor can be dissected as a frog can, but the thing dies in the process  
and the innards are discouraging to any but the pure scientific mind.

—E. B. White, author of *Charlotte's Web*

## PREFACE

During the final season of the TV series *Seinfeld*, Elaine dramatically stormed down to *The New Yorker* editor's office to demand an explanation for an ambiguous cartoon, only to find there wasn't one: the editor admitted he merely "liked the kitty." The episode, entitled "The Cartoon" and written by real-life *New Yorker* cartoonist Bruce Eric Kaplan, is undoubtedly a humorous illustration of art imitating life.

In an example of life imitating *art*, current cartoon editor Robert Mankoff was also recently asked to explain a cartoon that appeared in *The New Yorker* (Mankoff, 2012). He cites E. B. White's famous quote as a disclaimer before attempting to quantify the otherwise unquantifiable experience of comedy, warning readers that dissecting the cartoon would likely spoil its joke—and perhaps even all humor in general. Reluctantly, Mankoff ranks ambiguity and incongruity high among humor's essential ingredients.

This year, University of California San Diego psychologists joined forces with Mankoff to more scientifically examine comedy in a study of gender differences, revealing what some may argue is another key feature of humor. Pitting women's abilities against men's in the creation of humorous captions for a selection of *New Yorker* cartoons, men very slightly but significantly outperformed women in the comedy department, thereby providing a kernel of truth to the stereotype that men are funnier than women. And in a second study, which perhaps serves to perpetuate or even explain this stereotype, both men *and* women misremembered the more humorous captions as having been written by men (Mickes, Walker, Parris, Mankoff, & Christenfeld, 2012).

Previous psychologists who were also interested in examining what makes cartoons funny employed a novel and implicit manipulation of the “facial feedback hypothesis.” According to the hypothesis, emotions can begin on the *outside* and affect how we feel *inside* (Buck, 1980). In an effort to simulate emotions as outside-in, Strack, Martin, and Stepper (1988) contracted either participants’ “smiling muscles” or their “frowning muscles” to see if these artificially-originated facial expressions would influence participants’ experience of humor when assigning ratings to cartoons.

To achieve this, the researchers had participants grasp a pen between their teeth (producing a smile) or between their lips—creating a frown. Unlike previous scientists’ recreations of emotions through muscle and electrode stimulation, Strack et al.’s (1988) participants performed the cartoon rating exercise under the guise of a pilot study on coordination to help gain insight into workarounds for people with physical impairments. This cover story ensured that participants would be unaware that their smiling and frowning muscles had been implicitly activated and prevented participants’ realization that these activations might impact their humor ratings of cartoons.

It should not be surprising that Strack et al. (1988) found that participants rated cartoons as funnier when holding a pen between their teeth than those who held a pen between their lips. Moreover, the authors demonstrated that this effect occurred in the absence of participants’ awareness that these facial expressions had even been activated. In other words, judgments of humor were implicitly increased under the “facilitating condition” of a smile and implicitly decreased under the “inhibiting condition” of a frown. These results were replicated in a second study and, importantly, were limited to ratings made by participants who were instructed to rely on their affective and subjective

reactions to the cartoons vs. those who were asked to arrive at objective, cognitive-based humor ratings.

Around the same time, a Belgian researcher observed that attractiveness ratings for letters of the alphabet were also influenced by an implicit process that operated outside of participants' awareness: their positive self-associations for name letters (Nuttin, 1985; 1987). In other words, judgments of letter attractiveness were implicitly increased under the facilitating condition of shared letter similarity. In the years that followed, this "name letter effect" was found to extend to both mundane and important decision-making, while moderators that served to bolster or attenuate the effect were also discovered and still continue to be identified.

Drawing inspiration from the findings of Strack et al.'s (1988) humor judgments experiment and research on the name letter effect, the author of the present study wondered whether the "facilitating condition" of shared first name initials with cartoon caption writers would implicitly increase participant judges' humor ratings for these writers' captions in a "mock" *New Yorker* cartoon caption contest. Would participants' positive self-associations for their own name letters unwittingly influence their evaluations of cartoon captions written by writers who shared their initial letter, relative to captions written by writers with dissimilar initial letters and as compared to evaluations made by non-initial matching participants? If found, would this effect be influenced by both old and new moderating variables? The present study sought to answer these and other questions about the name letter effect in judgments of humor. To this pure scientific mind, dissecting the experience of humor in search of its implicit egotism innards would prove quite interesting indeed.

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## ABSTRACT

The present study demonstrates that implicit egotism is relevant to not only letter attractiveness ratings on the Name Letter Test (NLT), but also to judgments of humor—albeit to a lesser degree. Respondents participated as “mock” judges in a simulated cartoon caption contest and evaluated writers’ caption submissions for two cartoons. It was hypothesized that participants would exhibit biases toward captions submitted by writers with whom they shared a first initial letter, and additionally, their gender. A name letter effect was found in participants’ judgments of humor and on the NLT. Shared gender with a caption writer—when coupled with a shared initial—increased biases toward these writers’ captions, but not significantly so. The impact of implicit self-esteem on initial-letter biases was examined, with level of implicit self-esteem weakly predicting NLT biases, but not biases demonstrated toward captions submitted by same-initial writers. While name-letter preferences are believed to tap implicit self-esteem, less than one-third of participants demonstrated high implicit self-esteem, despite the very large name letter effect observed on the NLT. This challenges the notion that people overwhelmingly possess the positive self-attitudes thought to ignite implicit egotism. Recent researchers have suggested that the NLT is best understood as a measure for which it was first designed—implicit egotism, the tendency to display automatic self-positivity biases toward targets that share our self-attributes—instead of a measure of implicit self-esteem. This possibility is discussed and explored with analyses of the relationships between the NLT, explicit self-esteem, and implicit self-esteem.

## CHAPTER 1

### INTRODUCTION

By most standards, psychologists have only somewhat recently begun to examine people's profound affection for the attribute that is most closely related to the self, i.e., one's own name—a principle that human relations guru Dale Carnegie taught his students more than three quarters of a century ago. Carnegie (1936) asserted that “there is no sweeter sound than one's own name” almost 50 years before Nuttin (1985, 1987) would experimentally demonstrate people's preferences for their name letters over non-name letters in the social psychological phenomenon known as “the name letter effect.”

Researchers have since found name-letter preferences to extend beyond just mundane decision-making. This automatic and unconscious bias appears to manifest itself in not only everyday decisions, but important ones as well such as people's preferences for places of residence, careers, and even mates. Because choosing a place to live, career, or life partner because of shared name letters is unlikely to reflect a conscious decision, researchers have labeled this phenomenon “implicit egotism” (Pelham, Mirenberg, & Jones, 2002). The study of implicit egotism has predominantly focused on people's profound affection for their own name letters. However, some researchers argue that virtually any self-attribute, including something as trivial as shared birthdates or birthday numbers (Kityama & Karasawa, 1997; Pelham et al., 2002; see also Finch & Cialdini, 1989), can foster implicit egotism.

Interesting paradigms have been used to study the seemingly illogical yet predictable decisions we make as a result of our implicit positive biases we have toward our name letters and other self-attributes. In a series of ten different studies, Pelham et al. (2002) demonstrated that important life decisions can fall prey to implicit egotism. First, they discovered that people are disproportionately more likely to live in cities and states that resemble their names. Specifically, U.S. citizens were more likely to move to or reside in places that shared several letters of their first or last name (Louis's disproportionately populate St. Louis and Louisiana) or places that resembled their birthday numbers (a disproportionate number of people born on May 5<sup>th</sup> live in Five Points, Alabama). Follow-up studies have found similar effects where individuals' surnames match the street on which they choose to live (Pelham, Carvallo, DeHart, & Jones, 2003).

Another major life decision influenced by implicit egotism is that of career choice. Pelham et al. (2002) found that a disproportionate number of people whose names begin with "Den" (e.g., Denis and Denise) make their livelihoods as dentists, while names beginning with "La" (e.g., Larry and Laura) are overrepresented within the law profession. Similarly, authors of scholarly articles in the geosciences are more likely than chance to be named George and Geo<sup>ff</sup>rey.

But perhaps the most significant life decision studied by implicit egotism researchers to date is that of mate selection. Jones, Pelham, Carvallo, and Mirenberg (2004) found evidence that brides and grooms gravitate toward spouses whose first names share letters with their own. These authors also observed surname-matching effects for married couples, despite our culture's strong taboos against incest.

In sum, implicit egotism researchers have found the name letter effect to be quite robust, extending to both everyday and important decisions. Moreover, name-letter preferences are not merely an American phenomenon, but rather, have been demonstrated in other countries and cultures as well (Dijksterhuis, 2004; Kityama & Karasawa, 1997, Hoorens & Todorova, 1988; Nuttin, 1987). Researchers have examined several different possible mediating mechanisms underlying the name letter effect, each of which will be described below.

### **What Mediates the Name Letter Effect?**

#### **The Primacy Effect**

One explanation for name-letter preferences that has since been discounted is “the primacy effect.” Researchers Hoorens and Todorova (1988) compared Bulgarian students’ preferences for their name letters in their native and second languages. They hypothesized that if the name letter effect is simply due to a primacy effect of the name letters these students learned first in their native language, then they should not prefer their name letters over non-name letters in a later-learned second language. Preferences for name letters, however, were found for *both* the Bulgarians’ native Cyrillic alphabet and within the Roman alphabet they learned in a second, more recently acquired language (Hoorens & Todorova, 1988).

#### **The Mere Exposure Effect**

The “mere exposure effect” suggests that we prefer things with which we are familiar, therefore repeated exposure to them can increase both familiarity and ultimately liking (Zajonc, 1968). Moreover, instead of being a cognitive or post-cognitive phenomenon, this effect appears to be a purely affective experience, occurring

independently of any intervening cognitive processes (Zajonc, 1980). If we encounter our name letters more frequently than the other letters of the alphabet, then according to the mere exposure effect, perhaps this is why we prefer them.

Kitayama and Karasawa (1997), however, ruled out the mere exposure effect as a possible mediator of preferences for numbers that make up one's birthday, a form of implicit egotism similar to name-letter preferences. Participants in their study exhibited stronger preferences for birth date numbers 13-31 than the numbers 1-12, despite encountering the numbers 1-12 more frequently in everyday life (Kitayama & Karasawa, 1997). These findings demonstrate the link between number preferences and our positive bias toward objects related to the self, while simultaneously ruling out mere exposure as a mediating variable.

In another study, researchers Jones, Pelham, Mirenberg, and Hetts (2002) found that people prefer six of the most frequently used letters of the alphabet (A, E, I, N, S, and T) to the six least often used letters (J, K, Q, W, X, and Z). These preferences, however, were overshadowed by biases toward name letters, suggesting again that the name letter effect is a product of something much more than mere exposure. In fact, people with rare initials preferred their initial letters more than two of the most frequently occurring letters, E and S (Jones et al., 2002). These findings staunchly challenge mere exposure as a mediator of name-letter biases. If name-letter preferences were in fact due to the mere exposure effect, these preferences should be much *less* pronounced for people with uncommon initials since they are not exposed to these letters very frequently.

Finally, two other studies in which ownership of name-related and non-name-related "symbols" were induced provide additional evidence against mere exposure's role

in self-related preferences (Feys, 1991). In these experiments, participants were equally exposed to novel symbols that they were led to believe either belonged to their own or someone else's name. Name-related biases still occurred, i.e., participants preferred symbols associated with their own name more than those associated with another person's name without having been exposed to their name-related symbols more frequently (Feys, 1991).

### Evaluative Conditioning

Having failed to find support for mere exposure as a determinant of name-related preferences, Feys (1995) set out to test another hypothesis that might explain the name letter effect. The author was intrigued by the evidence he found that mere ownership can be temporarily induced in a laboratory setting and also that this newly-found ownership was strong enough to create preferences for symbols associated with the self. Moreover, this effect did not appear to be a function of classical conditioning since the author controlled for the duration of symbol induction (the number of trials it took participants to learn self- vs. other-related symbols) *or* mere exposure since participants were exposed equally to self- vs. other-related symbols. Feys (1995) next decided to examine the role “evaluative conditioning” might play in self-related preferences in a follow-up study.

Since implicit egotism embodies the positive biases we have toward objects related to the self, it implies a mediating relationship between exposure to these objects and our preferences for them. As such, it is believed to elicit an affective—albeit implicit—response to these attributes or objects en route. “Evaluative conditioning,” on the other hand, is not considered a “true” mediator since it occurs automatically without any affective or cognitive processing. For example, very early research by Syz (1926)

demonstrated that of fifty different stimuli presented, the one that achieved the most frequent (84% of the time) automatic galvanic skin responses for participants was his or her own name. For these reasons argues Fey (1995), evaluative conditioning—if supported—would be a much more “parsimonious” explanation than the mere ownership hypothesis because it is purely automatic and occurs without any intervening processes.

To examine evaluative conditioning as a potential mediator of name-related preferences, Feys (1995) visually paired Japanese Kanji symbols with participants’ name letters and non-name letters. In a control condition, participants were instructed to simply remember which symbols corresponded to their own versus another’s name without the aid of “visual pairing” (seeing the Kanji symbol alongside Roman alphabet letters). In both conditions, participants preferred the symbols associated with his or her own name more strongly than symbols associated with another person’s name. However, this effect was even stronger in the non-evaluative conditioning (no visual pairing) control condition where participants *simply remembered* that the symbols were name-related. These results led Feys (1995) to conclude that evaluative conditioning is not necessary to elicit name-related preferences (in his experiment, simply knowing the symbols were name-relevant was sufficient) and that “mere ownership” is a more likely determinant. Mere ownership, while similar to implicit egotism, differs in that it often embodies an *explicit* awareness that a stimulus object is “mine.” Implicit egotism, on the other hand, occurs outside of our conscious awareness, i.e., is driven by an implicit process or unconscious mechanism.

### Subjective Frequency

Researchers Hoorens and Nuttin (1993) tested the hypothesis that the name letter effect might be attributable to an exaggerated subjective frequency of exposure to one’s

name letters. As Nuttin (1985; 1987) was the first researcher to identify the name letter effect, the purpose of his and Hoorens's 1993 study was to test an alternative hypothesis that might explain this phenomenon. Previous researchers had already demonstrated that people's affective reactions to objects might be driven by their perceived (vs. their actual) familiarity with them (Matlin, 1971; Moreland & Zajonc, 1977, 1979). Based on this principle, Hoorens and Nuttin (1993) considered that name letters might be preferred simply because people overestimate their exposure to them and thus their familiarity with them.

While subjective frequency for name letters was somewhat exaggerated, the authors found no relationship between name-letter preferences and participants' reports of their subjective frequency of exposure to these letters (Hoorens & Nuttin, 1993). The authors did, however, find evidence of a name letter effect, which was strengthened when participants believed that their name a) would suit someone they admired from the same gender, b) would suit a member of a professional group to which they would like to belong, and c) carried a strong likelihood of being chosen as a name for themselves if they had had the opportunity to so choose.

### Implicit Self-Esteem

Intrigued by their discovery that numerous Dutch nationals still kept their ancestors' surnames that were chosen out of rebellion and carried unfavorable translations, Koole and Pelham (2003) also decided to investigate the underpinnings of the name letter effect. The authors believed that the Dutch's commitment to these made-up names suggested an underlying implicit favorable self-attitude that spilled over into a fondness for their surname self-attribute. After performing a comprehensive review of



studies on the name letter effect, they argue that name-letter preferences are most likely driven by one's implicit self-esteem for the following reasons. First, name letters represent a self-attribute. Secondly, letter preferences largely reflect a *positive* bias toward own-name letters. Thirdly, these preferences operate outside of one's awareness ("implicitly"), and lastly, no other factors can account for the numerous studies which have consistently found biases toward name letters (Koole & Pelham, 2003).

Since then, researchers have begun to explore the relationship between self-esteem and name-letter preferences. After subjecting participants to self-esteem-enhancing conditioning, Dijksterhuis (2004) found stronger name letter effects for these participants compared to controls. By subliminally pairing the word "I" with positive trait terms, the author was able to temporarily increase both implicit self-esteem and name-letter preferences (Dijksterhuis, 2004). Baccus, Baldwin, and Packer (2004) found a similar increase in implicit self-esteem and name-letter preferences in an experiment that paired self-relevant information with "socially-approving" happy faces. Finally, Jones et al. (2004) demonstrated that temporary threats to the self-concept increased attraction to a potential dating partner whose screen name shared letters of participants' names. Taken together, these studies appear to suggest at least some underlying component of self-esteem in name-letter biases.

### Implicit Egotism

While often used interchangeably with "implicit self-esteem," *implicit egotism* refers uniquely to our preferences for self-related objects, which are driven by our unconscious positive self-biases (Pelham et al., 2005). Whether these automatic positivity biases that are overwhelmingly displayed in peoples' preferences for self-related objects

suggest an equally overwhelming and underlying incidence of favorable implicit self-esteem is debatable. Recent researchers Buhrmester, Blanton, and Swann, (2011) argue that two widely used measures of implicit self-esteem, the Implicit Association Test (IAT; Greenwald & Farnham, 2000) and the Name Letter Test (NLT; Nuttin, 1985; 1987), *do not* measure participants' implicit global evaluations of the self. Rather, these authors believe that the IAT is a more appropriate measure of implicit affect and the NLT is best understood as a measure of the tendency to display automatic self-positivity biases, i.e., “implicit egotism”—instead of a tool to measure implicit self-esteem (Buhrmester et al., 2011).

At first, some researchers questioned the “implicitness” of their own findings after discovering the strongest effects occurred when entire names matched a target (e.g., a state where one chooses to live), however, their follow-up study provided irrefutable evidence that implicit egotism is indeed implicit. In their self-described “strictest test of implicit egotism” to date, these authors found men participants demonstrated increased attraction to a photograph of a woman after subliminally pairing her football jersey number with these participants' names (Jones et al., 2004).

Even outside of the psychology laboratory, consciously choosing a career, residence, or mate because of shared name letters is ultimately not only foolhardy, but improbable. Regardless of the mechanism underlying name-letter preferences, it seems as though most researchers agree that it is implicit. Pelham et al. (2005) point out that significant life decisions regarding careers, places of residence, and mate selection are highly unlikely to be a product of “explicit egotism” when these choices share letters with one's own name. And finally, in all of the above experiments studying name-letter

preferences, researchers very seldom ever reported that participants were able to guess the hypotheses of their studies. This further supports the notion that our bias toward our name letters operates outside of our awareness.

### **Moderators of the Name Letter Effect**

The identification of moderators of the name letter effect has further strengthened the evidence that our positive (in valence, but not necessarily in content) self-biases are indeed responsible for name-letter preferences. In addition to previous researchers' findings that unique name letters (Jones et al., 2002; Pelham et al., 2002) and self-esteem enhancement (Dijksterhuis, 2004; Baccus et al., 2004) can intensify the name letter effect, other variables such as gender (Pelham et al., 2002; Kitayama & Karasawa, 1997) and self-concept threats (Brendl, Chattopadhyay, Pelham, & Carvallo, 2005; Jones et al., 2004, 2002) have also been found to differentially influence preferences for name letters.

#### **Uniqueness of Name**

Having an uncommon name serves to strengthen the name letter effect and—as mentioned previously with respect to rare initials—simultaneously discounts mere exposure as a determinant of name-letter preferences (Jones et al., 2002; Pelham et al., 2002). For example, one study found that people with unique names exhibited stronger biases than people with more common names as evidenced by the disproportionate number of people residing in states that resembled their own (uncommon) name (Pelham et al., 2002). This finding again challenges mere exposure as an explanation for the effect because the mere exposure hypothesis argues that people develop preferences after repeated exposure to the object or stimulus. However, people with unique names do not

encounter their names with the same frequency as do those with more common names, therefore biases toward name-letter cannot be explained by mere exposure.

### Gender Differences

In addition to finding “letter position” effects, i.e., participants preferred the first letter of their names (initial) more strongly than other letters in their names, Kitayama and Karasawa (1997) found significant gender differences among name-letter preferences. Specifically, initial-letter preferences were stronger for men’s surname initial than for their first name initial, while females demonstrated stronger biases toward their first initial than their surname initial. The authors explain this gender difference as owing to men’s and women’s differential association between the self and their first vs. last names. Whereas Japanese men are expected to carry on their family names when married, women are expected to change their last name to that of their husband’s upon marriage. Therefore, a stronger sense of self is derived by men from their last names, while women derive their sense of self more from their first names as this is the name that will remain with them throughout their lifetime (Kitayama & Karasawa, 1997).

Gender differences were also found in Pelham et al.’s (2002) study which examined the likelihood of living in a state as a function of one’s first name. Males were 26% more likely than chance to reside in states that resembled their first name (i.e., Kenneths disproportionately populated the state of Kentucky), while females were 44% more likely than chance to live in first-name matching states. Additional analyses that focused on state immigration data and the populations of “Saint” cities also revealed stronger evidence of implicit egotism among females when these states and cities resembled their first name (Pelham et al., 2002). Because the tradition of women taking

their husbands' last names upon marriage is customary in the U.S. as well, it stands to reason that American females would also derive a stronger sense of self from their first names and exhibit—on an unconscious level—a greater attraction to and preference for places of residence which contain their first name letters.

### Explicit Self-Esteem and Self-Concept Threats

As previously discussed, name-letter preferences have been found to increase after self-esteem enhancement. Likewise, a “threat” to the self-concept can also bolster the name letter effect. Ostensibly, one way people protect their sense of self-worth in the face of self-concept threats is by automatically enhancing the value of self-associated symbols, including exhibiting preferences for name letters. For example, Jones et al. (2002) found differential name-letter and birthday-number preferences among participants with low and high explicit self-esteem after they experienced a self-concept threat. Specifically, evaluations of name letters and birthday numbers were strengthened after high explicit self-esteem participants were asked to write about a personal flaw. Born from a well-practiced need to self-enhance, preferring our name letters when confronted with them can help restore homeostasis to our temporarily injured self-esteem—much like a defense mechanism. This perspective views name-letter preferences as an unconscious form of self-regulation, with this type of self-serving bias particularly pronounced for those with high explicit self-esteem.

Brendl et al. (2005) also found differential effects for name-letter preferences after participants either wrote about something they wished to change about themselves vs. something positive about themselves. Namely, those who experienced a self-concept threat (wrote about a personal flaw) showed stronger biases toward a fictitious brand of

Japanese cracker that contained their name letters, while those in the self-affirmation condition preferred the name-letter and non-name letter brand equally (Brendl et al., 2005). Experiencing a threat to the self-concept in and of itself was a motive for participants to self-enhance when given the opportunity to evaluate a brand that shared their name letters.

### Self-Attitude Accessibility

Bosson, Swann, and Pennebaker (2000) found that explicit measures of self-esteem given prior to implicit measures—including an initial-preferences task—served to increase the correlation between the two types of self-esteem. In Krizan and Suls' (2008) meta-analysis of 10 different studies administering the NLT, the authors found a small but significant correlation between the NLT and explicit self-esteem measures. These authors also found the following order effects for the two types of measures: explicit measures of self-esteem administered prior to the NLT strengthened the correlation, while the correlation decreased but was still significant when the NLT was given first. This moderating effect of instrument order is explained as participants' self-attitudes becoming more accessible after the administration of explicit self-esteem measures and essentially priming one's own attitudes about oneself (Krizan & Suls, 2008).

### **Establishing Mediation and Moderation in**

### **Future Research on the Name Letter Effect**

In sum, when the study of name-letter preferences was still in its infancy, researchers endeavored to identify the type of mechanism underlying the name letter effect. Beginning with Hoorens and Todorova (1988), these authors failed to find support for mediation via the primacy effect when name-letter preferences were found in both the

participants' native Cyrillic alphabet and their later-learned Roman alphabet. Next, Feys (1991) discounted "the mere exposure effect" as a mediator of name-letter biases when he induced ownership of name-related and non-name related novel symbols among participants by holding the number of learning trials constant for both types of symbols, and in the end, still found strong biases toward name-related symbols. Kitayama and Karasawa (1997) also failed to find support for "the mere exposure effect" as instrumental in producing self-related preferences (birth date numbers) since participants still demonstrated biases toward the numbers 13-31, even though we are exposed to the numbers 1-12 with greater frequency. Jones et al.'s (2002) finding that people preferred their initial-letters more than other letters of the alphabet—even when their initials were rare—further discounts "the mere exposure effect" as a determinant of name-letter preferences. What this group of experiments tells us about the biases we exhibit toward our name letters is that these biases are not due to having been exposed to our name letters first in life, nor are they a product of more frequent exposure to our name letters.

Hoorens and Nuttin (1993) examined "subjective frequency" as a possible mediator of the name letter effect, testing the hypothesis that people prefer their name letters because of the false perception that they encounter them more frequently than they really do. While participants did somewhat exaggerate the frequency with which they encountered their name letters relative to non-name letters, the authors found no relationship between the name-letter preferences they observed and participants' subjective reports of their frequency of exposure to these letters. "Evaluative conditioning," while at first promising to be a more parsimonious explanation for name-letter preferences because it was devoid of both affective and cognitive processing, also

failed to be a substantiated mediator since subjects still preferred symbols associated with their own-name letters with or without their name letter serving as a visual cue (Feys, 1995). These researchers consistently found that participants preferred their name letters simply because they were just that, “theirs,” and thus “mere ownership” as a mediator of the name letter effect paved the way for a closely related determinant and what contemporary researchers now refer to as “implicit egotism.”

Implicit egotism is the underlying process responsible for the name letter effect whereby participants’ self-biases spill over into their evaluations of self-related stimuli, such as objects that share their name letters. Unlike “mere ownership,” however, implicit egotism operates outside of our awareness, whereas the mere ownership effect involves an awareness that an object is “mine” and consequently its value is overestimated (see Kahneman, Knetsch, & Thaler, 1990 for one famous empirical example of “the endowment effect”).

Early researchers’ painstaking efforts to uncover the mediating mechanism responsible for the name letter effect all point to an implicit bias toward our own self-attributes. Pioneering researchers of the name letter effect now believe that one of our many self-serving biases is at work and attempts to prove that other motivational and cognitive processes are at work would be unfruitful, if not impossible (Hoorens & Nuttin, 1993). Empirically establishing the mediating mechanism underlying self-related preferences is extraordinarily difficult due to the automatic and unconscious nature of people’s self-associations and/or self-positivity biases. Moreover, attempting to measure these self-associations and biases are liable to change name-letter preferences. Thus,



current implicit egotism scholars overwhelmingly focus their research instead on studying new moderating variables over mediators.

### **Is the NLT Simply a Measure of Implicit Self-Esteem?**

According to Pelham et al. (2002), implicit egotism reflects “an unconscious process grounded in people’s favorable self-associations” and is “an implicit judgmental consequence of people’s positive associations” (p. 106). The key assumption underlying name-letter preferences is that “people’s positive associations about themselves spill over into their evaluations of objects associated with the self” (Jones, Pelham, Mirenberg, & Hetts, 2002, p. 170). Whether the self-associations thought to drive implicit egotism and the name letter effect are positive (i.e., “favorable”) in content or are merely positively *valenced* has recently been debated. Implicit egotism and implicit self-esteem are often used interchangeably, thus measures of name-letter preferences are often used as indexes of people’s unconscious global self-attitudes, i.e., “implicit self-esteem.” Even though implicit egotism and implicit self-esteem are similar constructs, there remains some disagreement as to whether they are distinct, synonymous, or merely related.

Researchers such as Greenwald and Banaji (1995) argue that name-letter preferences are driven by individuals’ high self-esteem, with these preferences offering a glimpse into people’s global evaluations of themselves (Koole & Pelham, 2003). While these and other researchers have suggested that implicit egotism is driven by an underlying sense of high self-worth, more recent researchers, however, argue that measures like the Name Letter Test tap implicit egotism, but not necessarily implicit self-esteem (Buhrmester, Blanton, & Swann, 2011).

Buhrmester et al. (2011) are skeptical of the NLT's psychometric properties as a measure of implicit self-esteem because of its poor construct validity, low predictability of general well-being/depression, and low correlations with explicit measures of self-esteem, among other issues. The fact that name-letter preferences are overwhelmingly exhibited on the NLT—enough to compel researchers to administer it as a tool to measure self-esteem—suggests that respondents are likely relying on automatic, universal self-positivity biases instead of providing a window into their global self-worth. If name-letter preference tasks like the NLT *did* measure implicit self-esteem, one would expect significantly more variability in name-letter biases commensurate with the variability that surely exists in the population's implicit self-esteem. Surely the overwhelming majority of people are not fortunate enough to possess high implicit self-esteem. Or are they?

The jury is still out with respect to whether the NLT is a valid measure of implicit self-esteem, or whether it simply taps implicit egotism, the simple tendency to unconsciously gravitate toward objects which share our name letters—as it was “originally conceptualized” to do (Buhrmester et al., 2001). Examining the conditions under which name-letter preferences are bolstered, reduced, or even reversed is therefore of great theoretical interest and is important for future research on implicit egotism. Pelham et al. (2005) have hailed the role of implicit self-esteem in name-letter preferences as one of the next frontiers of implicit egotism, suggesting that biases might be reduced or even reversed for those who truly possess negative self-attitudes. If, however, Buhrmester et al. (2011) are correct in arguing that implicit egotism instead taps a tendency to display automatic self-positivity biases—regardless of one's level of implicit self-esteem—then implicit self-esteem should *not* impact name-letter

preferences. Thus, examining the role of implicit self-esteem in initial-letter biases was one of the aims of the present research, which are discussed in the following chapter.

## CHAPTER 2

### AIMS OF THE CURRENT STUDY

One of the aims of the current study was to offer a unique paradigm in which the name letter effect could be examined in everyday decision-making. Specifically, the researcher sought to investigate whether the otherwise subjective experience of humor would be influenced by an objective and predictable implicit process, our bias toward our name letters. A second aim was to examine two new variables that may influence the name letter effect: gender-matching and implicit self-esteem. To this end, and as described in the previous chapter, authors such as Pelham, Carvallo, and Jones (2005) have raised the question of whether implicit egotism researchers' observed name letter effects might be due to the majority of participants' good fortune of having unconscious favorable self-attitudes, and further underscore the possibility that their typical implicit egotism findings "would be reversed among people who possess truly negative self-associations, i.e., for those with low levels of implicit self-esteem" (Pelham et al., 2005, p. 109). Thus, research which seeks to pursue this possibility that typical implicit egotism findings could be reversed among individuals with unfavorable implicit self-esteem is extremely theoretically worthwhile and will make important contributions to social psychology and research on the name letter effect. However, measuring an implicit construct such as self-esteem is often tenuous and requires indirect assessment. Fortunately, a valid and efficient measure has recently been developed to measure implicit self-attitudes.

Authors Gebauer, Riketta, Broemer and Maio (2008) examined participants' liking for their entire names instead of individual name letters to devise a single-item measure of implicit self-esteem. For this measure, participants are simply asked to indicate how much they like their full name using a scale with endpoints of "not at all" to "very much." Comparisons with two widely used measures of implicit self-biases, the Self-Esteem IAT (Greenwald & Farnham, 2000) and the Name Letter Test (Nuttin, 1985; 1987), found Gebauer et al.'s (2008) single-item measure to be correlated with both of these instruments. Name-liking was also highly correlated with two other explicit measures of global self-attitudes, including results obtained from quick responses on explicit self-esteem indexes and under conditions of high cognitive load.

Gebauer et al. (2008) argue that name-liking is superior to the NLT as a measure of implicit self-esteem because a) more meaning is attached to groups of letters than individual letters, b) whole words instead of letters are stored in our memory, c) letter order determines the meaning of an object and how it is evaluated, d) there is no other attribute or object more closely related to the self than one's own name, and e) according to Gestalt psychology, the whole is generally considered to be greater than the sum of its parts. For these reasons, the authors believe that liking of one's own name is a more efficient and valid measure of implicit self-esteem than the Name Letter Test. Moreover, this single-item measure addresses some of Buhrmester, Blanton, and Swann's (2011) concerns with the NLT as a valid measure of implicit self-esteem. Namely, unlike the NLT, Gebauer et al.'s (2008) measure was correlated with general well-being and explicit self-esteem, demonstrated high test-retest reliability, and was not dependent on the order of administration of implicit-explicit measures. In addition, name-liking is more likely

than preferences for individual name letters to tap autobiographical information, addressing yet another one of Buhrmester et al.'s (2011) doubts regarding the validity of the NLT as a measure of implicit self-esteem.

Previously, Hoorens & Nuttin (1993) found differences in name-letter preferences were associated with participants' beliefs that their name a) would suit someone of the opposite gender, b) would suit a member of a professional group to which they would like to belong, and c) carried a strong likelihood of being chosen as a name for themselves if they would have had the opportunity to so choose. Their study provides conceptual justification for Gebauer et al.'s (2008) name-liking measure, but stopped short of establishing it as a valid measure of implicit self-esteem. Nevertheless, their study gives legs to the hypothesis and provides preliminary support that implicit-self attitudes—as measured by name-liking—might influence name-letter biases. Currently, Gebauer et al.'s (2008) measure and Hoorens and Nuttin's (1993) findings are just the tip of the implicit egotism iceberg. The present study explored the relationship between implicit self-esteem and name-letter preferences more deeply by examining—as Pelham et al. (2005) suggested—whether typical implicit egotism findings would be reversed among those who truly possess negative self-associations, i.e., poor implicit self-esteem. Differences in name-letter preferences were examined among participants with high and low implicit self-esteem, as measured by Gebauer et al.'s (2008) name-liking instrument.

Thus, the proposed study sought to a) fill this important gap in implicit egotism research by examining the role of implicit self-attitudes in biases toward name letters, b) foster a more comprehensive understanding of the self-concept and our resulting self-biases, and c) contribute theoretically to social psychology and research on implicit

egotism by challenging the notion that the name letter effects observed in previous studies are the result of an overwhelming majority of participants' good fortune of possessing favorable self-attitudes. By measuring implicit self-esteem, the proposed study also aimed to determine whether a) positive self-attitudes are the norm for the general population, b) one's level of implicit self-esteem impacts name-letter preferences, c) if biases toward name letters occur in general regardless of one's level of self-esteem, and d) if name-letter preferences are synonymous with implicit self-esteem.

Previous research has also demonstrated that name-letter biases can vary as a function of gender, with women demonstrating stronger name letter effects than men for their first names (Pelham et al., 2002; Kitayama & Karasawa, 1997). Therefore, another aim of the current study was to examine whether women participants would exhibit stronger preferences for captions submitted by writers with whom they shared a first name initial and if they would display stronger biases toward their initial letter on the NLT.

Another aim of the present study was to examine the effect of self-concept threats on name-letter preferences. Based on Jones et al.'s (2002) and Brendl et al.'s (2005) research, it was expected that participants who experienced a minor threat to the self-concept would exhibit modestly stronger biases compared to those who received a small boost to their self-concept. A self-concept threat was expected to serve as a motive for participants to self-enhance when given the opportunity to evaluate captions submitted by writers who shared their first name initial. Evidence of a name letter effect would support implicit egotism as one of our many self-serving biases. Explicit self-esteem, as measured by the Rosenberg Self-Esteem Scale (Rosenberg, 1989), would likely moderate the effect

of a self-concept threat. Participants high in explicit self-esteem were expected to exhibit stronger name-letter preferences than participants with low explicit self-esteem after experiencing the same threat. Born from the well-practiced need to self-enhance, preferences for captions submitted by writers with whom participants shared an initial letter would provide an opportunity for ordinarily high explicit self-esteem individuals to regulate their recently injured self-esteem.

Krizan and Suls' (2008) meta-analysis found the accessibility of participants' self-attitudes influenced name-letter preferences when explicit self-esteem measures were administered immediately prior to versus after the Name Letter Test. Specifically, explicit self-esteem measures were found to prime positive self-evaluations before the NLT because participants' self-attitudes became more accessible, resulting in stronger biases. The order of the current study's explicit self-esteem measure would thus be manipulated in the same fashion to test for differences in name-letter preferences between participants with high and low self-attitude accessibility.

A final aim of the proposed study—and representing a “new twist” on implicit egotism research—was to examine shared gender as a possible influence on name-letter biases. Previous research has found gender-matching effects in the areas of supervisor-supervisee relationships (Worthington, Jr. & Stern, 1985) and therapist-patient treatment outcomes (Zlotnick, Elkin, & Shea, 1998; Sterling, Gottheil, Weinstein & Serota, 1998), but gender-matching effects have never before been examined among biases demonstrated toward name letters. In a unique task designed to simultaneously measure both initial-letter biases *and* gender-biases, participants would rate cartoon captions ostensibly submitted by men and women writers who shared and did not share their first



initial. Analysis of participants' and caption writers' gender would provide a novel test of gender-matching as a potential influence on name-letter preferences and further advance the study of implicit egotism with new data.

In sum, the present study sought to add to the growing body of implicit egotism research by examining several influences (gender differences, self-attitude accessibility, and self-concept threats) on name-letter biases using a novel paradigm. In addition, two new variables were examined as potential influences on name-letter preferences: implicit self-esteem and shared gender. A special emphasis was placed on teasing out the role implicit self-attitudes play in initial-letter biases to determine if the name letter effect reflects a universal and automatic positive self-bias *or* if it reflects one's level of implicit self-esteem, with typical implicit egotism findings possibly being reversed for participants with negative implicit self-attitudes—an important empirical question for future research raised by Pelham et al. (2005).

Hypotheses reflecting the current study's aims are summarized in Table 1. Methods for experimentally manipulating variables will be described in more detail in the following chapter, along with descriptions of the study's measures and their administration.

Table 1. Hypotheses

	Reference	Analysis	Hypothesis
a. Primary analyses			
1) The Name Letter Test	Nuttin (1985, 1987)	<i>t</i> test	Initial-letter biases will be significantly different from zero.
2) Name-letter preferences in humor judgments	–	<i>t</i> test	Biases for captions written by same-initial writers will be significantly different from zero.
b. Secondary analyses			
3) Self-attitude accessibility	Krizan & Suls (2008)	<i>t</i> test	Initial-letter biases will be stronger among participants with high self-attitude accessibility.
4) Self-concept threat	Jones et al. (2002); Brendl et al. (2005)	ANOVA	Initial-letter biases will be stronger among participants who experience a threat to the self-concept.
5) Self-concept threat $\times$ explicit self-esteem interaction	Jones et al. (2002)	multiple regression	Initial-letter biases will be especially strong among participants with high explicit self-esteem after a self-concept threat.
6) Sex differences	Kityama & Karasawa (1997); Pelham et al. (2002)	<i>t</i> test	Initial-letter biases will be stronger among women.
7) Sex-matching	–	<i>t</i> test	Initial-letter biases will be stronger when caption writers' gender matches that of a participant.
8) Implicit self-esteem	–	regression	Level of implicit self-esteem will predict initial-letter biases.

## CHAPTER 3

### METHODS

#### **Participant Recruitment**

Two separate identical studies—one recruiting men and one recruiting women—were posted simultaneously in Experimetrix, Loyola University Chicago’s web-based psychology experiment scheduling portal. Both studies were titled “Judgments of Humor and Language in the 21<sup>st</sup> Century” and were designed to be completed in one hour or less as an online experiment. Respondents received one experiment participation credit hour for completing the experiment. Recruitment of participants began at the beginning of the fifth week of the Fall 2011 semester and lasted 11 weeks.

#### **Sample**

Participants were 503 men and women enrolled in Psychology 101 at Loyola University Chicago. The average weekly number of respondents who signed up for and completed the experiment was 45, range = 14 (week 11) to 74 (week 2). An average of 15 men completed the experiment each week, range = 10 (week 3) to 24 (week 10), while an average of 29 women completed the experiment each week, range = 3 (week 11) to 57 (week 6). For weekly participant recruitment frequencies for men and women, please see Table 2.

Table 2. Weekly participant recruitment frequencies

Week #	Men	Women	Total
1	14	40	54
2	20	54	74
3	10	38	48
4	20	38	58
5	13	11	24
6	11	57	68
7	12	29	41
8	19	25	44
9	14	10	24
10	24	19	43
11	11	3	14
Overall	168	324	492

Based on the appearance of traditionally feminine first and/or middle names in Experimetrix's participant sign-up list, at least 6 women signed up incorrectly for the men's version of the experiment, while no men incorrectly signed up for the women's version. These estimates, however, could vary when taking into account gender-neutral, ethnic, or other types of names. Because respondents' names could not be associated with their surveys, these six women's data were unable to be removed from the men's set of responses, however, they accounted for less than 4% of what will henceforth be considered the (all) men participants.

Nine participants' data were removed because they left their surveys blank either due to choice, a technological error, or because they closed the internet window during the experiment. Two respondents were excluded because they disagreed to participate in the experiment after signing up. The resulting  $N$  was 492 participants (324 women and 168 men).

### **Completion Times**

The amount of time participants took to complete the experiment was recorded online. On average, respondents spent 16.46 minutes completing the experiment (range = 2–624 minutes,  $SD = 35.76$ ). Fourteen cases were excluded from this average because they did not finish the experiment, thus a completion time was not recorded. Men ( $N = 166$ ) on average spent 14.32 minutes completing the experiment (range = 2–198 minutes,  $SD = 17.06$ ), while women ( $N = 312$ ) on average spent 17.61 minutes (range = 4–624 minutes,  $SD = 42.46$ ). Even though women on average spent more than 3 minutes longer than men completing the experiment, this difference was not statistically significant,  $t(476) = .96, p = .34, d = .10$ .

Immediately, the minimum and maximum completion times suggest, respectively, that some participants might not have complied with the instructions for parts or all of the experiment, while others might not have relied on their gut feelings when instructed to do so. It is also possible that respondents with longer completion times might have suspended their participation and returned to the experiment at a later time. As with all completion times, it is impossible to say which parts of the survey could have been compromised because completion times for each page of the experiment were not recorded.

Kolmogorov-Smirnov and Shapiro-Wilk tests of normality for the distribution of completion times were both significant,  $p < .001$ . Exclusion of outliers based on three or more standard deviations was not instructive. Because the experiment was designed to be completed well within a one-hour time frame, cases where participants took longer than one hour to finish the experiment ( $N = 10$ , 2.1%) were considered to be potential outliers on the completion time variable. Previous researchers, however, have refrained from using long completion times as an exclusionary criterion, arguing that their survey program did not record completion times for individual pages so it would be impossible to determine on which part(s) respondents lingered, if they simply opened the survey but began it at a another time, or if they began the survey and returned to it at a later time (Nosen & Woody, 2008). It is worth mention that no participants contacted the experimenter to request additional participation credit hours or to complain that the experiment took longer than the projected (maximum) time of one hour.

Short completion times raise similar concerns as to whether respondents conscientiously completed the experiment. Twenty-nine participants (6.1%) had completion times of less than 7 minutes, which fell under the 10<sup>th</sup> percentile in the distribution of respondent completion times. It is unclear as to whether meaningful participation could occur with experiment completion times under 7 minutes. However, because name-letter preferences reflect an automatic and highly efficient unconscious process, shorter experiment completion times are not particularly troubling.

As a precaution, the researcher took a conservative approach and analyzed experiment data both including and excluding data from participants with long and short completion times. Fortunately, results were extremely similar and often identical,

regardless of the amount of time respondents spent completing the experiment. Thus, results reported reflect data from the entire set of participants ( $N = 492$ ).

### **Procedure**

At the time of sign-up in Experimentrix, a hyperlink to the web study was displayed to participants so they could click on and begin the experiment. The study could be completed online using an internet connection from respondents' home computers, or if desired, from a campus computer. Experimentrix parameters were set so that respondents could only sign up for the experiment once and participation in either the men's or women's version of the web study precluded eligibility for participation in the women's or men's version, respectively.

The format of the experiment was developed in Opinio, a widely-used and Loyola-endorsed web research software tool. Loyola's Office of Research Services (ORS) and Information Technology Services (ITS) departments have worked in tandem with Opinio to create specifications that secure participants' anonymity and the data they enter according to the highest security standards possible. Data transmitted from respondents' computers are encrypted and stored on the Opinio server, which is protected inside Loyola's perimeter firewall. Participants' IP addresses were hidden from the researcher.

The ITS department was able to build a randomization process into the survey link provided to participants upon sign-up for the experiment. Clicking on the same web URL randomly directed men respondents to one of 24 different surveys, which featured slight differences in the order of measures and contained one of three self-concept threat conditions. A different web URL was provided to women participants who signed up for

the experiment, which also randomly directed them to one of 24 different surveys featuring the same slight differences in the order of measures and contained one of three of the men's same threat conditions. Again, men and women participants' experiments were identical in content, orders of measures, and threat conditions, but were listed separately as a means to merely eliminate the need to ask respondents to indicate their gender within the survey.

Participants completed the online experiment anonymously, with the exception of the identifying information of gender (determined by the version of the study they signed up for) and their first initial (which was queried in a survey question). No other identifying information was requested or collected from respondents.

### **Informed Consent**

On the first webpage of the survey, participants were given information about the nature, purpose, and type of questions that were included in the experiment. The consent process adhered to guidelines set forth by Loyola University Chicago's Institutional Review Board for web-based research and used Loyola's General Consent Form Template. Elements included a confidentiality disclaimer, information regarding participants' anonymity, contact information for the primary investigator and faculty sponsor, and respondents' inability to retrieve or discard responses from the database once submitted. Participation throughout the duration of the experiment was entirely voluntary and respondents were free to refrain from answering any question(s) for any reason or discontinue their participation in the study without penalty. Results of the study were made available to those who requested it.



## **Measures**

### **Humor Consumption Survey**

All respondents who consented to participate in the experiment first completed a two-item survey which measured their exposure to cartoons and other humorous media. Participants were asked to indicate how often they read cartoons in newspapers or magazines and how often they watched humorous television shows or movies using the provided scale of 1) every day, 2) once a week, 3) once a month, 4) a few times a year, and 5) almost never. The purpose of this survey was to help disguise the study's hypotheses and to help frame the experiment as a study of humor and language.

### **Self-Esteem Threat Manipulation**

Following an experimental manipulation used by Jones, Pelham, Carvallo, and Mirenberg (2002), participants were randomly assigned to surveys containing one of three conditions: a self-concept threat, affirmation, or control writing task. In the self-concept threat condition, respondents were asked to write at least three sentences about an aspect of themselves that they have found difficult to change, but would like to be different. Participants in the self-concept affirmation condition were asked to write at least three sentences about an important area of their lives in which they have always felt good about themselves and represented a positive, important, and stable aspect of who they are. Finally, respondents in the control condition were simply asked to write at least three sentences about the last movie they saw. The purpose of this component of the experiment was to deliver a mild threat to participants' self-concept and compare their name-letter preferences with the other two experimental conditions.

### Post-Manipulation Mood

Subsequent to the self-concept threat, affirmation, or control manipulation, all participants were asked to complete a brief mood questionnaire to ensure that observed differences in initial-letter biases could be attributable to the threat manipulation, instead of due to a manipulation of mood. The questionnaire was identical to the one used by Jones et al. (2002) for this same purpose and consisted of two items which asked respondents to indicate their current mood and how they were feeling at the current moment using the scale of 1) extremely sad to 7) extremely happy.

### Rosenberg Self-Esteem Scale

Next, one-half of participants were randomly assigned to complete the Rosenberg Self-Esteem Scale (Rosenberg, 1989), a brief widely-used measure of self-reported (“explicit”) self-esteem. It consists of 10 items which assess general self-value and self-worth. The remaining half of participants completed the measure at the end of the study as a means to vary self-attitude accessibility and to examine its impact on initial-letter biases. Please see Appendix A for the Rosenberg Self-Esteem Scale (RSES).

### Initial-Letter Biases

Participants were next shown two randomly-selected cartoons licensed from *The New Yorker* magazine’s “cartoon caption contest” collection, along with 20 different possible captions for each cartoon. More than one million people have entered their caption submissions for the over 300 contests the magazine has held since the competition was first introduced in 2005. All of the several thousand captions submitted for each weekly contest are publicly viewable online, however the identity of caption writers are kept anonymous. Names of caption writers are only included with their

submissions when they are selected as one of three finalists by the cartoon editorial staff for each week's cartoon. The public then votes on the top three entries, and based on these votes, a winner is chosen. As a "mock" version of the magazine's cartoon caption contest, participants were asked to play the role of a "judge" and evaluate 20 captions provided for each of the two cartoons (40 captions in total).

Captions for the cartoons were selected from the thousands of submissions posted online according to the following criteria: a) they referenced the juxtaposition of the two opposing elements featured in the cartoon, b) they were grammatically and semantically correct, and c) they were short in length, i.e., approximately 10–12 words. While the humor and interpretation of each caption would be largely subjective, adhering to the above criteria ensured that a) the captions resembled typical cartoon captions, b) they could be read quickly, and c) they were brief enough to allow participants to compare multiple captions when assigning ratings, if desired. Respondents were asked to individually and quickly rate each set of 20 captions using the scale of 1) not at all humorous to 7) extremely humorous, while using their "gut feelings."

Importantly, fictitious caption writers' names were referred to as "students who submitted each caption," with caption writers' names printed in **bold** beside each caption. Emboldening the names served to highlight the writers' names, while placing the names in a prominent position, i.e., "flush left" and preceding each caption, helped to further emphasize and draw attention to writers' names while participants were evaluating captions and making their ratings. These methods were used in a concerted effort to help ensure that participants attended to caption writers' names. However, because previous researchers have found name letter effects even when using subliminal (13.7 millisecond)

name-letter targets (Jones et al., 2004), whether participants attended to name-letter stimuli in the current study was not an immediate concern.

Each caption was randomly assigned a fictitious “submitted by” name that began with a different letter of the alphabet, with the exception of the letters O, Q, U, X, Y, and Z. These six letters were omitted because a) they are the rarest first name initials according to recent Census data ([census.gov/genealogy/names/](https://www.census.gov/genealogy/names/)), b) asking participants to rate 26 captions might have revealed the hypotheses of the study, c) including all 26 letters would have increased the length of the task and experiment, while d) requiring participants to rate a large, uncommon, and potentially unwieldy number of items. The caption-rating task was a unique method designed to simultaneously allow measurement of name-letter preferences while also manipulating the gender of caption writers. In this way, the current investigation became the first implicit egotism study to examine both types of biases in tandem.

Arrangement of captions could not be completely randomized using the survey software technology, so instead participants were randomly assigned to one of four cartoon caption order presentations: forward order for Cartoons 1 and 2; forward order for Cartoon 1 and backward order for Cartoon 2; backward order for Cartoon 1 and forward order for Cartoon 2; and backward order for Cartoons 1 and 2. Within these caption arrangements, the first letter of caption writers’ names (one name each for the letters A–Z, with the exceptions of O, Q, U, X, Y, and Z) were randomly positioned for Cartoon 1 and Cartoon 2. Gender of fictitious caption writers (10 men and 10 women), were randomly assigned for Cartoon 1, with the opposite gender assigned to that initial-

letter for Cartoon 2. Please refer to Appendix B for the two cartoons used in the study, with sets of captions presented in “forward” arrangement order for both cartoons.

### Hypothesis Awareness Questionnaire

After the caption-rating exercise, respondents completed an “influences on humor judgments” questionnaire as a test of hypothesis awareness. The questionnaire probed whether participants were aware of any role the similarity between their names and the caption writers’ might have played in their caption preferences and ratings. Respondents were asked to indicate anything that might have influenced their humor ratings in the preceding task by choosing one or more of the following variables: the cleverness of the caption’s language; the caption writer’s interpretation of the drawing; their individual sense of humor; the number of cartoons they read per week; the number of humorous television shows or films they watch per week; and a variable marked “other,” along with text space to type in a potential influence that was not already listed.

### Name Letter Test

Next, respondents completed the traditional Name Letter Test (NLT), which was described as a survey of “aesthetic judgments of lexical stimuli.” Because letter order could not be randomized for each participant, letters were displayed in one random fixed order for all respondents. Participants were asked to estimate how beautiful they found each of the 26 letters of the alphabet using their “gut feelings.” It was explained that even though it might seem unusual to evaluate letters in terms of their beauty, previous research studying these types of evaluations has fostered “a better understanding of language and human emotions.” Please see Appendix C for a copy of the NLT.

### Implicit Self-Esteem Measure

After the NLT, participants completed Gebauer, Riketta, Broemer, and Maio's (2008) single-item measure of implicit self-esteem. This measure asks respondents to simply indicate how much they like their full name in total (including their first and last name together) using the scale 1) not at all to 7) very much. It was purposely administered after the humor judgments exercise and NLT so as not to raise participants' suspicion that their name might play a role in or otherwise influence their preferences for cartoon captions and evaluations of alphabet letters.

### Self-Attitude Accessibility

Finally, to conclude the experiment, the remaining half of participants randomly assigned to the "low self-attitude accessibility" condition completed the RSES at this time. The other half of respondents who were randomly assigned to the "high self-attitude accessibility" condition had already completed this portion of the experiment prior to the humor judgments exercise. For participants in the "high self-attitude accessibility" condition, the experiment concluded after they completed the single-item implicit self-esteem measure described above. The experiment concluded for the "low self-attitude accessibility" participants upon completion of the RSES.

### Debriefing

All respondents were debriefed at the end of the experiment and were given additional information about the study's purpose, which was described as an investigation of self-related biases in humor judgments and aesthetic evaluations of letters of the alphabet. The debriefing statement explained that although some judgments and evaluations are often largely subjective, research on implicit egotism has shown that

people sometimes make both everyday and important decisions based on their preferences for self-related attributes, including one's name letters.

Respondents were made aware of the different experimental conditions to which they were randomly assigned. It was explained that these conditions consisted of measures developed to assess self-esteem and contained tasks designed to briefly deliver a mild threat or affirmation to their self-concept, or a control condition. Hypotheses underlying these manipulations were offered and participants were urged to refrain from discussing the experiment with other students.

Participants interested in learning more about implicit egotism were provided with a brief introductory-level journal article on the topic, which was accessible from Loyola University Chicago's e-journal database. They were also given email addresses for the Primary Investigator and her Faculty Sponsor, should they have any complaints, concerns, or questions about the study they participated in.

## CHAPTER 4

### PRELIMINARY ANALYSES

#### **Humor Consumption**

Respondents were first asked about their frequency of humor consumption in order to frame the study as one assessing judgments of humor and to help disguise the hypotheses of the experiment. All participants ( $N = 492$ ) indicated how often they read cartoons in newspapers or magazines by choosing one of the following responses: every day ( $N = 11$ ), once a week ( $N = 86$ ), once a month ( $N = 105$ ), a few times a year ( $N = 136$ ), or almost never ( $N = 154$ ).

Next, participants indicated how often they watched humorous television shows or movies by choosing from the same set of responses: every day ( $N = 165$ ), once a week ( $N = 264$ ), once a month ( $N = 45$ ), a few times a year ( $N = 13$ ), or almost never ( $N = 5$ ). Although respondents' consumption of humor was not of particular interest to this study and the purpose of these two questions was merely to "frame" the study as one examining judgments of humor, reported vs. expected frequencies of reading cartoons and watching humorous television shows or movies were significantly different,  $\chi^2 = 26.71$ ,  $df = 16$ ,  $p < .05$ , which is considered a small- to medium-sized effect,  $w^2 = .05$  (Cohen, 1988).

#### **Hypothesis Awareness**

The researcher probed for hypothesis awareness after participants completed the humor judgments exercise. Respondents were asked to choose one or more variables that might have influenced their caption ratings from the following list: the cleverness of the



captions' language, the caption writers' interpretations of the drawing, the participant's individual sense of humor, the number of cartoons he or she typically reads per week, and the number of humorous television shows or films watched per week, along with an "other" variable. If "other" was selected, respondents were asked to describe any additional variables that might have influenced their ratings that were not already listed. A blank text field was provided for respondents to list these additional variables.

Participants' individual sense of humor was the most frequently chosen variable as a possible influence on their previous cartoon caption ratings ( $N = 424$ ), followed by the cleverness of the captions' language ( $N = 395$ ), the caption writers' interpretations of the drawing ( $N = 297$ ), and the number of humorous television shows or films respondents watched per week ( $N = 55$ ). The number of cartoons participants read per week ( $N = 22$ ) and "other" ( $N = 22$ ) were the least frequently chosen influences. Analysis of "other" variables provided by respondents revealed neither an awareness of shared initials with caption writers as a possible influence on their caption ratings, nor any of the study's other hypotheses. Responses when indicated ( $N = 21$ ) included variables such as "puns," the "caption writer's originality," "irony," and the "ability to use common language to create a joke." Based on respondents' lack of awareness of the potential role shared initials with caption writers might have played in their judgments of humor, the researcher can conclude that a) the hypotheses of the experiment were not known or discovered by participants and b) any preferences that emerged toward initial-letter stimuli were implicit because they occurred outside of respondents' awareness.

## CHAPTER 5

### PRIMARY ANALYSES OF INITIAL-LETTER BIASES

#### **The Name Letter Test**

Initial-letter biases on the Name Letter Test (NLT) were computed using an ipsatized scoring algorithm (LeBel & Gawronski, 2009; Baccus, Baldwin, & Packer, 2004), which is a conservative approach to computing name-letter preferences because it double-corrects name-letter ratings at both the individual level and at the group level. First, the mean rating a participant assigned to all non-initial letters is subtracted from each of his or her letter ratings, including his or her initial letter. Next, normative letter baseline ratings for each letter are calculated by averaging the ipsatized letter ratings assigned by participants who *did not* have the letter as an initial. The final step involves subtracting the normative ipsatized baseline rating for a participant's initial letter from their ipsatized rating of that letter to obtain his or her "initial-letter bias." SPSS syntax for computing initial-letter biases with the I-algorithm (along with other algorithms) was obtained from Dr. LeBel's University of Western Ontario research website (2011). Computations were randomly and manually confirmed for accuracy, and were also confirmed exhaustively with Microsoft Excel.

For example, 420 participants rated the letter "A" from 1) not at all beautiful to (7) extremely beautiful. Of these, 55 participants' first initial was "A," so normative baselines were derived from the average of the 365 remaining non-initial matching participants' ipsatized ratings for the letter "A." Then, this normative ipsatized baseline

rating for the letter “A” was subtracted from each participant’s ipsatized rating for the letter “A” whose name began with that letter, with the result yielding their “initial-letter bias.” Please see Table 3 for a list of normative ipsatized baseline ratings, mean ipsatized ratings by initial-matching participants, and mean initial-letter biases for all 26 letters of the alphabet.

Out of the 492 total respondents, 420 (150 men and 270 women) provided their first initial for their first given name and had non-missing and non-redundant letter ratings on the Name Letter Test (NLT). Using the I-algorithm, results showed that participants displayed a very strong bias toward their own initial letters. The average difference between participants’ ipsatized ratings for their initial-letter and the normative ipsatized baseline value for that letter was +1.38 (range = –3.43–6.05,  $SD = 1.63$ ), which was statistically significant,  $t(419) = 17.28$ ,  $p < .001$ . This effect is considered large,  $d = .84$ , and is consistent with the medium to large effects found in recent studies on the name letter effect (LeBel & Gawronski, 2009). In sum, when it came to rating one’s first initial letter on the NLT, participants increased their ratings an average of 1.38 points on the 1) not at all beautiful to 7) extremely beautiful scale, relative to their average rating for non-initials and the corresponding normative ipsatized baseline rating for their initial.

Inspection of individual NLT scores that were three standard deviations below the mean (scores less than or equal to –3.52) and three standard deviations above the mean (scores greater than or equal to +6.27) produced no outliers. Examination of Tukey box plots, however, identified two potential outliers—one with an NLT score of –3.43 and one with an NLT score of +6.05. Removing them produced an identical average initial-

Table 3. Average NLT initial-letter biases for individual letters of the alphabet

Letter	Non-initial matching		Initial-matching		
	<i>N</i>	Normative baseline rating	<i>N</i>	$\bar{X}$ ipsatized rating	$\bar{X}$ initial-letter bias <sup>1</sup>
A	365	.88	55	1.89	1.01 ***
B	405	.34	15	2.87	2.52 ***
C	392	-.17	28	1.59	1.77 ***
D	406	-.89	14	1.35	2.23 ***
E	394	.08	26	1.71	1.62 ***
F	409	-.53	11	1.88	2.42 **
G	410	-.01	10	1.80	1.81 **
H	411	-.37	9	1.84	2.21 **
I	414	-.86	6	.19	1.06
J	366	.15	54	1.28	1.14 ***
K	394	.11	26	1.03	.92 **
L	406	.15	14	1.24	1.09
M	372	.52	48	1.65	1.13 ***
N	405	-.20	15	1.65	1.84 **
O	417	.15	3	.89	.74
P	409	-.49	11	1.62	2.11 **
Q	420	.46	0	—	—
R	410	.33	10	1.64	1.31 *
S	381	1.40	39	2.06	.66 **
T	409	-.41	11	.99	1.39 *
U	420	-.76	0	—	—
V	413	-.14	7	1.32	1.46 *
W	419	.18	1	1.96	1.78 <sup>2</sup>
X	419	.16	1	1.68	1.52 <sup>2</sup>
Y	417	-.10	3	2.73	2.83
Z	417	.25	3	2.43	2.18
Overall			420		1.38 ***

<sup>1</sup> initial-letter bias averages that deviate by  $\leq .02$  are due to rounding error.<sup>2</sup> significance could not be computed because the sum of caseweights = 1.\* indicates bias is significantly different from zero ( $p < .05$ , two-tailed).\*\* indicates bias is significantly different from zero ( $p < .01$ , two-tailed).\*\*\* indicates bias is significantly different from zero ( $p < .001$ , two-tailed).

letter bias of +1.38 (range =  $-2.83$ – $5.64$ ,  $SD = 1.60$ ), which was also again significant  $t(417) = 17.56$ ,  $p < .001$ ,  $d = .86$ . Because excluding these two respondents' data produced identical results, all participants were included in subsequent NLT analyses.

### **Initial-Letter Biases in Humor Judgments**

#### **Cartoon 1**

Biases toward captions written by writers who shared participants' first initials were computed in the same manner as initial-letter biases on the NLT, i.e., using the I-algorithm (LeBel & Gawronski, 2009; Baccus, Baldwin, & Packer, 2004). Results are based on data from respondents with complete and non-redundant ratings for the entire set of twenty Cartoon 1 captions, and those whose initial letter matched one of the twenty caption writers' initials ( $N = 428$ ).

The average initial-letter bias exhibited toward captions was  $+1.18$  (range =  $-3.46$ – $5.09$ ,  $SD = 1.59$ ). When it came to judging captions written by writers with whom participants shared a first initial, respondents on average increased their ratings by almost two-tenths of a point on the 1) not at all humorous to 7) extremely humorous scale, relative to ratings of captions written by non-initial matching writers and as compared to normative ipsatized baseline caption ratings. While this observed initial-letter bias was significant,  $t(427) = 2.35$ ,  $p = .02$  (two-tailed), the effect was small,  $d = .11$ . For a list of normative ipsatized baselines, mean ipsatized ratings by initial-matching participants, and mean initial-letter biases for Cartoon 1 captions, please refer to Table 4.

#### **Cartoon 2**

Biases toward captions written by writers who shared respondents' first initials were again computed using the I-algorithm (LeBel & Gawronski, 2009; Baccus, Baldwin,

& Packer, 2004). Results are based on participants with complete and non-redundant ratings for the entire set of twenty Cartoon 2 captions and those whose initial letter matched one of the twenty caption writers' initials ( $N = 429$ ).

Table 4. Average Cartoon 1 initial-letter biases for individual caption writers

Writer	Non-initial matching		Initial-matching		
	$N$	Normative baseline rating	$N$	$\bar{X}$ ipsatized rating	$\bar{X}$ initial-letter bias <sup>1</sup>
Kurt	401	-.19	27	-.45	-.27
Vanessa	422	-.69	6	-.45	.24
Amanda	371	-.43	57	-.23	.20
Ian	421	.14	7	.53	.38
Jessica	374	.59	54	.98	.39
Frank	416	-.73	12	-.36	.37
Thomas	416	-.15	12	-.45	-.30
Greg	418	1.17	10	1.42	.25
Helen	418	.10	10	-.23	-.33
Dennis	408	-.12	20	-.20	-.08
Pauline	416	-.74	12	-.61	.13
Bethany	412	-.39	16	-.27	.13
Regina	418	.41	10	.54	.13
Nicole	414	.10	14	-.35	-.45
Lauren	413	-.60	15	.17	.77
Samantha	393	.39	35	.52	.13
Eric	400	.93	28	.49	-.45
Wes	427	-.66	1	.84	1.51 <sup>2</sup>
Christopher	397	.60	31	1.25	.65
Matt	377	.40	51	.89	.50 *
Overall			428		.18 *

<sup>1</sup> initial-letter bias averages that deviate by  $\leq .01$  are due to rounding error.

<sup>2</sup> significance could not be computed because the sum of caseweights = 1.

\* indicates bias is significantly different from zero ( $p < .05$ , two-tailed).

The average bias exhibited toward Cartoon 2 captions was  $+.10$  (range =  $-3.81$ – $4.47$ ,  $SD = 1.48$ ). When it came to rating captions written by writers with whom participants shared a first initial, respondents on average increased their ratings by one-tenth of a point on the 1) not at all humorous to 7) extremely humorous scale, relative to their ratings of captions written by non-initial matching writers and as compared to normative ipsatized baseline ratings. While in the predicted direction, this bias did not reach statistical significance,  $t(428) = 1.35$ ,  $p = .18$  (two-tailed),  $d = .07$ . For a list of normative ipsatized baselines, mean ipsatized ratings by initial-matching participants, and mean initial-letter biases for Cartoon 2 captions, please refer to Table 5.

#### Both Cartoons

Participants demonstrated a statistically significant bias toward Cartoon 1 captions submitted by writers with whom they shared an initial letter. A similar effect was observed for Cartoon 2 captions, however this bias failed to reach statistical significance. After considering caption ratings for Cartoon 1 and Cartoon 2 separately, initial-letter biases were averaged together across both sets of captions. Data from respondents who shared an initial with caption writers and who had complete and non-redundant caption ratings for both Cartoon 1 *and* Cartoon 2 were included in analyses ( $N = 394$ ). Participants' average initial-letter biases were computed by the I-algorithm (LeBel & Gawronski, 2009; Baccus, Baldwin, & Packer, 2004).

The average bias exhibited across both sets of captions was  $+.15$  (range =  $-2.46$ – $3.12$ ,  $SD = 1.08$ ). When it came to judging captions written by writers with whom participants shared a first initial, respondents on average increased their ratings by  $.15$  points on the 1) not at all humorous to 7) extremely humorous scale, relative to their

Table 5. Average Cartoon 2 initial-letter biases for individual caption writers

Writer	Non-initial matching		Initial-matching		
	<i>N</i>	Normative baseline rating	<i>N</i>	$\bar{X}$ ipsatized rating	$\bar{X}$ initial-letter bias <sup>1</sup>
Scott	391	-.01	38	.29	.30
Gillian	420	.56	9	-.27	-.83
Meghan	375	.04	54	.13	.09
Vince	422	-.11	7	1.08	1.19
Brian	415	-.11	14	.63	.74 *
Isabel	423	.11	6	.32	.20
Andrew	371	-.61	58	-.62	-.01
Trisha	417	.03	12	.28	.25
Danielle	410	-.17	19	-.30	-.14
Peter	418	.71	11	1.38	.67
Felicia	418	-.10	11	-.33	-.24
Emma	398	.30	31	.72	.42
Logan	415	.78	14	1.08	.30
Kristen	403	-.58	26	-.30	.29
Ryan	418	-.78	11	-.96	-.18
Chelsea	400	.32	29	.17	-.14
Jacob	373	.34	56	.14	-.20
Hank	421	-.50	8	-.47	.03
Whitney	428	.70	1	2.74	2.04
Noah	415	-.95	14	-1.05	-1.10
Overall			429		.10

<sup>1</sup> initial-letter bias averages that deviate by  $\leq .01$  are due to rounding error.<sup>2</sup> significance could not be computed because the sum of caseweights = 1.\* indicates bias is significantly different from zero ( $p < .01$ , two-tailed).



ratings of captions written by non-initial matching writers and as compared to normative ipsatized baseline ratings. When biases were averaged together for Cartoon 1 and Cartoon 2, biases exhibited toward captions written by same-initial writers were statistically significant,  $t(393) = 2.67, p = .01$  (two-tailed). This effect, however, is considered small,  $d = .14$ .

### **Nickname Initial-Letter Biases**

At the conclusion of the experiment, respondents were asked to indicate the first initial letter of their first given name for use in name-letter preferences analyses. Because people oftentimes “go by” a name other than their first given name, participants were asked to indicate if that was the case, and if so, what the first initial of that name was. Of the 492 participants, 62 indicated that they “went by” another name with a different initial than their first given name, while 403 respondents confirmed that the name they “went by” began with the initial letter they already indicated for their first given name.

Twenty-seven respondents did not answer this “nickname” question, either because they discontinued the experiment prematurely or because they chose not to respond to this item. The purpose of the follow-up nickname question was to confirm that participants indeed “went by” their first given name initial in order to accurately identify the target of bias for name-letter preferences analyses. In so doing, an opportunity was presented to examine nickname initial biases separately and as distinct from first given name initial biases, becoming the first implicit egotism study to pursue this differentiation in name-letter targets.

### The Name Letter Test (Nickname Initial)

Nickname-initial analyses were based on 58 respondents who reported a nickname initial and had complete, non-redundant letter ratings on the NLT. Biases for nickname analyses were computed using the I-algorithm (LeBel & Gawronski, 2009; Baccus, Baldwin, & Packer, 2004). The average NLT bias based on participants' nickname initial was  $+.36$  (range =  $-3.61$ – $3.72$ ,  $SD = 1.76$ ). When it came to rating nickname initial letters, participants on average increased their ratings by over one-third of a point on the 1) not at all beautiful to 7) extremely beautiful scale, relative to their ratings of non-nickname initials and as compared to normative ipsatized baseline ratings. This bias, however, did not reach significance,  $t(57) = 1.55$ ,  $p = .13$  (two-tailed), and the effect was small,  $d = .20$ .

### Cartoon 1 (Nickname Initial)

Initial-letter biases for participants with nicknames were also computed using the I-algorithm (LeBel & Gawronski, 2009; Baccus, Baldwin, & Packer, 2004). Analyses were based on 51 respondents who reported a nickname initial and had complete, non-redundant caption ratings for Cartoon 1. The average nickname initial bias exhibited was  $-.38$  (range =  $-3.78$ – $3.80$ ,  $SD = 1.81$ ). Interestingly, when it came to rating captions submitted by writers who shared participants' nickname initial, respondents on average *decreased* their ratings by over one-third of a point on the 1) not at all humorous to 7) extremely humorous scale, relative to their ratings of captions written by non-nickname initial writers and as compared to normative ipsatized baseline ratings. This (negative) bias exhibited toward nickname-initial caption writers was in the opposite direction predicted and was nonsignificant,  $t(50) = -1.52$ ,  $p = .14$  (two-tailed),  $d = -.20$ .

### Cartoon 2 (Nickname Initial)

Initial-letter biases for participants with nicknames were again computed using the I-algorithm (LeBel & Gawronski, 2009; Baccus, Baldwin, & Packer, 2004). Analyses were based on 49 respondents who reported a nickname initial and had complete, non-redundant caption ratings for Cartoon 2. The average nickname initial-letter bias exhibited by participants was  $-.08$  (range =  $-3.58$ – $3.02$ ,  $SD = 1.30$ ). Consistent with the negative biases observed toward Cartoon 1 nickname-initial writers' captions, participants again on average *decreased* their ratings by nearly one-tenth of a point on the 1) not at all humorous to 7) extremely humorous scale, relative to their ratings of captions written by non-nickname initial writers and as compared to normative baseline ratings. Again, this negative bias was opposite to the direction predicted and nonsignificant,  $t(48) = -.41$ ,  $p = .68$  (two-tailed),  $d = -.06$ .

### Both Cartoons (Nickname Initial)

Cartoon 1 and Cartoon 2 nickname initial biases were again negative when averaged together across both sets of cartoon caption ratings ( $N = 45$ ,  $M = -.30$ , range =  $-3.30$ – $1.98$ ,  $SD = 1.03$ ). This effect was opposite to the direction predicted and, interestingly, was actually significant,  $t(44) = -1.98$ ,  $p = .05$  (two-tailed),  $d = -.29$ .

### **“Pure” First Given Name Letter Biases**

Savvy readers will have noted that previous analyses of initial-letter preferences reflected biases toward first given name initials, *irrespective* of the name respondents reported they “went by.” After separating participants who indicated they “went by” another name with a different initial letter in their own analysis (see preceding

“nickname” analyses), the researcher next re-analyzed the original initial-letter preference scores excluding participants who went by a different or “nickname” initial letter.

This was done for two reasons. First, initial-letter biases were vastly different between these two samples. Secondly, the original larger analysis examined first given name initial-letter biases among respondents who did *and* did not go by a name that began with this target letter. Therefore, nicknamed participants were excluded in a follow-up analysis to determine whether a “pure” subset of respondents who indeed went by their first given name would demonstrate even stronger biases when data was not muddled by noise from nicknamed participants.

#### The Name Letter Test (NLT)

Out of the 420 participants in the original NLT analysis, 58 respondents were excluded because they indicated they went by a name and initial that was different from that of their first given name. Twelve additional participants were excluded because they did not answer the question or confirm their first given name initial. A follow-up analysis of NLT biases toward respondent-confirmed first given name initials included 350 participants.

The average initial-letter bias was +1.42 for this subset of respondents (range = -2.83–6.05,  $SD = 1.61$ ). When it came to ratings of participants’ first initials *for their first given name*, participants rated that initial letter 1.42 points higher on the attractiveness scale, relative to non-initials and as compared to normative baselines. This observed bias was significant,  $t(349) = 16.44$ ,  $p < .001$ , and the effect is considered large,  $d = .88$ . Recall that this bias is very similar to the one demonstrated by the larger set of

respondents ( $N = 420$ ), which was based on first given name initial targets of participants who both went by their first given name *and* those who went by a different name that began with a different letter,  $+1.38$  (range =  $-3.43$ – $6.05$ ,  $SD = 1.63$ ),  $t(419) = 17.28$ ,  $p < .001$ ,  $d = .84$ . Based on previous and follow-up analyses of NLT scores, one can argue that initial-letter biases are relevant to first given name initials only, even when a person “goes by” another name beginning with a different letter.

#### Cartoon 1

Follow-up analyses of Cartoon 1 initial-letter biases were also conducted to determine whether a “pure” subset of participants who indeed “went by” their first given name (and the target initial used in the computation of previous initial-letter biases) would demonstrate these biases after excluding respondents who went by a name other than their first given one (and thus a different target initial that was used in the analysis). Out of the 428 participants in the original Cartoon 1 analysis, 54 respondents were excluded because they indicated they went by a name and initial that was different from their first given name. Thirteen additional participants were excluded because they did not answer the question or confirm their first given name initial. Follow-up analyses of Cartoon 1 initial-letter biases included 361 respondents who confirmed that they “went by” their first given name.

The average initial-letter bias exhibited by this subset of participants was  $+1.16$  (range =  $-3.46$ – $5.09$ ,  $SD = 1.61$ ). When it came to rating captions submitted by writers who shared their first initial *of their first given name*, respondents rated that caption .16 points higher on the humor scale, relative to captions submitted by writers who did not share their first given name initial and as compared to normative baselines. This finding

was marginally significant,  $t(360) = 1.89, p = .06$  (two-tailed) and the effect is considered very small,  $d = .10$ . Recall that these results were similar in findings for the larger set of participants, which included those who both went by their first given name *and* those who went by a nickname ( $N = 428$ ). Their average initial-letter bias was  $+.18$  (range =  $-3.46$ – $5.09$ ,  $SD = 1.59$ ), which was significant,  $t(427) = 2.35, p = .02$  (two-tailed) and effect sizes for both samples were small and nearly identical,  $d = .11$ . Based on previous and follow-up analyses of Cartoon 1 caption ratings, one can again argue that initial-letter biases are relevant to first given name initials only—even when a person “goes by” another name beginning with a different letter.

#### Cartoon 2

Out of the 429 participants from the original Cartoon 2 analysis, 53 respondents were excluded because they indicated they went by a name and initial that was different from their first given name. Twelve additional participants were excluded because they did not answer the question or confirm their first given name initial. Follow-up analyses of Cartoon 2 initial-letter biases based on respondent-confirmed first given name initials included 364 participants who indicated they “went by” their first given name.

The new average initial-letter bias was  $+.13$  (range =  $-3.81$ – $4.47$ ,  $SD = 1.50$ ). When it came to rating captions submitted by writers who shared their first initial *of their first given name*, participants rated that caption .13 points higher on the humor scale relative to captions submitted by writers who did not share their first given name initial, and as compared to normative baselines. This finding, although in the direction predicted, did not reach statistical significance,  $t(363) = 1.66, p = .10$  (two-tailed),  $d = .09$ . Recall that these results are similar in findings to those found among the larger set of

participants ( $N = 429$ ), which examined biases toward initials for respondents who both went by their first given name *and* those who went by a nickname,  $+ .10$  (range =  $-3.81$ – $4.47$ ,  $SD = 1.48$ ),  $t(428) = 1.35$ ,  $p = .18$ , two-tailed,  $d = .07$ .

Based on previous and follow-up analyses of Cartoon 2 initial-letter biases (and although nonsignificant), one can again argue that these biases were relevant to first given name initials only, even when a person “went by” another name beginning with a different letter.

#### Both Cartoons

This follow-up analysis was based on 333 respondents who indicated that they “went by” their first given name. Out of the 394 participants from the original analysis, 49 were excluded because they indicated they went by a name and initial that was different from their first given name. Twelve additional respondents were excluded because they did not answer the question or confirm their first given name initial letter.

The average initial-letter bias for this subset of participants was  $+ .15$  (range =  $-2.19$ – $3.12$ ,  $SD = 1.09$ ). When it came to rating captions submitted by writers who shared their first initial *of their first given name*, respondents rated that caption  $.15$  points higher on the humor scale, relative to captions submitted by writers who did not share their first given name initial and as compared to normative baseline ratings. This observed bias was significant,  $t(332) = 2.47$ ,  $p = .01$  (two-tailed),  $d = .14$ . Recall that this bias was identical to the first given name initial-letter bias exhibited by the larger sample of participants ( $N = 394$ ), which included those who went by their first given name, as well as those who went by a nickname ( $M = +.15$ , range =  $-2.46$ – $3.12$ ,  $SD = 1.08$ ). This original initial-

letter bias was also statistically significant,  $t(393) = 2.67, p = .01$  (two-tailed), and the magnitude of the effect was identical,  $d = .14$ .

In sum, based on previous and follow-up analyses of biases exhibited on the NLT, Cartoon 1 and Cartoon 2 (individually and when averaged together), the findings of this study suggest that initial-letter biases are relevant to first given name initials only, even when a person “goes by” another name beginning with a different initial. Moreover, biases toward nickname initials were reduced and nonsignificant on the NLT, and were negative on the cartoon caption rating task (which were statistically significant when averaged together for Cartoon 1 and 2).



## CHAPTER 6

### EXPLICIT SELF-ESTEEM AND SELF-CONCEPT THREAT CONDITIONS

Explicit self-esteem was measured using the Rosenberg Self-Esteem Scale (RSES; Rosenberg, 1989), a widely used ten-item self-report index that assesses respondents' overall level of self-value and self-worth. Both the timing of the RSES administration and its scores were examined in relation to initial-letter preferences. For the 474 participants who completed all ten items on the RSES, the average score was 20.72 (range = 5–30,  $SD = 5.02$ ), based on a total possible score of 30 with 3 points for each question. Half of the items (#1, #3, #4, #7, and #10) are scored as follows: “strongly agree” = 3 points, “agree” = 2 points, “disagree” = 1 point, and “strongly disagree” = 0 points. The remaining half of items (#2, #5, #6, #8, and #9) are reversed in valence and scored as: “strongly agree” = 0 points, “agree” = 1 point, “disagree” = 2 points, and “strongly disagree” = 3 points. For a list of RSES items, please refer to Appendix A.

#### **Explicit Self-Esteem Gender Differences**

For men ( $N = 161$ ), the average RSES score was 21.22 (range = 7–30,  $SD = 4.94$ ) and for women ( $N = 313$ ), the average score was 20.47 (range = 5–30,  $SD = 5.05$ ). Differences in men's and women's explicit self-esteem were not statistically significant,  $t(472) = 1.54$ ,  $p = .13$  (two-tailed),  $d = .15$ .

### Explicit Self-Esteem and Initial-Letter Biases

The influence of explicit self-esteem on initial-letter biases was next examined as a continuous variable (total score on the RSES) for those with complete data on this measure. First, initial-letter biases demonstrated on the Name Letter Test (NLT) were regressed on RSES scores.

#### The Name Letter Test

Explicit self-esteem, as measured by RSES scores, did not predict differences in NLT biases based on first given name initials,  $R^2 = .001$ ,  $F(1, 408) = .34$ ,  $p = .56$ . It also failed to predict NLT biases in analyses based on nickname initials,  $R^2 = .002$ ,  $F(1, 55) = .13$ ,  $p = .72$ .

#### Cartoon Caption Ratings

When initial-letter preferences for Cartoon 1 were regressed on RSES scores, explicit self-esteem approached significance as a predictor of biases exhibited toward captions written by writers who shared participants' first given name initial,  $R^2 = .01$ ,  $F(1, 417) = 3.26$ ,  $p = .07$ . Explicit self-esteem was not a significant predictor of Cartoon 1 initial-letter biases when analyses were based on respondents' nickname initial,  $R^2 = .01$ ,  $F(1, 48) = .28$ ,  $p = .60$ .

Cartoon 2 initial-letter preference scores were also regressed on RSES scores, but explicit self-esteem—as measured by total RSES scores—did not predict biases exhibited toward captions written by writers who shared participants' first given name initial,  $R^2 = .003$ ,  $F(1, 419) = 1.22$ ,  $p = .27$ , or nickname initial,  $R^2 = .03$ ,  $F(1, 47) = 1.49$ ,  $p = .23$ .

When initial-letter biases were averaged together for Cartoon 1 and Cartoon 2, explicit self-esteem was not a significant predictor of biases demonstrated toward

captions submitted by writers who shared respondents' first given name initials, however, it did approach significance,  $R^2 = .01$ ,  $F(1, 385) = 3.33$ ,  $p = .07$ . Explicit self-esteem did not predict average Cartoon 1 and Cartoon 2 initial-letter biases when analyses were based on participants' nickname initial,  $R^2 = .03$ ,  $F(1, 43) = 1.27$ ,  $p = .26$ .

### **Threat Condition Manipulation Check**

A total of 492 men and women respondents were randomly assigned to complete a brief writing task in one of three different experimental conditions: self-concept threat ( $N = 158$ ); self-concept affirmation ( $N = 163$ ); or a control condition ( $N = 171$ ). The researcher evaluated participants' responses to the writing task in the three different threat conditions according to a) the number of sentences written, b) whether they followed directions, and c) whether they provided complete responses.

Recall that following Jones, Pelham, Mirenberg, and Hetts (2002), participants who were randomly assigned to the self-concept threat condition were asked to write at least three sentences about an aspect of themselves that they have found difficult to change and would like to be different. The aspect should have reflected something important about themselves that they wished they could change, but have not been able to. In the self-concept affirmation condition, respondents were asked to write at least three sentences about an important area of their life where they have always felt good about themselves, which represents a positive, important, and stable aspect of who they are. Finally, participants randomly assigned to the control condition were asked to simply write at least three sentences about the last movie they saw.

On average, respondents wrote 3.19 sentences (range = 1–7,  $SD = .67$ ), based on 486 participants who completed the writing task for one of the three experimental

conditions. Of these, 469 participants wrote at least three sentences as instructed, 10 wrote only two sentences, and 7 respondents wrote just one sentence. Further analysis of the content of responses revealed that 461 participants wrote at least three sentences that were relevant to the threat condition to which they were assigned, 17 participants wrote less than three relevant and complete sentences, 4 did not follow the instructions<sup>1</sup> and another 4 respondents had misaligned or duplicate data. Only those participants who followed directions by writing at least three sentences on a topic relevant to the experimental condition to which they were assigned were included in self-concept threat analyses ( $N = 461$ ).

A one-way ANOVA was performed to determine whether the number of sentences respondents wrote significantly differed between the experimental conditions. Participants randomly assigned to the self-concept threat condition ( $N = 143$ ) wrote an average of 3.20 sentences ( $SD = .51$ ). Those randomly assigned to the self-concept affirmation condition ( $N = 154$ ) wrote an almost identical number of sentences ( $M = 3.19$ ,  $SD = .59$ ), while participants randomly assigned to the control condition ( $N = 164$ ) wrote an average of 3.32 sentences ( $SD = .67$ ). Differences in the average number of sentences written among the three threat conditions approached significance,  $F(2, 458) = 2.66$ ,  $p = .07$ ,  $\eta^2 = .01$ , with participants in the control condition (who were asked to write about a recent movie they saw) writing the most number of sentences.

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<sup>1</sup> All 4 of these participants were in the self-concept threat condition and instead of writing about a personal flaw, they either a) put a positive spin on a shortcoming or b) indicated that there was nothing about themselves that they wished to change. Neither of these outcomes represented a threat to the self-concept, and thus these respondents' data were not included in threat analyses.

Subsequent to the writing task, respondents were asked to rate their current mood and how they were feeling at the present moment using a scale from 1) extremely negative to 7) extremely positive (Jones et al., 2002). Two participants did not answer both questions, so they were excluded from analyses. The average rating given by participants across conditions was 4.92 for self-assessments of current mood ( $N = 459$ , range = 1–7,  $SD = 1.08$ ) and 4.72 for self-assessments of how they were currently feeling ( $N = 459$ , range = 1–7,  $SD = 1.15$ ).

On average, participants who were randomly assigned to the self-concept threat condition and completed the mood item ( $N = 142$ ) rated their current mood a 4.80 (range = 1–7,  $SD = 1.13$ ). Those assigned to the self-concept affirmation condition ( $N = 154$ ) rated their current mood a 4.93 (range = 2–7,  $SD = 1.10$ ), and the average rating for participants in the control condition ( $N = 163$ ) was 5.01 (range = 2–7,  $SD = 1.00$ ). A one-way ANOVA did not reveal any significant differences between the average mood ratings assigned by respondents in the self-concept threat, affirmation, or control conditions,  $F(2, 456) = 1.37, p = .25, \eta^2 = .006$ .

Participants in the self-concept threat condition assigned an average rating of 4.64 (range = 1–7,  $SD = 1.11$ ) to how they were feeling at the current moment. Those in the self-concept affirmation condition, on average, rated how they were currently feeling a 4.72 (range 2–7,  $SD = 1.15$ ), while those in the control condition had an average rating of 4.78 (range = 1–7,  $SD = 1.19$ ). A one-way ANOVA again did not reveal any significant differences in how respondents in the different experimental conditions rated they were currently feeling,  $F(2, 456) = .55, p = .58, \eta^2 = .002$ . Based on analyses of the two items

in the mood questionnaire, one can infer that the experimental manipulation was a valid threat to the self-concept, instead of merely a manipulation of mood.

### **The Name Letter Test**

#### **Differences in Biases Across Threat Conditions**

Respondents who wrote at least three sentences relevant to the experimental threat condition to which they were randomly assigned and those with complete data on the Name Letter Test (NLT) were included in these first analyses ( $N = 396$ ). The average initial-letter bias on the NLT for these participants was +1.34 (range =  $-3.43$ – $5.64$ ,  $SD = 1.63$ ).

Participants in the self-concept threat condition ( $N = 119$ ) had an average NLT bias of +1.41 (range =  $-2.83$ – $5.42$ ,  $SD = 1.60$ ). Those assigned to the self-concept affirmation condition ( $N = 132$ ) demonstrated an average bias of +1.29 (range =  $-2.79$ – $4.88$ ,  $SD = 1.66$ ), while respondents assigned to the control condition ( $N = 145$ ) exhibited an average bias of +1.34 (range =  $-3.43$ – $5.64$ ,  $SD = 1.64$ ). A one-way ANOVA was used to test for differences in NLT initial-letter biases between the three threat conditions and was nonsignificant,  $F(2, 393) = .16$ ,  $p = .86$ ,  $\eta^2 = .0008$ .

#### **Interaction of Threat Condition $\times$ Explicit Self-Esteem**

To examine whether the regression of NLT biases on threat condition varied as a function of participants' explicit self-esteem scores, the researcher next used Aiken & West's (1991) methods for testing for interactions. This analysis was based on data from respondents who completed the NLT, threat condition writing task, and those with complete data on the Rosenberg Self-Esteem Scale (RSES). The average initial-letter bias for this group of participants ( $N = 388$ ) was +1.33 (range =  $-3.43$ – $5.64$ ,  $SD = 1.62$ ).

First, each participant's RSES score was centered by subtracting the sample mean from each score. The RSES sample mean was 20.74 (range 5–30,  $SD = 5.03$ ) for this analysis. Next, threat condition was dummy-coded with the control group serving as the comparison group. A regression analysis was performed to test the  $b_1$  coefficient, which compared the means for the control and self-concept threat conditions, and the  $b_2$  coefficient, which compared the means for the control and self-concept affirmation conditions. Results were similar in findings and logic to the one-way ANOVA reported above, i.e., threat condition did not predict NLT biases,  $R^2 = .001$ ,  $F(2, 385) = .11$ ,  $p = .89$ .

Next, the effect of explicit self-esteem as measured by centered RSES scores was added to the multiple regression equation. RSES scores did not help predict NLT initial-letter biases when added to the model,  $R^2_{change} = .000$ ,  $F(1, 384) = .01$ ,  $p = .91$ . Finally, the interactions of RSES scores and the two threat condition comparisons were added to the model to determine whether NLT biases, when regressed on threat condition, varied as a function of explicit self-esteem. The overall model including the dummy variables (threat conditions), continuous variable (centered RSES scores), and their interactions was nonsignificant,  $R^2 = .008$ ,  $F(5, 382) = .60$ ,  $p = .70$ . Joint tests of the  $b_4$  and  $b_5$  coefficients (corresponding to the interaction of the self-concept threat vs. control condition  $\times$  RSES scores, and the interaction of the self-concept affirmation vs. control condition  $\times$  RSES scores, respectively), did not significantly improve the multiple regression model,  $R^2_{change} = .007$ ,  $F(2, 382) = 1.39$ ,  $p = .25$ .

In sum, when initial-letter biases exhibited on the NLT were regressed on threat condition, biases did not vary as a function of participants' explicit self-esteem. Because

there was no presence of an interaction between threat conditions and RSES scores, a simple slopes test to examine the nature of an interaction was not warranted. Please refer to Table 6 for the progression of the multiple regression equations regressing NLT biases on threat condition as a function of explicit self-esteem.

Table 6. Progression of multiple regression equations of NLT initial-letter biases on threat condition as a function of explicit self-esteem

a. Test of dummy variables only

$$\hat{Y} = b_1D_1 + b_2D_2 + b_0$$

$$\hat{Y} = (-.018)(D_1) + (-.089)(D_2) + 1.363$$

$$\text{Joint test of } b_1, b_2: R^2 = .001, F(2, 385) = .11, p = .89$$

$$\text{Test of } b_1: t(385) = -.09, p = .93$$

$$\text{Test of } b_2: t(385) = -.46, p = .65$$

$$SS_{reg} = .594; SS_{res} = 1,014.897$$

b. Test of dummy variables and continuous variable

$$\hat{Y} = b_1D_1 + b_2D_2 + b_3RSES + b_0$$

$$\hat{Y} = (-.016)(D_1) + (-.089)(D_2) + (.002)RSES + 1.362$$

$$\text{Joint test of } b_1, b_2, b_3: R^2 = .001, F(3, 384) = .08, p = .97$$

$$\text{Test of } b_3: R^2_{change} = .000, F(1, 384) = .01, p = .91$$

$$SS_{reg} = .629; SS_{res} = 1,014.863$$

c. Test of dummy variables, continuous variable, and interaction

$$\hat{Y} = b_1D_1 + b_2D_2 + b_3RSES + b_4(D_1 \times RSES) + b_5(D_2 \times RSES) + b_0$$

$$\hat{Y} = (-.029)D_1 + (-.108)D_2 + (-.034)RSES + (.043)(D_1 \times RSES) + (.067)(D_2 \times RSES) + 1.380$$

$$\text{Joint test of } b_1-b_5: R^2 = .008, F(5, 382) = .60, p = .70$$

$$\text{Joint test of } b_4, b_5: R^2_{change} = .007, F(2, 382) = 1.39, p = .25$$

$$\text{Test of } b_3: t(382) = -1.20, p = .23$$

$$\text{Test of } b_4: t(382) = 1.09, p = .28$$

$$\text{Test of } b_5: t(382) = 1.64, p = .10$$

$$SS_{reg} = 7.963; SS_{res} = 1,007.529$$

Based on methods by Aiken & West (1991); dummy variables = self-concept threat conditions, continuous variable = centered RSES score, and the interaction of the two.  $N = 388$ , observed statistical power for final model = .88.



## Cartoon 1 Initial-Letter Biases

### Differences in Biases Across Threat Conditions

Respondents who wrote at least three sentences relevant to the experimental threat condition to which they were randomly assigned and those with complete ratings for the entire set of Cartoon 1 captions were included in these first analyses ( $N = 404$ ). The average Cartoon 1 initial-letter bias for these participants was  $+.20$  (range =  $-3.46$ – $5.09$ ,  $SD = 1.58$ ).

Participants in the self-concept threat condition ( $N = 126$ ) demonstrated an average bias of  $+.25$  (range =  $-3.39$ – $3.49$ ,  $SD = 1.49$ ). Those assigned to the self-concept affirmation condition ( $N = 133$ ) exhibited an average bias of  $+.13$  (range =  $-3.46$ – $4.44$ ,  $SD = 1.62$ ), while participants assigned to the control condition ( $N = 145$ ) had an average bias of  $+.21$  (range =  $-3.13$ – $5.09$ ,  $SD = 1.63$ ). A one-way ANOVA was used to test for differences in initial-letter biases between the three threat conditions and was nonsignificant,  $F(2, 401) = .21$ ,  $p = .81$ ,  $\eta^2 = .001$ .

### Interaction of Threat Condition $\times$ Explicit Self-Esteem

This next analysis was based on data from respondents who completed one of the threat condition writing tasks, those with complete data on the RSES, and participants who had complete ratings for the entire set of twenty Cartoon 1 captions ( $N = 396$ ). The average initial-letter bias for captions submitted by writers who shared participants' first initial was  $+.22$  for this analysis (range =  $-3.46$ – $5.09$ ,  $SD = 1.58$ ) and the average RSES score was  $20.71$  (range =  $5$ – $30$ ,  $SD = 4.97$ ).

To examine whether the regression of Cartoon 1 initial-letter biases on threat condition varied as a function of participants' explicit self-esteem scores, the researcher

again used Aiken & West's (1991) multiple regression methods. First, each respondent's RSES score was centered by subtracting the sample mean from each score, which was 20.71 (range = 5–30,  $SD = 4.97$ ) for this sample. Next, threat condition was dummy-coded with the control group serving as the comparison group. A regression analysis was performed to test the  $b_1$  coefficient, which compared the mean initial-letter biases for the control and self-concept threat conditions, and the  $b_2$  coefficient, which compared the mean biases for the control and self-concept affirmation conditions.

Previously, threat condition did not predict differences in NLT biases. In the current analysis, threat condition also did not predict biases toward captions submitted by writers who shared participants' first initial, with results similar in findings and logic to the one-way ANOVA reported above, i.e., threat condition did not predict Cartoon 1 initial-letter biases,  $R^2 = .002$ ,  $F(2, 393) = .41$ ,  $p = .66$ .

Next, the effect of explicit self-esteem as measured by centered RSES scores was added to the multiple regression equation. Again, RSES scores did not help predict initial-letter biases for Cartoon 1 when this variable was added to the model,  $R^2_{change} = .007$ ,  $F(1, 392) = 2.97$ ,  $p = .09$ . Finally, the interactions of RSES scores and the two threat condition comparisons were added to the model to determine whether biases, when regressed on threat condition, varied as a function of participants' explicit self-esteem. The overall model including the dummy variables (threat conditions), continuous variable (centered RSES scores), and their interactions was not significant,  $R^2 = .010$ ,  $F(5, 390) = .79$ ,  $p = .56$ . Joint tests of the  $b_4$  and  $b_5$  coefficients (corresponding to the interaction of the self-concept threat vs. control condition  $\times$  RSES scores, and the interaction of the self-concept affirmation vs. control condition  $\times$  RSES scores, respectively), did not

significantly improve the multiple regression model,  $R^2_{change} = .000$ ,  $F(2, 390) = .09$ ,  $p = .92$ . Because there was no presence of interactions between threat conditions and RSES scores, a simple slopes test to examine the nature of an interaction was not performed. In sum, when initial-letter biases for Cartoon 1 were regressed onto threat condition, biases did not vary as a function of participants' explicit self-esteem. Please refer to Table 7 for the progression of equations regressing initial-letter biases for Cartoon 1 on threat condition as a function of explicit self-esteem.

### **Cartoon 2 Initial-Letter Biases**

#### **Differences in Biases Across Threat Conditions**

Respondents who wrote at least three sentences relevant to the experimental threat condition to which they were randomly assigned and those with complete ratings for the entire set of Cartoon 2 captions were included in these first analyses ( $N = 408$ ). The average initial-letter bias for these participants was  $+.10$  (range =  $-3.81$ – $4.47$ ,  $SD = 1.47$ ).

Participants in the self-concept threat condition ( $N = 126$ ) demonstrated an average bias of  $+.11$  (range =  $-3.68$ – $3.48$ ,  $SD = 1.41$ ). Those assigned to the self-concept affirmation condition ( $N = 141$ ) also had an average bias of  $+.11$  (range =  $-2.62$ – $3.42$ ,  $SD = 1.35$ ), while participants assigned to the control condition ( $N = 141$ ) exhibited an average bias of  $+.08$  (range =  $-3.81$ – $4.47$ ,  $SD = 1.65$ ). A one-way ANOVA was used to test for differences in initial-letter biases between the three threat conditions and was nonsignificant,  $F(2, 405) = .02$ ,  $p = .98$ ,  $\eta^2 = .00009$ .

#### **Interaction of Threat Condition $\times$ Explicit Self-Esteem**

This next analysis was based on data from respondents who completed a threat condition writing task, those who had complete ratings for the entire set of twenty

Cartoon 2 captions, and those with complete data on the RSES ( $N = 401$ ). The average bias exhibited toward Cartoon 2 captions submitted by writers who shared participants' first initial was +.11 for this sample (range =  $-3.81$ – $4.47$ ,  $SD = 1.47$ ). The average score on the RSES was 20.72 (range 5–30,  $SD = 5.03$ ), which was used to center participants' explicit self-esteem scores prior to multiple regression analyses.

Table 7. Progression of multiple regression equations of Cartoon 1 initial-letter biases on threat condition as a function of explicit self-esteem

a. Test of dummy variables only

$$\hat{Y} = b_1D_1 + b_2D_2 + b_0$$

$$\hat{Y} = (.068)(D_1) + (-.110)(D_2) + .236$$

$$\text{Joint test of } b_1, b_2: R^2 = .002, F(2, 393) = .41, p = .66$$

$$\text{Test of } b_1: t(393) = .35, p = .73$$

$$\text{Test of } b_2: t(393) = -.58, p = .56$$

$$SS_{reg} = 2.048; SS_{res} = 978.107$$

b. Test of dummy variables and continuous variable

$$\hat{Y} = b_1D_1 + b_2D_2 + b_3RSES + b_0$$

$$\hat{Y} = (.083)(D_1) + (-.103)(D_2) + (.027)RSES + .229$$

$$\text{Joint test of } b_1, b_2, b_3: R^2 = .010, F(3, 392) = 1.26, p = .29$$

$$\text{Test of } b_3: R^2_{change} = .007, F(1, 392) = 2.97, p = .09$$

$$SS_{reg} = 9.393; SS_{res} = 970.762$$

c. Test of dummy variables, continuous variable, and interaction

$$\hat{Y} = b_1D_1 + b_2D_2 + b_3RSES + b_4(D_1 \times RSES) + b_5(D_2 \times RSES) + b_0$$

$$\hat{Y} = (.086)D_1 + (-.102)D_2 + (.031)RSES + (.001)(D_1 \times RSES) + (-.014)(D_2 \times RSES) + .228$$

$$\text{Joint test of } b_1 - b_5: R^2 = .010, F(5, 390) = .79, p = .56$$

$$\text{Joint test of } b_4, b_5: R^2_{change} = .000, F(2, 390) = .09, p = .92$$

$$\text{Test of } b_3: t(390) = 1.15, p = .25$$

$$\text{Test of } b_4: t(390) = .04, p = .97$$

$$\text{Test of } b_5: t(390) = -.35, p = .73$$

$$SS_{reg} = 9.835; SS_{res} = 970.320$$

Based on methods by Aiken & West (1991); dummy variables = self-concept threat conditions, continuous variable = centered RSES score, and the interaction of the two.  $N = 396$ , observed statistical power for final model = .85.

Aiken & West's (1991) multiple regression methods were again used to examine whether the regression of Cartoon 2 initial-letter biases on threat condition varied as a function of participants' explicit self-esteem. The same dummy codes for the three threat conditions in the prior two analyses were again used in this analysis. Previously, threat condition did not predict differences in NLT initial-letter biases or differences in biases toward Cartoon 1 captions submitted by same-initial writers. In the current analysis, threat condition also did not predict biases toward Cartoon 2 captions submitted by writers with whom respondents shared a first initial,  $R^2 = .000$ ,  $F(2, 398) = .05$ ,  $p = .95$ . This analysis was similar in logic and result to the ANOVA reported above.

The effect of explicit self-esteem as measured by centered RSES scores was also nonsignificant when added to the regression model,  $R^2_{change} = .003$ ,  $F(1, 397) = 1.08$ ,  $p = .30$ . The overall model including the dummy variables (threat conditions), continuous variable (centered RSES scores), and their interactions was nonsignificant,  $R^2 = .006$ ,  $F(5, 395) = .48$ ,  $p = .79$ . Adding the interaction terms did not significantly improve the model,  $R^2_{change} = .003$ ,  $F(2, 395) = .61$ ,  $p = .54$ . In sum, when Cartoon 2 biases were regressed on threat condition, biases did not vary as a function of participants' explicit self-esteem. For the progression of multiple regression equations, please refer to Table 8.

### **Average Cartoon 1 and Cartoon 2 Initial-Letter Biases**

#### **Differences in Biases Across Threat Conditions**

Participants who wrote at least three sentences relevant to the experimental threat condition to which they were randomly assigned and those with complete ratings for the entire set of Cartoon 1 *and* Cartoon 2 captions were included in these first analyses ( $N =$

373). The average initial-letter bias for these participants was +.16 (range = -2.19–3.07,  $SD = 1.07$ ).

Table 8. Progression of multiple regression equations of Cartoon 2 initial-letter biases on threat condition as a function of explicit self-esteem

a. Test of dummy variables only

$$\hat{Y} = b_1D_1 + b_2D_2 + b_0$$

$$\hat{Y} = (.056)(D_1) + (.019)(D_2) + .090$$

$$\text{Joint test of } b_1, b_2: R^2 = .000, F(2, 398) = .05, p = .95$$

$$\text{Test of } b_1: t(398) = .31, p = .76$$

$$\text{Test of } b_2: t(398) = .11, p = .92$$

$$SS_{reg} = .208; SS_{res} = 858.929$$

b. Test of dummy variables and continuous variable

$$\hat{Y} = b_1D_1 + b_2D_2 + b_3RSES + b_0$$

$$\hat{Y} = (.066)(D_1) + (.027)(D_2) + (.015)RSES + .084$$

$$\text{Joint test of } b_1, b_2, b_3: R^2 = .003, F(3, 397) = .39, p = .76$$

$$\text{Test of } b_3: R^2_{change} = .003, F(1, 397) = 1.08, p = .30$$

$$SS_{reg} = 2.542; SS_{res} = 856.594$$

c. Test of dummy variables, continuous variable, and interaction

$$\hat{Y} = b_1D_1 + b_2D_2 + b_3RSES + b_4(D_1 \times RSES) + b_5(D_2 \times RSES) + b_0$$

$$\hat{Y} = (.066)D_1 + (.034)D_2 + (.031)RSES + (-.037)(D_1 \times RSES) + (-.010)(D_2 \times RSES) + .077$$

$$\text{Joint test of } b_1 - b_5: R^2 = .006, F(5, 395) = .48, p = .79$$

$$\text{Joint test of } b_4, b_5: R^2_{change} = .003, F(2, 395) = .61, p = .54$$

$$\text{Test of } b_3: t(395) = 1.25, p = .21$$

$$\text{Test of } b_4: t(395) = -1.06, p = .29$$

$$\text{Test of } b_5: t(395) = -.26, p = .79$$

$$SS_{reg} = 5.178; SS_{res} = 853.959$$

Based on methods by Aiken & West (1991); dummy variables = self-concept threat conditions, continuous variable = centered RSES score, and the interaction of the two.  $N = 401$ , observed statistical power for final model = .90.

Respondents in the self-concept threat condition ( $N = 117$ ) demonstrated an average bias of  $+.18$  (range =  $-2.12$ – $2.39$ ,  $SD = 1.02$ ). Those assigned to the self-concept affirmation condition ( $N = 123$ ) had an average bias of  $+.16$  (range =  $-2.19$ – $3.02$ ,  $SD = 1.06$ ), while participants assigned to the control condition ( $N = 133$ ) exhibited an average bias of  $+.14$  (range =  $-2.13$ – $3.07$ ,  $SD = 1.14$ ). A one-way ANOVA was used to test for differences in initial-letter biases between the three threat conditions and was nonsignificant,  $F(2, 370) = .05$ ,  $p = .96$ ,  $\eta^2 = .0002$ .

#### Interaction of Threat Condition $\times$ Explicit Self-Esteem

Next, initial-letter biases for Cartoon 1 and Cartoon 2 were averaged together and regressed onto threat condition to determine whether biases varied as a function of participants' explicit self-esteem. Aiken & West's (1991) multiple regression methods were again used and results are based on data from respondents who completed a threat condition writing task, those who had complete ratings for Cartoon 1 and Cartoon 2 captions, and those with complete data on the RSES ( $N = 367$ ). The average bias exhibited toward Cartoon 1 and 2 captions submitted by writers who shared participants' first initial was  $+.17$  for this sample (range =  $-2.19$ – $3.07$ ,  $SD = 1.07$ ). The average score on the RSES was  $20.78$  (range  $5$ – $30$ ,  $SD = 5.00$ ) and was used to center participants' explicit self-esteem scores prior to multiple regression analyses.

The same dummy codes from the prior multiple regression analyses were used in the current one. Previously, threat condition did not predict differences in NLT initial-letter biases, or—individually—biases toward Cartoon 1 or 2 captions submitted by same-initial writers. In the current analysis, threat condition also failed to predict average

biases demonstrated toward Cartoon 1 and Cartoon 2 captions submitted by writers with whom participants shared a first initial,  $R^2 = .001$ ,  $F(2, 364) = .12$ ,  $p = .89$ .

RSES scores also did not significantly predict average Cartoon 1 Cartoon 2 initial-letter biases when this variable was added to the model,  $R^2_{change} = .008$ ,  $F(1, 363) = 2.97$ ,  $p = .09$ . The overall model including the dummy variables (threat conditions), continuous variable (centered RSES scores), and their interactions was nonsignificant,  $R^2 = .012$ ,  $F(5, 361) = .85$ ,  $p = .52$ . Joint tests of the interaction terms were nonsignificant as well,  $R^2_{change} = .003$ ,  $F(2, 361) = .52$ ,  $p = .60$ , but the test of the  $b_3$  coefficient for the contribution of RSES scores to the final multiple regression model approached significance,  $t(361) = 1.84$ ,  $p = .07$ . For the progression of multiple regression equations regressing initial-letter biases for both Cartoon 1 and Cartoon 2 on threat condition as a function of participants' explicit self-esteem, please refer to Table 9.

### **The Name Letter Test (Nicknames)**

#### **Differences in Biases Across Threat Conditions**

Respondents who wrote at least three sentences relevant to the experimental threat condition to which they were randomly assigned, those with complete ratings on the NLT, and those who indicated an alternate initial for the name they “went by” that was different from that of their first given name were included in these first analyses ( $N = 53$ ). The average nickname initial-letter bias demonstrated by these participants was  $+.28$  (range =  $-3.61$ – $3.72$ ,  $SD = 1.79$ ).

Participants in the self-concept threat condition ( $N = 16$ ) demonstrated an average bias of  $-.16$  (range =  $-2.54$ – $2.99$ ,  $SD = 1.75$ ). Those assigned to the self-concept affirmation condition ( $N = 20$ ) had an average bias of  $+.08$  (range =  $-3.61$ – $2.71$ ,  $SD =$



1.87), while respondents assigned to the control condition ( $N = 17$ ) exhibited an average bias of  $+ .93$  (range =  $-2.37$ – $3.72$ ,  $SD = 1.63$ ). A one-way ANOVA was used to test for differences in initial-letter biases between the three threat conditions and was nonsignificant,  $F(2, 50) = 1.79$ ,  $p = .18$ ,  $\eta^2 = .07$ .

Table 9. Progression of multiple regression equations of average Cartoon 1 and Cartoon 2 initial-letter biases on threat condition as a function of explicit self-esteem

a. Test of dummy variables only

$$\hat{Y} = b_1D_1 + b_2D_2 + b_0$$

$$\hat{Y} = (.061)(D_1) + (.003)(D_2) + .150$$

$$\text{Joint test of } b_1, b_2: R^2 = .001, F(2, 364) = .12, p = .89$$

$$\text{Test of } b_1: t(364) = .45, p = .65$$

$$\text{Test of } b_2: t(364) = .03, p = .98$$

$$SS_{reg} = .280; SS_{res} = 416.510$$

b. Test of dummy variables and continuous variable

$$\hat{Y} = b_1D_1 + b_2D_2 + b_3RSES + b_0$$

$$\hat{Y} = (.071)(D_1) + (.009)(D_2) + (.019)RSES + .146$$

$$\text{Joint test of } b_1, b_2, b_3: R^2 = .009, F(3, 363) = 1.07, p = .36$$

$$\text{Test of } b_3: R^2_{change} = .008, F(1, 363) = 2.97, p = .09$$

$$SS_{reg} = 3.657; SS_{res} = 413.133$$

c. Test of dummy variables, continuous variable, and interaction

$$\hat{Y} = b_1D_1 + b_2D_2 + b_3RSES + b_4(D_1 \times RSES) + b_5(D_2 \times RSES) + b_0$$

$$\hat{Y} = (.073)D_1 + (.012)D_2 + (.035)RSES + (-.022)(D_1 \times RSES) + (-.025)(D_2 \times RSES) + .142$$

$$\text{Joint test of } b_1 - b_5: R^2 = .012, F(5, 361) = .85, p = .52$$

$$\text{Joint test of } b_4, b_5: R^2_{change} = .003, F(2, 361) = .52, p = .60$$

$$\text{Test of } b_3: t(361) = 1.84, p = .07$$

$$\text{Test of } b_4: t(361) = -.84, p = .40$$

$$\text{Test of } b_5: t(361) = -.91, p = .37$$

$$SS_{reg} = 4.845; SS_{res} = 411.946$$

Based on methods by Aiken & West (1991); dummy variables = self-concept threat conditions, continuous variable = centered RSES score, and the interaction of the two.  $N = 367$ , observed statistical power for final model = .85.

### Interaction of Threat Condition $\times$ Explicit Self-Esteem

For this analysis, initial-letter biases were based on participants' biases toward their nickname initial. Data were included from respondents who a) indicated they "went by" a name other than their first given name such as a "nickname," b) indicated their nickname initial, c) completed a threat condition writing task, d) had complete data on the NLT, and e) had complete data on the RSES ( $N = 52$ ). The average initial-letter bias for participants in this sample on the NLT was  $+ .24$  (range  $= -3.61$ – $3.72$ ,  $SD = 1.78$ ). The sample mean for RSES scores was  $20.15$  (range  $7$ – $29$ ,  $SD = 4.91$ ) and was used to center participants' RSES scores prior to multiple regression analyses (Aiken & West, 1991).

Previously, threat condition did not predict differences in initial-letter biases on the NLT or differences in biases toward same-initial writers' captions for Cartoon 1 or Cartoon 2 (individually or when averaged together). In the current analysis, threat condition also did not predict nickname initial biases on the NLT,  $R^2 = .054$ ,  $F(2, 49) = 1.40$ ,  $p = .26$ . Centered RSES scores also did not significantly predict biases toward nickname initials when this variable was added to the model,  $R^2_{change} = .000$ ,  $F(1, 48) = .01$ ,  $p = .94$ . The overall model including threat conditions, centered RSES scores, and their interactions was nonsignificant,  $R^2 = .148$ ,  $F(5, 46) = 1.60$ ,  $p = .18$ . The interaction terms did not together significantly improve the multiple regression model, although they approached significance,  $R^2_{change} = .094$ ,  $F(2, 46) = 2.53$ ,  $p = .09$ . In sum, when biases demonstrated toward nickname initial letters on the NLT were regressed on threat condition, biases did not vary as a function of participants' explicit self-esteem. For the progression of multiple regression equations, please refer to Table 10.

Table 10. Progression of multiple regression equations of NLT nickname initial biases on threat condition as a function of explicit self-esteem

a. Test of dummy variables only

$$\hat{Y} = b_1D_1 + b_2D_2 + b_0$$

$$\hat{Y} = (-.997)(D_1) + (-.753)(D_2) + .836$$

$$\text{Joint test of } b_1, b_2: R^2 = .054, F(2, 49) = 1.40, p = .26$$

$$\text{Test of } b_1: t(49) = -1.60, p = .12$$

$$\text{Test of } b_2: t(49) = -1.27, p = .21$$

$$SS_{\text{reg}} = 8.745; SS_{\text{res}} = 152.724$$

b. Test of dummy variables and continuous variable

$$\hat{Y} = b_1D_1 + b_2D_2 + b_3\text{RSES} + b_0$$

$$\hat{Y} = (-1.002)(D_1) + (-.754)(D_2) + (-.004)\text{RSES} + .838$$

$$\text{Joint test of } b_1, b_2, b_3: R^2 = .054, F(3, 48) = .92, p = .44$$

$$\text{Test of } b_3: R^2_{\text{change}} = .000, F(1, 48) = .01, p = .94$$

$$SS_{\text{reg}} = 8.765; SS_{\text{res}} = 152.705$$

c. Test of dummy variables, continuous variable, and interaction

$$\hat{Y} = b_1D_1 + b_2D_2 + b_3\text{RSES} + b_4(D_1 \times \text{RSES}) + b_5(D_2 \times \text{RSES}) + b_0$$

$$\hat{Y} = (-.836)D_1 + (-.733)D_2 + (-.006)\text{RSES} + (.217)(D_1 \times \text{RSES}) + (-.083)(D_2 \times \text{RSES}) + .838$$

$$\text{Joint test of } b_1-b_5: R^2 = .148, F(5, 46) = 1.60, p = .18$$

$$\text{Joint test of } b_4, b_5: R^2_{\text{change}} = .094, F(2, 46) = 2.53, p = .09$$

$$\text{Test of } b_3: t(46) = -.07, p = .95$$

$$\text{Test of } b_4: t(46) = 1.52, p = .14$$

$$\text{Test of } b_5: t(46) = -.74, p = .47$$

$$SS_{\text{reg}} = 23.895; SS_{\text{res}} = 137.575$$

Based on methods by Aiken & West (1991); dummy variables = self-concept threat conditions, continuous variable = centered RSES score, and the interaction of the two.  $N = 52$ , observed statistical power for final model = .81.

## Cartoon 1 Initial-Letter Biases (Nicknames)

### Differences in Biases Across Threat Conditions

Respondents who wrote at least three sentences relevant to the experimental threat condition to which they were randomly assigned, those with complete caption ratings on Cartoon 1, and those who indicated an alternate initial for the name they “went by” that

was different from that of their first given name were included in these first analyses ( $N = 45$ ). The average nickname initial-letter bias demonstrated by these participants was  $-.28$  (range =  $-3.78$ – $3.80$ ,  $SD = 1.88$ ).

Participants in the self-concept threat condition ( $N = 15$ ) demonstrated an average bias of  $+.19$  (range =  $-3.13$ – $3.80$ ,  $SD = 1.94$ ). Those assigned to the self-concept affirmation condition ( $N = 16$ ) had an average bias of  $-.52$  (range =  $-3.28$ – $3.21$ ,  $SD = 1.87$ ), while participants assigned to the control condition ( $N = 14$ ) exhibited an average bias of  $-.50$  (range =  $-3.78$ – $3.79$ ,  $SD = 1.86$ ). A one-way ANOVA was used to test for differences in biases between the three threat conditions and was nonsignificant,  $F(2, 42) = .69$ ,  $p = .51$ ,  $\eta^2 = .03160779$ .

#### Interaction of Threat Condition $\times$ Explicit Self-Esteem

This analysis was based on data from respondents who: a) indicated they “went by” a name other than their first given name such as a “nickname,” b) indicated their nickname initial, c) completed a threat condition writing task, d) completed all ratings for Cartoon 1, and e) had complete data on the RSES ( $N = 44$ ). The average nickname initial bias for participants in this sample was  $-.37$  (range =  $-3.78$ – $3.80$ ,  $SD = 1.79$ ). The average score on the RSES was  $20.45$  (range  $8$ – $30$ ,  $SD = 4.60$ ) for this sample and was used to center explicit self-esteem scores prior to multiple regression analyses (Aiken & West, 1991).

Previously, threat condition did not predict differences in NLT initial-letter biases (for nickname initials or first given name initials) or differences in initial-letter biases on

Cartoon 1 or Cartoon 2 (individually or when averaged together). In the current analysis, threat condition also did not predict Cartoon 1 biases toward nickname initials,  $R^2 = .057$ ,  $F(2, 41) = 1.23$ ,  $p = .30$ . RSES scores also did not significantly predict biases toward nickname initials when this variable was added to the model,  $R^2_{change} = .001$ ,  $F(1, 40) = .03$ ,  $p = .86$ . The addition of the interaction terms also failed to improve the multiple regression model,  $R^2_{change} = .035$ ,  $F(2, 38) = .73$ ,  $p = .49$ , and the overall model including threat conditions, centered RSES scores, and their interactions was nonsignificant,  $R^2 = .092$ ,  $F(5, 38) = .77$ ,  $p = .58$ . For the progression of equations regressing Cartoon 1 nickname initial biases on threat condition as a function of explicit self-esteem, please refer to Table 11.

### **Cartoon 2 Initial-Letter Biases (Nicknames)**

#### **Differences in Biases Across Threat Conditions**

Respondents who wrote at least three sentences relevant to the experimental threat condition to which they were randomly assigned, those with complete caption ratings on Cartoon 2, and those who indicated an alternate initial for the name they “went by” that was different from that of their first given name were included in these first analyses ( $N = 43$ ). The average nickname initial-letter bias demonstrated by these participants was  $-.19$  (range =  $-3.58$ – $3.02$ ,  $SD = 1.24$ ).

Participants in the self-concept threat condition ( $N = 13$ ) demonstrated an average bias of  $+.03$  (range =  $-2.06$ – $1.93$ ,  $SD = 1.10$ ). Those assigned to the self-concept affirmation condition ( $N = 18$ ) had an average bias of  $-.29$  (range =  $-3.58$ – $2.01$ ,  $SD = 1.35$ ), while participants assigned to the control condition ( $N = 12$ ) also exhibited an average bias of  $-.29$  (range =  $-1.98$ – $3.02$ ,  $SD = 1.30$ ). A one-way ANOVA was used to

test for differences in initial-letter biases between the three threat conditions and was nonsignificant,  $F(2, 40) = .28, p = .76, \eta^2 = .013745$ .

Table 11. Progression of multiple regression equations of Cartoon 1 nickname initial biases on threat condition as a function of explicit self-esteem

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a.	Test of dummy variables only
	$\hat{Y} = b_1D_1 + b_2D_2 + b_0$
	$\hat{Y} = (1.020)(D_1) + (.312)(D_2) + -.832$
	Joint test of $b_1, b_2$ : $R^2 = .057, F(2, 41) = 1.23, p = .30$
	Test of $b_1$ : $t(41) = 1.51, p = .14$
	Test of $b_2$ : $t(41) = .47, p = .64$
	$SS_{reg} = 7.812$ ; $SS_{res} = 130.378$
b.	Test of dummy variables and continuous variable
	$\hat{Y} = b_1D_1 + b_2D_2 + b_3RSES + b_0$
	$\hat{Y} = (1.007)(D_1) + (.322)(D_2) + (-.011)RSES + -.831$
	Joint test of $b_1, b_2, b_3$ : $R^2 = .057, F(3, 40) = .81, p = .50$
	Test of $b_3$ : $R^2_{change} = .001, F(1, 40) = .03, p = .86$
	$SS_{reg} = 7.914$ ; $SS_{res} = 130.275$
c.	Test of dummy variables, continuous variable, and interaction
	$\hat{Y} = b_1D_1 + b_2D_2 + b_3RSES + b_4(D_1 \times RSES) + b_5(D_2 \times RSES) + b_0$
	$\hat{Y} = (1.004)D_1 + (.396)D_2 + (.110)RSES + (-.132)(D_1 \times RSES) + (-.181)(D_2 \times RSES) + -.842$
	Joint test of $b_1-b_5$ : $R^2 = .092, F(5, 38) = .77, p = .58$
	Joint test of $b_4, b_5$ : $R^2_{change} = .035, F(2, 38) = .73, p = .49$
	Test of $b_3$ : $t(38) = .91, p = .37$
	Test of $b_4$ : $t(38) = -.78, p = .44$
	Test of $b_5$ : $t(38) = -1.20, p = .24$
	$SS_{reg} = 12.711$ ; $SS_{res} = 125.479$

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Based on methods by Aiken & West (1991); dummy variables = self-concept threat conditions, continuous variable = centered RSES score, and the interaction of the two.  $N = 44$ , observed statistical power for final model = .87.

### Interaction of Threat Condition $\times$ Explicit Self-Esteem

For this set of analyses, initial-letter biases were based on respondents' preferences for their nickname initial. Participants were included in the analyses if a) they indicated they "went by" a name other than their first given name such as a "nickname," b) they reported their nickname initial, c) completed a threat condition writing task, d) rated all twenty captions for Cartoon 2, and e) had complete data on the RSES ( $N = 43$ ). The mean nickname initial bias was the same average reported for the above sample:  $-.19$  (range =  $-3.58$ – $3.02$ ,  $SD = 1.24$ ). The average score on the RSES was  $20.84$  for the present sample (range  $8$ – $30$ ,  $SD = 4.43$ ), which was used to center participants' explicit self-esteem scores prior to multiple regression analyses (Aiken & West, 1991).

Biases toward captions submitted by writers who shared participants' first nickname initial were regressed on threat condition as a function of explicit self-esteem. Thus far, threat condition failed to predict initial-letter biases in all analyses, and in the current analysis, threat condition also did not predict Cartoon 2 nickname initial biases,  $R^2 = .014$ ,  $F(2, 40) = .28$ ,  $p = .76$ . RSES scores also did not significantly predict biases when this variable was added to the model,  $R^2_{change} = .015$ ,  $F(1, 39) = .62$ ,  $p = .44$ . Addition of the interactions between threat conditions and RSES scores did not significantly improve the multiple regression model,  $R^2_{change} = .068$ ,  $F(2, 37) = 1.38$ ,  $p = .26$ , and the overall model including threat conditions, centered RSES scores, and their interactions was nonsignificant,  $R^2 = .097$ ,  $F(5, 37) = .79$ ,  $p = .56$ . For the progression of multiple regression equations regressing Cartoon 2 nickname initial biases on threat condition as a function of explicit self-esteem, please refer to Table 12.

Table 12. Progression of multiple regression equations of Cartoon 2 nickname initial biases on threat condition as a function of explicit self-esteem.

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a.	Test of dummy variables only
	$\hat{Y} = b_1D_1 + b_2D_2 + b_0$
	$\hat{Y} = (.313)(D_1) + (-.001)(D_2) + -.288$
	Joint test of $b_1, b_2$ : $R^2 = .014, F(2, 40) = .28, p = .76$
	Test of $b_1$ : $t(40) = .62, p = .54$
	Test of $b_2$ : $t(40) = .00, p > .99$
	$SS_{reg} = .891$ ; $SS_{res} = 63.931$
b.	Test of dummy variables and continuous variable
	$\hat{Y} = b_1D_1 + b_2D_2 + b_3RSES + b_0$
	$\hat{Y} = (.249)(D_1) + (.017)(D_2) + (-.036)RSES + -.276$
	Joint test of $b_1, b_2, b_3$ : $R^2 = .029, F(3, 39) = .39, p = .76$
	Test of $b_3$ : $R^2_{change} = .015, F(1, 39) = .62, p = .44$
	$SS_{reg} = 1.883$ ; $SS_{res} = 62.939$
c.	Test of dummy variables, continuous variable, and interaction
	$\hat{Y} = b_1D_1 + b_2D_2 + b_3RSES + b_4(D_1 \times RSES) + b_5(D_2 \times RSES) + b_0$
	$\hat{Y} = (.382)D_1 + (.103)D_2 + (.073)RSES + (-.042)(D_1 \times RSES) + (-.170)(D_2 \times RSES) + -.311$
	Joint test of $b_1, b_2, b_3$ : $R^2 = .097, F(5, 37) = .79, p = .56$
	Joint test of $b_4, b_5$ : $R^2_{change} = .068, F(2, 37) = 1.38, p = .26$
	Test of $b_3$ : $t(37) = .76, p = .45$
	Test of $b_4$ : $t(37) = -.30, p = .77$
	Test of $b_5$ : $t(37) = -1.50, p = .14$
	$SS_{reg} = 6.262$ ; $SS_{res} = 58.560$

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Based on methods by Aiken & West (1991); dummy variables = self-concept threat conditions, continuous variable = centered RSES score, and the interaction of the two.  $N = 43$ , observed statistical power for final model = .87.

## Average Cartoon 1 and 2 Initial-Letter Biases

### Differences in Biases Across Threat Conditions

Respondents who wrote at least three sentences relevant to the experimental threat condition to which they were randomly assigned, those with complete caption ratings on both Cartoon 1 *and* Cartoon 2, and those who indicated an alternate initial for the name



they “went by” that was different from that of their first given name were included in these first analyses ( $N = 39$ ). The average nickname initial bias demonstrated by these participants was  $-.32$  (range =  $-3.30$ – $1.98$ ,  $SD = 1.08$ ).

Participants in the self-concept threat condition ( $N = 13$ ) exhibited an average bias of  $-.05$  (range =  $-1.28$ – $1.68$ ,  $SD = 1.06$ ). Those assigned to the self-concept affirmation condition ( $N = 15$ ) had an average bias of  $-.52$  (range =  $-3.30$ – $1.98$ ,  $SD = 1.31$ ), while respondents assigned to the control condition ( $N = 11$ ) demonstrated an average bias of  $-.37$  (range =  $-1.39$ – $1.20$ ,  $SD = .73$ ). A one-way ANOVA was used to test for differences in biases between the three threat conditions and was nonsignificant,  $F(2, 36) = .67$ ,  $p = .52$ ,  $\eta^2 = .03572313$ .

#### Interaction of Threat Condition $\times$ Explicit Self-Esteem

This analysis was based on average Cartoon 1 and Cartoon 2 biases exhibited toward nickname initials. Data were included from respondents who a) indicated they “went by” a name other than their first given name such as a “nickname,” b) reported their nickname initial, c) completed a threat condition writing task, d) rated all captions for both Cartoon 1 and Cartoon 2, and e) had complete data on the RSES ( $N = 39$ ). The mean nickname initial bias demonstrated by participants in this sample was the same average reported above:  $-.32$  (range =  $-3.30$ – $1.98$ ,  $SD = 1.08$ ). The average score on the RSES was  $20.87$  (range  $8$ – $30$ ,  $SD = 4.43$ ) and was used to center participants’ scores prior to multiple regression analyses (Aiken & West, 1991).

Biases toward captions submitted by writers who shared respondents’ nickname initial were regressed on threat condition as a function of explicit self-esteem. Previously, threat condition did not predict differences in NLT initial-letter biases (based on

participants' initial for their first given name or their nickname initial) or differences in biases on Cartoon 1 or Cartoon 2 individually (for initials of first given names or nicknames). In the current analysis, threat condition also did not predict average Cartoon 1 and Cartoon 2 biases toward captions written by same nickname initial writers,  $R^2 = .036$ ,  $F(2, 36) = .67$ ,  $p = .52$ . RSES scores also did not significantly predict biases when this variable was added to the model,  $R^2_{change} = .008$ ,  $F(1, 35) = .30$ ,  $p = .59$ . The interactions of threat condition and RSES scores also failed to significantly improve the multiple regression model,  $R^2_{change} = .095$ ,  $F(2, 33) = 1.82$ ,  $p = .18$ . The overall model including threat conditions, centered RSES scores, and their interactions was nonsignificant,  $R^2 = .139$ ,  $F(5, 33) = 1.06$ ,  $p = .40$ . For the progression of equations regressing average Cartoon 1 and Cartoon 2 nickname initial biases on threat condition as a function of explicit self-esteem, please refer to Table 13.

In sum, despite previous studies which have found initial-letter biases to diverge among those with low and high explicit self-esteem after experiencing a self-concept threat (Jones et al., 2002), the current investigation did not find evidence of such an interaction on any of the initial-letter preference tasks. Name-letter preferences' function of self-esteem regulation was not supported, even for those participants with high explicit self-esteem.

Table 13. Progression of multiple regression equations of average Cartoon 1 and Cartoon 2 nickname initial biases on threat condition as a function of explicit self-esteem

a. Test of dummy variables only

$$\hat{Y} = b_1D_1 + b_2D_2 + b_0$$

$$\hat{Y} = (.319)(D_1) + (-.154)(D_2) + -.367$$

$$\text{Joint test of } b_1, b_2: R^2 = .036, F(2, 36) = .67, p = .52$$

$$\text{Test of } b_1: t(36) = .71, p = .48$$

$$\text{Test of } b_2: t(36) = -.36, p = .73$$

$$SS_{reg} = 1.587; SS_{res} = 42.838$$

b. Test of dummy variables and continuous variable

$$\hat{Y} = b_1D_1 + b_2D_2 + b_3RSES + b_0$$

$$\hat{Y} = (.270)(D_1) + (-.151)(D_2) + (-.023)RSES + -.352$$

$$\text{Joint test of } b_1, b_2, b_3: R^2 = .044, F(3, 35) = .53, p = .66$$

$$\text{Test of } b_3: R^2_{change} = .008, F(1, 35) = .30, p = .59$$

$$SS_{reg} = 1.945; SS_{res} = 42.480$$

c. Test of dummy variables, continuous variable, and interaction

$$\hat{Y} = b_1D_1 + b_2D_2 + b_3RSES + b_4(D_1 \times RSES) + b_5(D_2 \times RSES) + b_0$$

$$\hat{Y} = (.475)D_1 + (-.047)D_2 + (.054)RSES + (.027)(D_1 \times RSES) + (-.142)(D_2 \times RSES) + -.403$$

$$\text{Joint test of } b_1-b_5: R^2 = .139, F(5, 33) = 1.06, p = .40$$

$$\text{Joint test of } b_4, b_5: R^2_{change} = .095, F(2, 33) = 1.82, p = .18$$

$$\text{Test of } b_3: t(33) = .61, p = .55$$

$$\text{Test of } b_4: t(33) = .22, p = .83$$

$$\text{Test of } b_5: t(33) = -1.38, p = .18$$

$$SS_{reg} = 6.155; SS_{res} = 38.270$$

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Based on methods by Aiken & West (1991); dummy variables = self-concept threat conditions, continuous variable = centered RSES score, and the interaction of the two.  $N = 39$ , observed statistical power for final model = .85.

## CHAPTER 7

### SELF-ATTITUDE ACCESSIBILITY AND INITIAL-LETTER BIASES

Participants were randomly assigned to surveys containing one of two administration orders of the Rosenberg Self-Esteem Scale (RSES; Rosenberg, 1989): before vs. after the initial-letter preference tasks. The purpose of this scale was to examine respondents' conscious self-worth ("explicit self-esteem") and its relation to initial-letter biases, as discussed in the previous chapter. Administration *order* of this explicit measure was manipulated to compare differences in biases between participants whose self-attitudes were accessible with biases of those who completed the RSES after the implicit tasks, corresponding to the "high" and "low" self-attitude accessibility conditions, respectively. Based on previous research, respondents whose self-attitudes had been primed prior to the Name Letter Test (NLT) and cartoon caption rating task were expected to exhibit stronger initial-letter biases than their low self-attitude accessibility counterparts (Krizan & Suls, 2008).

#### **The Name Letter Test**

Of the 420 participants with complete and non-redundant data on the NLT, 410 of these respondents also had complete data on the RSES. The average NLT initial-letter bias demonstrated by this group of participants was +1.36 (range = -3.43–6.05,  $SD = 1.62$ ). Participants assigned to the "high self-attitude accessibility" condition ( $N = 214$ ) on average exhibited a +1.43 initial-letter bias (range = -2.83–5.64,  $SD = 1.62$ ), while those assigned to the "low self-attitude accessibility" condition ( $N = 196$ ) on average

demonstrated a +1.29 bias toward their first given name initial (range = -3.43–6.05,  $SD = 1.63$ ). While in the predicted direction, high self-attitude accessibility participants did not exhibit significantly stronger initial-letter biases on the NLT than their low self-attitude accessibility counterparts,  $t(408) = .89, p = .38$  (two-tailed),  $d = .09$ .

### **Cartoon 1 Initial-Letter Biases**

Of the 428 respondents with complete and non-redundant ratings for Cartoon 1 captions, 419 of these participants also had complete data on the RSES. The average initial-letter bias exhibited by this group of participants was +.20 (range = -3.46–5.09,  $SD = 1.58$ ). High self-attitude accessibility respondents ( $N = 220$ ) on average demonstrated a +.17 bias on Cartoon 1 (range = -3.46–4.30,  $SD = 1.61$ ), while low self-attitude accessibility participants ( $N = 199$ ) on average exhibited a +.24 bias toward the caption written by the writer who shared their first given name initial (range = -3.45–5.09,  $SD = 1.56$ ). Differences in initial-letter biases between participants with high and low self-attitude accessibility were not in the predicted direction and were not statistically significant,  $t(417) = -.47, p = .64$  (two-tailed),  $d = -.04$ .

### **Cartoon 2 Initial-Letter Biases**

Of the 429 respondents with complete and non-redundant ratings for Cartoon 2 captions, 421 of these participants also had complete data on the RSES. The average initial-letter bias demonstrated by this group of participants was +.10 (range = -3.81–4.47,  $SD = 1.47$ ). High self-attitude accessibility participants ( $N = 213$ ) on average exhibited a +.26 initial-letter bias (range = -3.14–4.47,  $SD = 1.47$ ), while low self-attitude accessibility respondents ( $N = 208$ ) demonstrated an average *negative* bias of -.05 toward the caption written by the writer who shared their first given name initial

(range =  $-3.81$ – $3.48$ ,  $SD = 1.46$ ). Differences in biases between participants with high and low self-attitude accessibility were in the predicted direction and were statistically significant,  $t(419) = 2.16$ ,  $p = .03$  (two-tailed),  $d = .21$ .

### **Average Cartoon 1 and 2 Initial-Letter Biases**

Of the 394 respondents with complete and non-redundant ratings for both Cartoon 1 and Cartoon 2 captions, 387 of these participants also had complete data on the RSES. The average bias exhibited by this group of participants was  $+.16$  (range =  $-2.46$ – $3.12$ ,  $SD = 1.08$ ). High self-attitude accessibility participants ( $N = 200$ ) on average demonstrated a  $+.23$  initial-letter bias across both sets of cartoon captions (range =  $-2.46$ – $3.07$ ,  $SD = 1.12$ ), while low self-attitude accessibility respondents ( $N = 187$ ) exhibited an average bias of  $+.08$  toward captions written by writers who shared their first given name initial (range =  $-2.19$ – $3.12$ ,  $SD = 1.03$ ). Differences in biases between participants with high and low self-attitude accessibility were in the direction predicted, but were not statistically significant,  $t(385) = 1.33$ ,  $p = .19$  (two-tailed),  $d = .14$ .

Overall, biases demonstrated on the initial-letter preference tasks did not significantly differ between respondents who completed the explicit measure prior to vs. after the implicit tasks, except on Cartoon 2. For average initial-letter biases exhibited by low and high self-attitude accessibility participants on all initial-letter preference tasks (including biases demonstrated toward nickname initials), please refer to Table 14.

Table 14. Self-attitude accessibility and initial-letter biases

Analysis	Low		High	
	<i>N</i>	$\bar{X}$	<i>N</i>	$\bar{X}$
NLT	196	1.29	214	1.43
Cartoon 1	199	.24	220	.17
Cartoon 2*	208	−.05	213	.26
Both Cartoons	187	.08	200	.23
NLT nickname	20	−.15	37	.57
Cartoon 1 nickname	14	−.41	36	−.49
Cartoon 2 nickname	15	−.48	34	.10
Both cartoons nickname	13	−.56	32	−.20

\* differences in biases between low and high self-attitude accessibility groups were significantly different ( $p < .05$ , two-tailed).

## CHAPTER 8

### GENDER DIFFERENCES AND GENDER BIASES

Based on previous research investigating gender differences in name-letter preferences (Kityama & Karasawa, 1997; Pelham, Mirenberg, & Jones, 2002), women were expected to demonstrate stronger initial-letter biases than men. Ostensibly, women have a stronger affinity toward their first names because these names will remain with them throughout their lifetime. This study was also the first to examine whether gender *biases* serve to strengthen the name letter effect. First, differences in initial-letter biases between men and women were examined.

#### **Gender Differences**

##### The Name Letter Test

As predicted, initial-letter biases on the Name Letter Test (NLT) were significantly stronger for women than for men,  $t(418) = 2.51, p = .01$  (two-tailed),  $d = .25$ . On average, women exhibited a +1.52 bias toward their initial letter (range = -2.83–6.05,  $SD = 1.63$ ;  $N = 270$ ), while men demonstrated a +1.11 bias toward their initial (range = -3.43–4.97,  $SD = 1.61$ ;  $N = 150$ ). Individually, women's and men's initial-letter biases on the NLT were both significantly different from zero,  $t(269) = 15.38, p < .001, d = 1.14$  (women), and  $t(149) = 8.44, p < .001, d = .69$  (men).

When biases toward nickname initials were examined, men unexpectedly demonstrated stronger biases than did women. Men's average biases ( $M = +.64$ ,



range =  $-3.61$ – $4.97$ ,  $SD = 1.61$ ;  $N = 25$ ) and women's average biases ( $M = +.14$ , range =  $-2.70$ – $3.72$ ,  $SD = 1.86$ ;  $N = 33$ ) differed by half of a scale point, however, this difference was not statistically significant,  $t(56) = -1.08$ ,  $p = .29$  (two-tailed),  $d = .29$ . Individually, men's nickname initial biases were marginally significantly different from zero (with 0 = no bias),  $t(24) = 2.00$ ,  $p < .06$  (two-tailed),  $d = .40$ ; but women's nickname initial biases were not,  $t(32) = .43$ ,  $p = .67$ , (two-tailed)  $d = .08$ . This finding is interesting because women, as expected, demonstrated stronger initial-letter biases on the NLT than men when the initial for their first given name was used in analyses, but men exhibited stronger biases than women when it came to evaluations of their nickname initial.

#### Cartoon 1 Initial-Letter Biases

Average biases exhibited toward captions written by same-initial writers were identical between women ( $M = +.18$ , range =  $-3.46$ – $5.09$ ,  $SD = 1.64$ ;  $N = 282$ ) and men ( $M = +.18$ , range =  $-.75$ – $3.94$ ,  $SD = 1.48$ ;  $N = 146$ ). Individually, women's initial-letter biases were marginally significant,  $t(281) = 1.86$ ,  $p = .06$  (two-tailed),  $d = .11$ , but men's biases were not,  $t(145) = 1.44$ ,  $p = .15$  (two-tailed),  $d = .12$ .

Men did not demonstrate a significant bias toward captions submitted by same-nickname initial writers for Cartoon 1 ( $M = +.002$ , range =  $-3.02$ – $3.80$ ,  $SD = 1.95$ ;  $N = 20$ ),  $t(19) = .01$ ,  $p = .996$  (two-tailed),  $d = .001$ . An unexpected finding, however, emerged for women. Not only did they not prefer captions written by writers who shared their nickname initial, they rated them as significantly *less humorous* than captions submitted by writers who did not share their nickname initial ( $M = -.63$ , range =  $-3.78$ – $3.21$ ,  $SD = 1.69$ ;  $N = 31$ ),  $t(30) = -2.09$ ,  $p < .05$  (two-tailed),  $d = -.37$ . Whereas men were

essentially unbiased toward captions written by same-nickname initial writers, women were unexpectedly biased *against* these writers' captions, and significantly so.

#### Cartoon 2 Initial-Letter Biases

This time, men demonstrated stronger biases ( $M = +.14$ , range =  $-3.81$ – $3.44$ ,  $SD = 1.47$ ;  $N = 143$ ) than did women ( $M = +.07$ , range =  $-3.25$ – $4.47$ ,  $SD = 1.49$ ;  $N = 286$ ) toward captions submitted by same-initial writers, however this gender difference was not statistically significant,  $t(427) = -.428$ ,  $p = .67$  (two-tailed),  $d = .05$ . Neither men's nor women's individual biases were significantly different from zero,  $t(142) = 1.14$ ,  $p = .26$ , two-tailed (men),  $d = .10$ , and  $t(285) = .85$ ,  $p = .40$ , two-tailed (women),  $d = .05$ .

In nickname initial analyses, neither women nor men demonstrated biases toward captions submitted by writers who shared their nickname initial. Women, on average, exhibited a  $-.10$  bias (range =  $-2.06$ – $3.02$ ,  $SD = 1.33$ ;  $N = 32$ ), rating captions by same-initial writers slightly *lower* than captions written by different-initial writers. Men, on average, demonstrated a slightly less negative bias of  $-.04$  (range =  $-3.58$ – $2.01$ ,  $SD = 1.26$ ;  $N = 17$ ). Women's and men's difference in biases was not significant,  $t(47) = -.16$ ,  $p = .88$  (two-tailed),  $d = .05$ .

#### Average Cartoon 1 and 2 Initial-Letter Biases

When initial-letter biases for Cartoon 1 and Cartoon 2 were averaged together, women and men both exhibited a slight bias toward captions submitted by same-initial writers. On average, women rated both captions written by same-initial writers  $+.14$  points higher (range =  $-2.46$ – $3.07$ ,  $SD = 1.08$ ;  $N = 263$ ) relative to their ratings of non-initial writers' captions and as compared to normative ipsatized baseline ratings. This bias toward initial-letter writers' captions was significantly different from zero,  $t(262) =$

2.05,  $p = .04$ ,  $d = .13$ . Men rated captions written by same-initial writers on average +.16 points higher (range =  $-2.01$ – $3.12$ ,  $SD = 1.09$ ;  $N = 131$ ) than non-initial writers' captions and as compared to normative baselines. Men's average initial-letter biases, however, only approached significance,  $t(130) = 1.72$ ,  $p = .09$ ,  $d = .15$ . Taken together, both women and men demonstrated small biases toward both captions written by same-initial writers, and women's and men's biases did not significantly differ from each other,  $t(392) = -.23$ ,  $p = .82$  (two-tailed),  $d = .02$ .

Finally, in nickname analyses, neither women nor men exhibited biases toward their nickname initial letter. On the contrary, women on average exhibited a  $-.29$  *negative* bias toward captions written by writers who shared their nickname initial (range =  $-2.52$ – $1.98$ ,  $SD = .98$ ;  $N = 29$ ), while men exhibited a slightly stronger negative bias of  $-.34$  toward nickname-initial writers' captions (range =  $-3.30$ – $1.68$ ,  $SD = 1.16$ ;  $N = 16$ ). While both women and men unexpectedly exhibited negative nickname initial biases, neither of these biases were individually significantly different from zero,  $t(28) = -1.57$ ,  $p = .13$  (two-tailed),  $d = -.30$  (women), and  $t(15) = -1.17$ ,  $p = .26$ , (two-tailed),  $d = -.29$  (men). Gender differences in biases were also not statistically significant,  $t(43) = .27$ ,  $p = .87$  (two-tailed),  $d = .05$ .

### Same-Gender Biases

Next, biases toward captions submitted by writers who were the same gender as participants were measured by computing a "same-gender preference score," or "same-gender bias." The self-corrected algorithm (S-algorithm; LeBel & Gawronski, 2009) was used instead of the ipsatized algorithm because computation of normative baseline ratings for captions would not be a neutral comparison metric because these baselines would be

comprised of caption ratings made by participants of the opposite gender. Instead, the average rating men assigned to captions submitted by women writers (“opposite gender”) was subtracted from the average rating men assigned to captions submitted by men writers (“same gender”) to obtain their “same-gender bias” at the individual level. This procedure was repeated for women writers, with the average rating assigned to captions written by men subtracted from the average rating women assigned to captions written by other women.

#### Cartoon 1

The average same-gender bias for men respondents with complete caption ratings on Cartoon 1 was  $+.22$  (range =  $-1.20$ – $2.20$ ,  $SD = .65$ ;  $N = 146$ ). Men’s self-corrected increases in liking for captions written by men relative to ratings of captions written by women were significantly different from zero,  $t(145) = 4.14$ ,  $p < .001$ ,  $d = .34$ . Examination of women participants’ same-gender biases presented a different, opposite picture. Average biases ranged between  $-2.40$  and  $1.80$ , with a Mean of  $-.28$  ( $SD = .66$ ;  $N = 282$ ), indicating a bias toward captions submitted by *men* writers. While unexpected and not in the predicted direction, women respondents’ bias toward captions submitted by men was highly significant,  $t(281) = -7.106$ ,  $p < .001$ ,  $d = -.42$ . The difference in same-gender biases exhibited between men and women participants ( $N = 428$ ) was also highly significant,  $t(426) = 7.50$ ,  $p < .001$ ,  $d = .76$ .

#### Cartoon 2

These analyses were repeated for Cartoon 2 caption ratings and the opposite pattern of results emerged. This time, men respondents who rated all captions for Cartoon 2 demonstrated a bias toward captions written by *women* ( $M = -.18$ , range =  $-2.20$ – $1.00$ ,

$SD = .53$ ;  $N = 143$ ). Although not in the predicted direction, this bias was significant,  $t(142) = -4.07$ ,  $p < .001$ ,  $d = -.34$ . Women participants were biased toward captions written by other women for Cartoon 2 ( $M = +.27$ , range =  $-1.70$ – $2.60$ ,  $SD = .63$ ;  $N = 286$ ) and this same-gender bias was significant,  $t(285) = 7.33$ ,  $p < .001$ ,  $d = .43$ . The difference in same-gender biases exhibited between men and women participants on Cartoon 2 was significant,  $t(427) = -7.41$ ,  $p < .001$ ,  $d = .77$ .

#### Average of Cartoon 1 and 2

When same-gender biases for Cartoon 1 and 2 were averaged together, men were unbiased toward captions submitted by men ( $M = +.02$ , range =  $-1.15$ – $1.35$ ,  $SD = .41$ ;  $N = 131$ ), which was nonsignificant,  $t(130) = .50$ ,  $p = .62$ ,  $d = .05$ . Women were similarly unbiased toward women-written captions for Cartoon 1 and Cartoon 2 when biases were averaged together ( $M = +.004$ , range =  $-1.60$ – $1.80$ ,  $SD = .46$ ;  $N = 263$ ),  $t(262) = .15$ ,  $p = .88$ ,  $d = .01$ . The difference in same-gender biases exhibited between men and women participants ( $N = 394$ ) was also nonsignificant,  $t(392) = .29$ ,  $p = .77$ ,  $d = .04$ .

#### Shared Gender and Initial-Letter Biases

To determine whether initial-letter biases were even stronger when participants shared both the same first initial *and* the same gender as a caption writer, same-initial caption writer ratings were next coded as “same gender” vs. “different gender” for each respondent. For example, for a male participant with the initial letter “A,” his ipsatized and baseline-corrected humor rating assigned to the caption written by “Andrew” would be coded as “same gender” for Cartoon 1. His humor rating for the caption writer “Amanda” on Cartoon 2 would be coded as “different gender.” Equal numbers of each gender of caption writers (10 women’s names and 10 men’s names) for each of the 20

included letters of the alphabet were assigned to captions in random order for Caption

1. The opposite caption writer gender was assigned to that initial letter for Cartoon 2, thus, participants' gender matched that of their same-initial writer for either Cartoon 1 or Cartoon 2, but never both.

#### Cartoon 1

Using the above coding scheme for the 428 respondents with complete Cartoon 1 humor ratings, 205 participants' gender matched that of their same-initial caption writer and 223 participants' gender did *not* match that of their same-initial caption writer. On average, respondents' initial-letter biases were + .20 points higher when their initial-matching caption writer was also the same gender ( $M = +.29$ ,  $SD = 1.61$ ), compared to initial-letter biases exhibited by their initial-matching only counterparts ( $M = +.09$ ,  $SD = 1.56$ ). This difference in biases was in the direction predicted, however, it was not statistically significant,  $t(426) = 1.31$ ,  $p = .19$  (two-tailed),  $d = .13$ .

#### Cartoon 2

Out of 429 respondents with complete Cartoon 2 caption ratings, 217 participants' gender also matched that of their initial-matching caption writer's, while 212 participants' gender did *not* match that of their initial-matching caption writer. Mean initial-letter biases for initial- and gender-matching participants was +.13 ( $SD = 1.52$ ), while mean name-letter preferences for initial-only matching respondents was +.06 ( $SD = 1.45$ ). This difference, while again in the predicted direction, was not significant,  $t(427) = -.49$ ,  $p = .63$ ,  $d = .05$ .

### Cartoon 1 Nicknames

This analysis was based on 51 respondents who reported a nickname initial and had complete ratings for Cartoon 1. Participants whose gender also matched that of a nickname initial-matching caption writer ( $N = 31$ ) exhibited an average bias of  $-.27$  ( $SD = 1.74$ ), while those who only shared a nickname initial with a caption writer ( $N = 20$ ), demonstrated an average bias of  $-.56$  ( $SD = 1.94$ ). Both biases were unexpectedly negative, but gender-matching participants' nickname-initial biases were less negative than those for gender-mismatching participants. This difference, however, was not statistically significant,  $t(49) = .57, p = .58, d = .16$ .

### Cartoon 2 Nicknames

Of the 49 respondents who provided their nickname initial, 21 shared the same gender as their nickname initial-matching caption writer. Their average initial-letter bias was  $-.17$  ( $SD = 1.48$ ), compared to the average bias of  $-.01$  ( $SD = 1.16$ ) demonstrated by the 28 participants who only shared a nickname initial with a caption writer. This difference was not in the direction predicted and was also not statistically significant,  $t(47) = .437, p = .66, d = -.12$ .

## CHAPTER 9

### IMPLICIT SELF-ESTEEM AND INITIAL-LETTER BIASES

Implicit self-esteem was assessed using Gebauer, Riketta, Broemer, and Maio's (2008) single-item measure of name-liking to examine its association with initial-letter biases. Respondents rated how much they liked their full name from 1) not at all to 7) very much as an implicit measure of their unconscious global self-value and self-worth. The average name-liking rating was 5.59 (range = 1–7,  $SD = 1.30$ ;  $N = 479$ ). Men demonstrated slightly higher implicit self-esteem than did women, rating their name on average .13 points higher ( $M = 5.68$ , range = 2–7,  $SD = 1.25$ ;  $N = 166$ ) than women participants ( $M = 5.55$ ,  $SD = 1.33$ ;  $N = 313$ ). Men's and women's difference in implicit self-esteem scores, however, was not statistically significant,  $t(477) = 1.07$ ,  $p = .28$  (two-tailed),  $d = .10$ .

To examine the effect of implicit self-esteem on initial-letter biases, respondents' name-liking scores were first standardized. Cut-off values for "low implicit self-esteem" were equivalent to 1.00 or more standard deviations below the mean (set at zero), values greater than  $-1.00$  standard deviation, but less than  $+1.00$  standard deviation were categorized as "average implicit self-esteem," and values 1.00 standard deviation or more above the mean were considered to be suggestive of "high implicit self-esteem." Of the 479 participants who completed the single-item measure, 89 (18.6%) had



standardized scores indicative of “low implicit self-esteem,” 247 participants (51.6%) had “average implicit self-esteem,” and 143 had “high implicit self-esteem” (29.9%).

### **The Name Letter Test**

Scores on the Name Letter Test (NLT) were regressed on participants’ level of implicit self-esteem. Level of implicit self-esteem (“low,” “average,” and “high”) was dummy coded, with the “average implicit self-esteem” group serving as the comparison group. As predicted, level of implicit self-esteem predicted initial-letter biases on the NLT, but only accounted for 2% of the variance in biases,  $R^2 = .02$ ,  $F(2, 417) = 3.58$ ,  $p = .03$ . Also as predicted, the dummy-coded variable comparing the low implicit self-esteem group to the average implicit self-esteem group was negatively associated with initial-letter biases, however, it was not a significant predictor,  $t(417) = -.62$ ,  $p = .54$ . The dummy-coded variable comparing the high and average implicit self-esteem groups, however, was a significant and positive predictor of initial-letter biases,  $t(417) = 2.28$ ,  $p = .02$ . A follow-up regression analysis comparing only the low implicit self-esteem group to the high implicit self-esteem group was also significant,  $R^2 = .03$ ,  $F(1, 202) = 5.17$ ,  $p = .02$ .

When initial-letter biases exhibited on the NLT were regressed on level of implicit self-esteem (“low,” “average,” and “high”) for participants with nicknames ( $N = 58$ ), implicit self-esteem did not significantly predict biases,  $R^2 = .04$ ,  $F(2, 55) = 1.05$ ,  $p = .36$ .

### **Cartoon 1**

Cartoon 1 initial-letter biases were also regressed on participants’ level of implicit self-esteem (“low,” “average,” and “high”). The same dummy codes were used from the

above analysis, with the “average implicit self-esteem” group serving as the comparison group. Level of implicit self-esteem did not predict biases toward captions written by same-initial writers,  $R^2 = .002$ ,  $F(2, 425) = .46$ ,  $p = .63$ . The dummy-coded variable comparing the low implicit self-esteem group to the average implicit self-esteem group was again negatively associated with initial-letter biases, however, it was not a statistically significant predictor,  $t(425) = -.45$ ,  $p = .66$ . The dummy-coded variable comparing the high implicit self-esteem group to the average implicit self-esteem group was positively associated with initial-letter biases, but this predictor also was not statistically significant,  $t(425) = .67$ ,  $p = .51$ . A follow-up regression analysis comparing only the low implicit self-esteem group to the high implicit self-esteem group was non-significant as well,  $R^2 = .004$ ,  $F(1, 208) = .88$ ,  $p = .35$ .

When Cartoon 1 initial-letter biases were regressed on level of implicit self-esteem for respondents with nicknames ( $N = 51$ ), implicit self-esteem also did not significantly predict biases,  $R^2 = .01$ ,  $F(2, 48) = .18$ ,  $p = .83$ .

### **Cartoon 2**

Level of implicit self-esteem did not predict biases toward Cartoon 2 captions written by same-initial writers,  $R^2 = .01$ ,  $F(2, 426) = 1.16$ ,  $p = .31$ . The dummy-coded variable comparing the low implicit self-esteem group to the average implicit self-esteem group was again negatively associated with biases, however, it was not a significant individual predictor,  $t(426) = -1.42$ ,  $p = .16$ . The dummy-coded variable comparing the high implicit self-esteem group to the average implicit self-esteem group also was not statistically significant,  $t(426) = .08$ ,  $p = .93$ . A follow-up regression analysis comparing

only the low implicit self-esteem group to the high implicit self-esteem group did not reach statistical significance,  $R^2 = .01$ ,  $F(1, 204) = 1.96$ ,  $p = .16$ .

When Cartoon 2 initial-letter biases were regressed on level of implicit self-esteem (“low,” “average,” and “high”) for respondents with nicknames ( $N = 49$ ), implicit self-esteem did not significantly predict biases,  $R^2 = .04$ ,  $F(2, 46) = .96$ ,  $p = .39$ .

### **Cartoon 1 and Cartoon 2**

Level of implicit self-esteem did not predict biases toward captions written by same-initial writers when initial-letter biases were averaged together for both cartoons,  $R^2 = .01$ ,  $F(2, 391) = 1.45$ ,  $p = .24$ . The dummy-coded variable comparing the low implicit self-esteem group to the average implicit self-esteem group was again negatively associated with biases, however, it did not reach statistical significance as an individual predictor,  $t(391) = -1.43$ ,  $p = .16$ . The dummy-coded variable comparing the high implicit self-esteem group to the average implicit self-esteem group was positively associated with biases, but this predictor was again nonsignificant,  $t(391) = .45$ ,  $p = .66$ . A follow-up regression analysis comparing only the low implicit self-esteem group to the high implicit self-esteem group did not reach statistical significance,  $R^2 = .01$ ,  $F(1, 188) = 2.73$ ,  $p = .10$ .

When average Cartoon 1 and Cartoon 2 initial-letter biases were regressed on level of implicit self-esteem for participants with nicknames ( $N = 45$ ), implicit self-esteem was a marginally significant predictor of nickname-initial biases,  $R^2 = .13$ ,  $F(2, 42) = 3.01$ ,  $p = .06$ . This time, the dummy-coded variable comparing the low implicit self-esteem group to the average implicit self-esteem group was *positive* and significant,  $t(42) = 2.41$ ,  $p = .02$ . The variable comparing the high implicit self-esteem group to the

average implicit self-esteem group was also positive, but it was not significant,  $t(42) = .41, p = .69$ . Differences in initial-letter biases among participants with low, average, and high implicit self-esteem are summarized in Table 15.

Table 15. Level of implicit self-esteem and initial-letter biases

Analysis	Low		Average		High	
	<i>N</i>	$\bar{X}$	<i>N</i>	$\bar{X}$	<i>N</i>	$\bar{X}$
NLT*	76	1.14	216	1.27	128	1.69
Cartoon 1	81	.07	218	.16	129	.28
Cartoon 2	78	-.13	223	.14	128	.16
Both Cartoons	72	-.04	204	.17	118	.22
NLT nickname	11	.63	25	.61	22	-.07
Cartoon 1 nickname	9	-.15	23	-.55	19	-.30
Cartoon 2 nickname	9	.44	23	-.27	17	-.09
Both cartoons nickname	8	.46	21	-.53	16	-.39

\* biases differed significantly among levels of implicit self-esteem ( $p < .05$ , two-tailed).

## CHAPTER 10

### DISCUSSION

The aims of the present investigation were to provide a novel test of the name letter effect and to introduce two new variables that might influence initial-letter biases. The current study opened “the black box” of the subjective experience of humor and found preliminary empirical evidence that judgments of humor can fall prey to an automatic, unconscious and objective bias in predictable ways: via implicit egotism. Instead of being just another demonstration of the name letter effect, the relationship between implicit self-esteem and name-letter preferences was also examined in order to shed light on the current debate as to whether the Name Letter Test (NLT) measures one’s underlying implicit global sense of self-worth or is best understood as a measure of implicit egotism: the tendency to gravitate toward objects that share our self-attributes such as our name letters.

Past theory on implicit egotism and research on the name letter effect suggests that these biases are based on people’s implicit self-attitudes, which are overwhelmingly positive in content. However, recent researchers have suggested that perhaps there is another, overlooked side to the implicit egotism coin: name-letter biases might very well be reversed for those who possess truly negative unconscious self-attitudes, i.e., “low implicit self-esteem” (Pelham, Carvallo, & Jones, 2005). These authors recently raised important questions for future research on the name-letter effect: are name-letter

preferences due to participants' good fortune of possessing favorable implicit self-attitudes, or would biases be smaller or even reversed for those with low levels of implicit self-esteem? Pelham et al. (2005) emphasized that answers to these questions would be important to both social as well as clinical psychology. Thus, measurement of participants' level of implicit self-esteem and its influence on initial-letter biases was a critical aim of the current study. If name-letter preferences *are* indeed representative of one's underlying self-attitudes, then participants with low implicit self-esteem should not exhibit biases toward their name letters—perhaps even demonstrating a “reverse” name letter effect. However, if name letter effects occurred regardless of respondents' level of implicit self-worth, then name-letter preferences as a measure of implicit egotism—as it was originally designed—would be supported.

Biases toward persons of the same gender have been demonstrated in supervisor-supervisee relationships (Worthington & Stern, 1985) and therapist-patient treatment outcomes (Zlotnick, Elkin, & Shea, 1998; Sterling, Gottheil, Weinstein, & Serota, 1998), but had never before been examined within the field of implicit egotism. The current investigation aimed to determine whether shared gender with fictitious caption writers would further increase initial-letter biases, becoming the first implicit egotism experiment to examine both types of biases in tandem.

By examining these two potential new influences on initial-letter biases, as well as providing additional tests of previously identified variables' impact on these biases, the present study sought to increase the extant implicit egotism knowledge base by taking another look at name-letter preferences. Biases toward nickname initials were also measured, with results challenging the notion that virtually any self-related attribute can

ignite implicit egotism. Finally, while many researchers have and continue to administer the Name Letter Test (NLT) as a measure of implicit self-esteem, the current study sought to examine the overlap between it and a newer measure of implicit self-esteem: Gebauer, Riketta, Broemer, & Maio's (2008) single-item measure of name-liking.

### **General Initial-Letter Biases**

As predicted, name letter effects were found on both the NLT and in judgments of humor. Participants demonstrated a significant bias toward their first given name initial on the Name Letter Test (NLT), increasing their ratings by +1.38 points relative to non-initials and group-averaged normative baselines on the 1) not at all beautiful to 7) extremely beautiful scale. Also as predicted, respondents exhibited initial-letter biases in their humor ratings for cartoon captions submitted by same-initial writers. The average bias demonstrated across both sets of captions was +.15 points on the 1) not at all humorous to 7) extremely humorous scale. While only a small initial letter effect was found among ratings for captions submitted by same-initial writers, this finding suggests that the otherwise subjective experience of humor operates at least partially outside of our conscious awareness via one of our unconscious self-related biases—our implicit preference for our name letters.

The present study offered a rare opportunity to examine biases toward initials of names that participants indicated they “went by” and were different from the initial indicated for their first given name, i.e., “nicknames.” The purpose of this distinction in name-letter bias targets was two-fold. First, it allowed the researcher to confirm that respondents indeed “went by” the initial they indicated for their first given name in order

to establish the correct target name letter for analyses. Secondly, if they *did not* “go by” their first given name, participants were asked to indicate the first letter for another name such as a “nickname” if this was the name they routinely went by. In this way, the current investigation became the first implicit egotism experiment to examine nickname initial biases as distinct from first given name initial-letter biases.

As compared to nickname initials, the present study found that name-letter preferences are relevant to first given name initials only—regardless of whether one goes by their first given name or not. Nickname-initial biases on the NLT did not reach statistical significance and an unexpected pattern of results emerged when it came to biases demonstrated toward captions submitted by same-nickname initial writers. Participants actually preferred captions submitted by writers with whom they *did not* share a nickname initial, rating same-initial writers’ captions on average .30 points *lower* across both cartoons. Curiously, this “reverse nickname-letter effect” was significant. This is intriguing, especially when researchers such as Pelham, Mirenberg, and Jones (2002) have argued virtually “anything that people associate with the self” can prompt implicit egotism (p. 470). Moreover, authors such as Brendl, Chattopadhyay, Pelham, and Carvallo (2005) later eliminated participants from their name-letter analyses if they went by another name such as a nickname, because “the NL (name-letter) brand should have been constructed from the nickname” (p. 409). Why biases *against* nickname initials were found on the caption-rating tasks is counterintuitive and raises several empirical questions. As the first implicit egotism study to examine nickname initial biases specifically—and as distinct from first given name initial biases—a few possible explanations are offered.



First, perhaps name letter effects were too small to translate to nickname-initial biases. Only a small initial letter effect was found in the humor judgments exercise, thus it is possible that biases exhibited toward captions submitted by writers who shared participants' first given name initials were too small to translate to nickname initials. However, biases observed on the NLT for first given name initials were very large, yet they still did not translate into significant biases toward nickname initials on the NLT.

Nickname initial biases were decreased and sometimes even reversed. This begs the question as to whether name-letter and nickname-letter preferences assess general name-esteem, implicit egotism, or are a measure of implicit self-esteem. Perhaps participants held negative attitudes toward their nicknames and/or nickname initials and thus when confronted with nickname letter stimuli, their ratings reflected this negative attitude toward this particular self-attribute. However, it seems odd for a person to "go by" a name they do not particularly like. Implicit self-esteem for respondents who indicated they went by their first given name and for those who went by a nickname was nearly identical. However, this measure is based on participants' liking for their entire name (including their first and last name together) and it is unclear on which name respondents based their ratings, i.e., their first given name and last name or their nickname and last name. A name-liking measure which taps liking for participants' nicknames specifically might or might not provide a window into these respondents' implicit sense of self-worth, depending on how much of their identity is derived from their unofficial name. But in the end, implicit self-esteem—as measured by full name liking—only explained 1–2% of the variance in name-letter preferences anyway.

Another potential explanation for decreased nickname initial letter biases is that nicknames and nickname initials might not be as closely linked to one's identity as are first given names and their corresponding initials. Perhaps because a nickname is presumably bestowed upon a person later in life after one's identity is already established, it is thus less "self-defining" despite being a name one regularly "goes by." Unfortunately, it is impossible to know at which point in a participant's life a nickname was assumed in the current study with respect to either age or stage of identity formation. Previous studies that have ruled out the primacy effect as integral to name-letter biases challenge this explanation of later-bestowed nicknames carrying less self-identity punch and thereby mitigating implicit egotism. Namely, Hoorens and Todorova (1988) found name-letter effects among Bulgarian students' first *and* second languages with unique alphabets. These authors argue that students do not typically begin their acquisition of a second language with the mastery of name letters, and even if they did, it would be "less thrilling" the second time around (Hoorens & Todorova, 1988). And representing perhaps the strongest argument against nickname initials as not being self-defining enough are Feys' (1991; 1995) studies which were able to induce mere ownership in a laboratory setting, including biases for newly-learned name-related symbols.

Finally, perhaps nicknames are more self-defining for men than women. While men may be slower to mature, women—especially during the college years, which was the demographic of the current study's sample—might be more eager to eschew a childhood nickname. While this explanation is speculative, men did exhibit a marginally significant bias toward their nickname initial on the NLT and were nearly equivocal with respect to the humor ratings they assigned to nickname-initial and non-nickname initial

writers' captions. Women, on the other hand, demonstrated statistically significant *negative* biases toward their nickname initial on Cartoon 1. While only the negative Cartoon 1 bias displayed by women reached significance, this otherwise consistent trend in differences between men's and women's nickname initial letter biases is difficult to ignore. Men's vs. women's self-attitudes toward their nicknames and how "self-defining" they are for both genders await future research.

Whatever the reason nickname initial biases in the current study departed from that of typical name letter effects, these results challenge previous researchers' arguments that virtually any self-attribute—provided one indeed associates this attribute with the self—can predispose one to gravitate toward the object, person, or place that shares that self-attribute. More research is needed on nickname letters' link to one's identity and whether these stimuli offer the same opportunity for people to self-enhance when confronted with these letters. Interestingly, when initial-letter biases for first given names were re-examined *without* participants who indicated they went by a different (i.e., nickname) initial, results were nearly identical. Based on this, it is quite possible that name-letter preferences are relevant to first given name initials only, *even when* a person "goes by" another name beginning with a different letter. This lends support to the third explanation, i.e., that nickname initials are not as closely linked to one's identity as are initials for first given names and thus do not provide the same opportunity to self-enhance when confronted with only somewhat self-related stimuli.

### **Self-Attitude Accessibility**

A recent meta-analysis by Krizan and Suls (2008) found self-attitude accessibility increased initial-letter biases when explicit self-esteem measures were administered prior

to an initial-letter preference task. While participants did exhibit somewhat stronger biases when they completed the Rosenberg Self-Esteem Scale (RSES; Rosenberg, 1989) prior to the NLT and caption-rating task in the current study, this effect was only statistically significant for Cartoon 2. Why increasing respondents' accessibility of their self-attitudes did not replicate previous researchers' findings across all implicit tasks—especially the NLT—is curious because the same explicit measure (RSES) and implicit measure (NLT) were used. It is possible that the RSES measure was not administered in close enough proximity to the name-letter preference tasks to demonstrate an effect on the NLT (which participants completed after the caption-rating task). However, based on the average amount of time it took participants to complete the experiment—namely, 16.5 minutes on average—this is unlikely. Moreover, a self-attitude accessibility effect was found for the second set of cartoon captions but not the first, which further discounts attenuation due to increased temporal proximity and a diminished accessibility of self-attitudes. In addition, Cartoon 2's overall name letter effect was nonsignificant and even smaller than the name letter effect observed for Cartoon 1 (which *was* significant), yet a self-attitude accessibility effect was still found for Cartoon 2 (but not Cartoon 1). In fact, differences in biases between participants with high and low self-attitude accessibility were in *the opposite direction predicted* for Cartoon 1. In sum, the current study's mixed results offer very little support—if any—for increased name-letter preferences when self-attitudes are more accessible.

### **Gender Differences**

As predicted, women demonstrated significantly stronger first given name initial-letter biases than men, but only on the NLT. And as discussed previously, men's

nickname biases were marginally significant on the NLT, while women's were not.

Unexpectedly, when it came to nickname initials, both men and women exhibited either no bias or a *negative bias* toward these letters on the caption-rating task, with women's negative nickname initial biases reaching significance for Cartoon 1. These lack of biases—and sometimes even “reversed” biases—that participants demonstrated toward nickname initials in their judgments of humor again raises questions as to whether implicit egotism is theoretically relevant to nickname-related stimuli and/or if nicknames are as strongly associated with the self for women as they are for men. At least for humor judgments in the current study, it appears as though they are not.

### **Same-Gender Biases**

A recent study using a similar variation of a cartoon caption contest found that both men *and* women rated men's captions as funnier than those created by women, and both sexes misattributed humorous captions to having been written by men in a recall test (Mickes, Walker, Parris, Mankoff, & Christenfeld, 2012). In the present study, men exhibited a significant bias toward captions written by other men for Cartoon 1, however, men significantly preferred captions written by *women* for Cartoon 2. The opposite pattern of results emerged for women: they significantly preferred captions written by other women for Cartoon 2, but significantly preferred captions written by men for Cartoon 1. It is possible that more humorous captions were assigned to these genders for these cartoons, however, the gender of caption writers was randomly assigned. Another possible explanation is that a man was featured in the first cartoon, and consequently both men and women participants unconsciously or consciously favored men writers' perspectives on humorous caption content, while both men *and* women were featured in

the second cartoon. This explanation is still tenuous, since a significant preference for women-written captions emerged among both men's and women's ratings for Cartoon 2.

While in the direction predicted, shared gender with caption writers did not significantly increase first given name initial-letter biases, but produced mixed results for nickname initial biases. Same gender somewhat mitigated the observed negative bias toward nickname initials observed for Cartoon 1, while increasing them for Cartoon 2. It is not surprising that the self-relevant trait of gender failed to compound a name-letter effect for nickname initials when results of the current study failed to support the notion that nicknames are as closely linked to an individual's identity as are first given names (regardless of whether a person "goes by" that name or not) and are thus not as likely to fall prey to implicit egotism. Although shared gender *did* somewhat "buffer" the negative biases demonstrated toward nickname initials on the humor task, whether name-letter and gender biases work stronger in tandem awaits additional research.

### **Implicit Self-Esteem**

When NLT initial-letter biases were regressed on implicit self-esteem, participant's level of implicit self-esteem significantly predicted first given name initial preferences, however, it only accounted for 2% of the variance in these biases. Level of implicit self-esteem did not significantly predict initial-letter biases for Cartoon 1 or Cartoon 2, accounting for only 1% of the variance in these name-letter preferences. With implicit self-esteem predicting only 1–2% of the variance in the biases demonstrated on both initial-letter preference tasks in the current study, it is questionable as to whether name-letter preferences are indeed representative of and/or dependent upon one's implicit self-attitudes. Because a rather large name letter effect was exhibited on the NLT, one

would expect more than one-third of the present investigation's participants to have the favorable implicit self-attitudes thought to drive these observed initial-letter biases. Moreover, increasing the accessibility for what some researchers have argued are the overwhelming positive self-attitudes responsible for name-letter preferences did not consistently or significantly increase respondents' initial-letter ratings, or their ratings for captions submitted by same-initial writers. While implicit self-esteem may play some role in name-letter biases—in the current study, its role was very small indeed—it cannot fully account for our predisposition to gravitate toward self-related objects.

### **Self-Concept Threats**

Casting further doubt on the role of self-esteem in name-letter preferences, a threat to the self-concept did not increase biases—not even for those with high explicit self-esteem. Previous researchers have found name-letter biases to diverge among those with high vs. low explicit self-esteem after experiencing a temporary threat to the self-concept. Again, tempering a threat to the self-concept is thought to be well-practiced among those with high explicit self-esteem, suggesting classical conditioning underpinnings of the name letter effect. To this end, researchers argue that name letters and other self-related symbols are “fundamentally rewarding” and “the rough psychological equivalent of meat powder to a hungry puppy” (Jones et al., 2004, p. 680). Despite using these researchers' method for delivering a self-concept threat and the same name-letter preference measure, a threat to the self-concept was *not* a strong enough motive for participants to self-enhance when confronted with their name letters, regardless of their level of explicit self-esteem. Results of the present investigation do not support name-letter preferences' functional purpose of a self-regulation bias.

## Conclusion

What does the present study contribute to our understanding of implicit egotism and name-letter preferences? It has demonstrated that implicit egotism is still a prevalent social psychological phenomenon that exists not only among letter attractiveness ratings on the NLT, but also in judgments of humor—albeit to a lesser degree. Additional support for gender differences in NLT initial-letter biases was found, while findings for the role self-attitude accessibility played across name-letter preference tasks were weak and inconsistent. Previous findings of name-letter biases diverging among participants with high and low explicit self-esteem in response to a self-concept threat were not replicated. The impact of implicit self-esteem on initial-letter biases was examined, with level of implicit self-esteem only accounting for less than a fraction of the variance in name-letter preferences. Shared gender, when coupled with a shared initial with a caption writer, increased biases toward these writers' captions, but not significantly so. But perhaps one of the most surprising and interesting findings was that significant name letter effects were not found for nickname initials on the NLT or on the humor judgments task.

Does the NLT simply measure implicit self-esteem? Buhrmester, Blanton, and Swann (2011) argue that it does not. While the NLT is often used by researchers as an index of implicit self-esteem, these authors suggest that using the NLT to assess implicit self-attitudes is not recommended and that it is best understood as a measure of *implicit egotism*. In their review, Buhrmester et al. (2011) cite studies where participants appeared to have conscious access to unconscious feelings of self-worth (Gailliot & Schmeichel, 2006) and instances where approximately half of respondents recognized the self-



referential nature of the NLT (Krizan, 2008). These and other examples have led Buhrmester et al. (2011) to believe that instruments researchers have been using as implicit measures of self-esteem (including the NLT) might not be immune to the self-presentational biases inherent in explicit measures and “might be contaminated with conscious content that is not of theoretic interest” (p. 371). Consequently, implicit instruments such as the NLT may not be tapping the unconscious processes it purports to measure. Finally, because implicit self-esteem measures—including the NLT—did not predict general well-being and depression nowhere near as well as explicit measures did in their meta-analysis, the authors are justifiably skeptical about the NLT’s ability to tap self-esteem.

Implicit egotism researchers explain the name letter effect as a consequence of people overwhelmingly possessing positive self-attitudes (Pelham, Carvallo, & Jones, 2005). Buhrmester et al. (2011) argue that such an explanation is questionable because not everyone can have high implicit self-esteem. Results of the current study showed less than one-third of participants had the high implicit self-esteem that—up until recently—has been widely accepted as the mechanism driving name-letter preferences, challenging the notion that people overwhelmingly possess favorable self-attitudes. To further undercut such a rose-colored view of a population with “uniformly positive” implicit self-attitudes, Buhrmester et al. (2011) argue that other researchers (Cassidy, 1988; Sroufe, 1989; Diener & Diener, 1995) have found that as much as one-third of children possess insecure parent attachments, which have predicted incidence of low self-esteem later in life. These insecure attachments styles are the very ones that authors such as

DeHart, Pelham, and Tennen (2006) have found to be associated with low implicit self-esteem in adulthood.

The improbability and lack of evidence in the current study that positive implicit self-esteem is the norm, coupled with previous researchers' correlations between implicit and explicit measures increasing under cognitive load (Koole, Dijksterhuis, & van Knippenberg, 2001) all suggest that the first available self-referential piece of information is used in making implicit global self-assessments. Depleting participants' cognitive resources compromises what Buhrmester et al. (2011) refer to as the "depth of self-insight" necessary for making such an assessment and thus "removing the self from self-relevant responding" (p. 376). These same researchers argue that the NLT in particular also precludes "breadth of self-insight" because self-esteem is extremely multifaceted, yet the measure examines only a single aspect of participants' self-regard—namely, for their initials—in hopes that this one facet will translate to a much larger global view of self-worth.

It is more probable that compromising cognitive resources triggers results in line with implicit egotism because participants rely instead on automatic positivity biases rather than an authentic implicit global view of self. Activation of a universal positivity bias might explain the prevailing belief that people overwhelmingly possess favorable self-attitudes, i.e., "positive implicit self-esteem." Robust name-letter preferences need not be dependent on/reflective of one's level of implicit self-esteem. As to what the NLT measures exactly, Buhrmester et al. (2011) again suggest that it is best understood as a measure for which it was first designed, implicit egotism—the tendency to display automatic positivity biases—instead of a measure of implicit self-attitudes.

What were the correlations between the NLT, implicit self-esteem, and explicit self-esteem? In the present study, level of implicit self-esteem significantly predicted initial-letter biases on the NLT, but only accounted for 2% of the variance in biases,  $R^2 = .02$ ,  $F(2, 417) = 3.58$ ,  $N = 420$ ,  $p = .03$ . And when implicit self-esteem was treated as a continuous variable, the correlation between NLT biases and implicit self-esteem as measured by name-liking ratings became nonsignificant,  $r = .09$ ,  $N = 420$ ,  $p = .06$ . NLT scores were unrelated to explicit self-esteem as measured by the Rosenberg Self-Esteem Scale (RSES; Rosenberg, 1989),  $r = .03$ ,  $N = 410$ ,  $p = .56$ , while implicit self-esteem as measured by name-liking (Gebauer et al., 2008) was highly and significantly correlated with explicit self-esteem ( $r = .26$ ,  $N = 468$ ,  $p < .001$ ). Results of the present study do not support name-letter preferences as dependent on or wholly reflective of one's self-esteem—either implicit or explicit.

A closer look at the semi-partial correlations between level of implicit self-esteem and initial-letter biases while controlling for the effect of NLT scores on biases presents somewhat discouraging information about the unique variance that Gebauer et al.'s (2008) implicit self-esteem measure accounted for. Recall that level of implicit self-esteem was significantly but weakly correlated with NLT initial-letter biases and significantly predicted biases on this name-letter preference task only. And while implicit self-esteem level was associated with biases on the remaining tasks in the predicted directions (biases were negatively associated with low implicit self-esteem and positively associated with high implicit self-esteem), level of implicit self-esteem did not reach significance as a predictor in any of the comparisons. Nevertheless, because implicit self-esteem level accounted for a significant (yet very modest) proportion of the variance in

NLT biases, the relationship between implicit self-esteem and initial-letter biases on the remaining name-letter preference tasks was examined while controlling for the effects of NLT scores.

While the zero-order correlations were not statistically significant to begin with, the semi-partial correlations decreased further once the effect of NLT biases were residualized from biases exhibited on the cartoon caption rating task. For example, zero-order correlations between low vs. high implicit self-esteem and cartoon caption biases ranged from  $r = .07-.12$  (all  $ps > .10$ ). But when the effect of NLT biases was residualized from caption biases, semi-partial correlations decreased to  $r = .03-.06$ , all of which were also nonsignificant (all  $ps > .46$ ). While the zero-order correlations were originally also nonsignificant, it does appear that at least some—even as much as half—of the variance in initial-letter biases exhibited on the cartoon caption rating tasks can be accounted for by NLT biases.

Implicit self-esteem was significantly correlated with explicit self-esteem as measured by name-liking and the Rosenberg Self-Esteem Scale (RSES), respectively,  $r = .26, p < .001$ . Explicit self-esteem was only weakly related to average biases exhibited on the caption-rating tasks,  $r = .09, p = .07$ , and was *unrelated* to NLT biases,  $r = .03, p = .56$ , even though previous research has found a weak, yet significant correlation of  $r = .12$  between the NLT and RSES (Krizan & Suls, 2008). Two diametrically opposed inferences can be made for low correlations: either the implicit measure did not tap the intended construct—thereby challenging the measure’s validity—or the two measures tap two conceptually different constructs—supporting the measure’s discriminant validity (LeBel & Gawronski, 2009).

If the NLT does tap some aspect of self-esteem, then when compared to well-established measures of explicit self-esteem such as the RSES, this study's results support the NLT's discriminant validity. However, such a conclusion is tenuous at best because lack of support for convergent validity does not discriminant validity make. Buhrmester et al. (2011) said it best when criticizing a crafty approach for enhancing another implicit measure's test-retest reliability estimates and the same notion applies here: such an assumption would be "similar to elevating one's estimate of a basketball player's shooting ability based on his poor dribbling skills" (p. 367).

On the other hand, high correlations—such as the one observed in the current study between the implicit self-esteem name-liking measure and the explicit index of self-esteem (RSES)—carry a different set of possible interpretations. Either the implicit measure of self-esteem really did tap the intended construct of unconscious global self-worth—demonstrating its convergent validity with the explicit self-esteem measure—or the implicit measure was contaminated with explicit processes (LeBel & Gawronski, 2009). However, because name-liking likely taps autobiographical information—at least arguably more so than unconscious name-letter preferences—this gives support for the implicit measure's convergent validity with the RSES. Moreover, just as name-letter biases did not significantly increase when the RSES was administered prior to the NLT, name-liking also was not stronger among participants whose explicit self-attitudes had been primed, i.e., made more accessible, prior to the implicit self-esteem (name-liking) measure as compared to those who completed the name-liking measure first,  $t(466) = -.55, p = .58$ . Therefore, the implicit self-esteem measure could not have been contaminated by readily available explicit self-attitudes.

## **Limitations of the Current Study**

### **Web-Based Research**

One limitation of the current investigation is the lack of control and consistency over the experimental setting. Because the study was designed to be completed online, participants were free to take part in the experiment from any campus or personal computer that had a connection to the internet. An inherent pitfall of web-based research is that respondents may have been distracted by any number of factors, which could have potentially compromised their attention to the experiment. Because a researcher could not be present during the administration of the experiment in this type of environment, it is unknown whether such distractions may have prevented participants from providing their undivided attention to the study. However, the implicit processes underlying name-letter preferences are not dependent upon large amounts of cognitive resources and occur very efficiently.

In a similar vein, the amount of time participants spent completing the experiment could not be monitored and as a result, a small percentage of respondents had completion times that exceeded the amount of credit they were compensated for their participation (one hour). Unlike participants who might have spent too little time completing the experiment, long completion times suggest that participants might have begun the experiment and returned to it at a later time and/or did not rely on their “gut” feelings when assigning relevant attractiveness or humor ratings to name-letter stimuli. Thus, a conservative approach to these issues was to perform parallel analyses both with and without data from respondents with unusually short or long completion times. Fortunately, these concerns were unsubstantiated. Results were extremely similar, if not

identical, when analyses both excluded or included data from participants with short or long completion times. Moreover, the key dependent variables in the present study required little more than a heuristic-based judgment based on a rapid “gut” feeling, thus even under conditions of high cognitive load, an effect would still likely have been found. Indeed, rather large effect sizes were found for initial-letter biases, at least on the NLT.

#### A Note on Randomization

Men and women participants were randomly assigned to complete 1 of 24 surveys. The surveys were identical in content (with the exception of three different threat conditions) and consisted of two orders of explicit-implicit measures and four orders of captions. Every effort was taken to make aspects of the experiment random, however, all features were not possible to randomize using the technology of the web-based research tool. For example, while some researchers administer the NLT using one fixed random display order of alphabet letters (LeBel & Campbell, 2009), others have used several different random alphabet orders (Stieger & LeBel, 2012; Koole, et al, 2001; Nuttin, 1987) or even individual random orders (Dijksterhuis, 2004).

Ideally, at least several random orders would have been used in the present study, however, the number of surveys required to allow randomization of even the most essential experiment features already resulted in 48 surveys, including 24 surveys for men and 24 surveys for women. Manipulating alphabet order even at a minimum level with two different display orders would have required nearly 100 unique surveys. Because component parts of individual questions could not be randomized in the current study without creating numerous additional versions of the survey, the researcher focused on randomizing only the most critical experiment features, i.e., assignment of participants

to one of three self-concept threat conditions, one of two explicit-implicit measure orders, and one of four cartoon caption orders.

Captions were displayed in one of two random orders for each cartoon—forward or backward—with the forward order displaying a random arrangement of captions and the backward order displaying these same captions, but in reverse order. Thus, participants were randomly assigned to surveys that featured one of four caption orders for Cartoon 1 and Cartoon 2, i.e., forward-forward, forward-backward, backward-backward, and backward-forward. Statistically, a one-way ANOVA revealed no significant differences among average Cartoon 1 and Cartoon 2 initial-letter biases between the four different caption orders,  $F(3, 390) = 1.76, p = .16$ . Methodologically, this caption arrangement marks an improvement over previous researchers' methods which utilized a single fixed random order of name-letter candybars for all participants (Brendl et al., 2005).

#### Is the I-Algorithm Too Conservative?

When scoring the NLT, most researchers use the baseline-corrected algorithm ("B-algorithm") to compute name-letter preferences (LeBel & Gawronski, 2009). This method involves simply subtracting normative baseline letter ratings from participants' initial-letter ratings. While widely used, the B-algorithm has several issues including the production of skewed distributions and large standard deviations, while failing to control for participants' individual response tendencies such as those due to differences in mood or affect (LeBel & Gawronski, 2009).

LeBel and Gawronski (2009) recommend the use of the ipsatized double-correction algorithm ("I-algorithm") when computing name-letter preference scores,



based on results obtained after re-analyzing 18 different sets of studies that used the NLT. To compute scores, letter ratings are first ipsatized by subtracting the mean rating assigned to non-initial letters from all letter ratings—including both non-initials and initials. Next, normative ipsatized baseline letter ratings (assigned by participants who did not have that initial) are subtracted from respondents' ipsatized initial-letter rating.

This algorithm offers both methodological advantages as well as statistical benefits. Methodologically, the I-algorithm double-corrects at the individual *and* group level. First, it controls for participants' individual response tendencies—such as those due to mood, affect, or acquiescence—by ipsatizing letter ratings as described above. Second, it controls for aesthetic variability in name-letter stimuli by subtracting normative ipsatized baseline ratings from respondents' ipsatized initial-letter rating. In the current study, the I-algorithm's methodological superiority was especially desirable when it came to computing initial-letter biases in the cartoon caption tasks because it a) controlled for individual response tendencies due to mood, affect, and acquiescence in the first step of the calculation, and b) controlled for baseline differences in captions' humor in the second step. And in LeBel & Gawronski's (2009) meta-analysis, the I-algorithm boasted the following statistical advantages: high reliability estimates, the lowest levels of skewness/kurtosis, zero outliers, while introducing very little error due to the algorithm's control over participants' individual response tendencies and baseline letter attractiveness.

Despite these benefits and its well-documented optimality, some might argue that the I-algorithm is an overly conservative method for computing initial-letter biases because it doubly corrects initial-letter ratings. Doing so might serve to attenuate a name

letter effect that might have been otherwise observed if a less conservative algorithm had been used, such as the B-algorithm or S-algorithm. To examine this possibility, overall initial-letter biases were re-analyzed using each of the four other scoring algorithms (the B-algorithm, S-algorithm, D-algorithm, and Z-algorithm).

Initial-letter biases on the NLT were highly significant using all five algorithms (all  $ps < .001$ ), with effect sizes ranging from  $d = .68$ – $1.01$ . Biases on Cartoon 1 were significant on all algorithms (all  $ps < .05$ ; all  $ds = .10$ – $.19$ ), except for the D-algorithm, which produced a negative but nonsignificant name letter effect. Interestingly, the D-algorithm was the only algorithm that found significant initial-letter biases on Cartoon 2, but these biases were also negative in valence. When biases on both Cartoon 1 and Cartoon 2 captions were averaged together, all algorithms except for the B-algorithm found significant name letter effects (all  $ps < .01$ ), with effect sizes ranging from  $d = -.16$  (again, the D-algorithm) to  $d = .17$  (S-algorithm).

The S-algorithm was the only algorithm to find a significant and positive name letter effect for nickname initials on the NLT ( $p < .001$ ,  $d = .37$ ), while only the D-algorithm's computation of negative initial-letter biases observed for the cartoon caption tasks consistently reached significance across all analyses (all  $ps < .05$ , all  $ds = -.32$  to  $-.43$ ). Negative nickname-initial biases computed with the I-algorithm were significant ( $p = .05$ ,  $d = -.29$ ) only when scores were averaged across both sets of captions. Please refer to Table 16 for average initial-letter biases as computed by each of the five scoring algorithms.

Based on these supplementary analyses, the I-algorithm does not appear to have been too conservative. Results using the I-algorithm were similar to those achieved with

Table 16. Initial-letter biases as calculated by the five different scoring algorithms

	NLT		Cartoon 1		Cartoon 2		Both Cartoons	
	$\bar{X}$	<i>SD</i>	$\bar{X}$	<i>SD</i>	$\bar{X}$	<i>SD</i>	$\bar{X}$	<i>SD</i>
a. First given name initial								
I-algorithm	1.38****	1.63	.18**	1.59	.10	1.48	.15***	1.08
B-algorithm	1.38****	1.66	.18**	1.78	.10	1.77	.12	1.42
S-algorithm	1.63****	1.61	.32****	1.67	.05	1.55	.20***	1.15
D-algorithm	.34****	.50	-.06	.68	-.11***	.68	-.09***	.55
Z-algorithm	.88****	.97	.11**	.93	.07	.96	.10***	.66
b. Nickname initial								
I-algorithm	.36	1.76	-.38	1.81	-.08	1.30	-.30*	1.03
B-algorithm	.34	1.97	-.39	1.86	-.16	1.68	-.32	1.43
S-algorithm	.66***	1.80	-.40	1.93	.02	1.29	-.28	1.00
D-algorithm	.02	.60	-.26**	.74	-.23***	.71	-.27***	.63
Z-algorithm	.24	.97	-.18	.96	-.09	.89	-.17	.65

\* biases were marginally significantly different from zero ( $p = .05$ , two-tailed).

\*\* biases were significantly different from zero ( $p < .05$ , two-tailed).

\*\*\* biases were significantly different from zero ( $p < .01$ , two-tailed).

\*\*\*\* biases were significantly different from zero ( $p < .001$ , two-tailed).

two of the less conservative scoring algorithms (the B-algorithm and S-algorithm). In fact, initial-letter biases computed using the I-algorithm were very highly correlated with both of these algorithms, as well as the two other algorithms (all  $r_s > .87$ , all  $p_s < .001$ ).

Of the five algorithms, the D-algorithm produced the weakest initial-letter biases, however this may be because it produced the largest number of extreme values on the low end of the distribution—as many as 12 potential NLT outliers and 7 outliers in the analysis of average initial-letter biases on both cartoons' captions. LeBel and Gawronski (2009) also found the D-algorithm fared the worst out of the five algorithms with respect to the production of outliers. In the current study, the I-algorithm produced one of the lowest numbers of potential outliers, second only to the B-algorithm—which does not correct for individual response tendencies. Based on past and current research, the I-algorithm seems to indeed be the optimal scoring method for computing name-letter biases in this and future studies on the name letter effect.

### **Future Research on Name-Letter Biases**

One of the most surprising and interesting findings of the present study was that significant name letter effects were not found for nickname initials on the NLT or on the humor judgments task and only a small, marginally significant nickname letter bias was found for men on the NLT. Based on the difference observed between men's and women's affinity for nickname initials, are nicknames more self-defining for men than they are for women? Or, are nicknames equally self-defining for men and women, but women possess negative affect toward their nicknames and nickname initials? The latter explanation seems tenable—at least when it came to assigning humor ratings to captions submitted by same-nickname initial writers. Women demonstrated a consistent trend of

negative biases toward same-nickname initial writers' captions, which reached significance for Cartoon 1. While only the Cartoon 1 nickname-initial biases exhibited by women reached significance, this otherwise consistent trend is difficult to ignore.

Based on the present study's unexpected, yet consistent findings for either negative or zero-biases demonstrated toward nickname letters, it behooves future researchers examining name-letter biases to make the distinction between first given names and nicknames as participants' target of bias. Subsuming both types of names under a single category is cautioned because it is unclear to what degree observed name-letter preferences thus far may have been attenuated by nicknames which might not be imbued with the same self-defining properties as one's first given name—even for people who “go by” a nickname. Researchers would be wise to query participants directly for their first given name—if different from the name they go by—to avoid making a Type II error. Moreover, doing so will provide additional opportunities to examine nickname letters as self-referential stimuli or whether confrontation with nickname letters triggers intervening cognitive processes that may serve to mitigate or even reverse the name letter effect, as was the case with the current study.

Likewise, and based on the present investigation's results, single-item measures of implicit self-esteem might be adapted to assess participants' nickname-liking specifically, if they go by such a name. It is unclear which name respondents used to make their name-liking ratings on this measure and, consequently, whether this measure accurately examined nicknamed participants' implicit self-esteem. If nicknames are not as self-referential as are first given names like the findings of this study suggest—or alternatively, if they are self-defining, but carry negative affect for some women—then

we might not have an accurate picture of participants' implicit self-esteem if they relied on liking of their nicknames while completing this measure.

More research is needed to differentiate the NLT as an implicit self-esteem instrument versus a measure of implicit egotism. Does it measure one's unconscious global self-worth, or is it best understood merely as an index of one's predisposition to rely on universal automatic self-positivity biases? While the tendency to rely on automatic positivity biases might be universal, implicit self-esteem cannot. Initial-letter biases were smaller—and in some instances even reversed—among those with low implicit self-esteem, however, this effect did not reach statistical significance. Level of implicit self-esteem as a whole only predicted a very small proportion (1–2%) of first given name initial-letter biases on both of the name-letter preference tasks, and NLT biases were uncorrelated with explicit self-esteem. Results of the current study suggest that name-letter preferences are not synonymous with implicit self-esteem and that favorable self-attitudes—whether implicit or explicit—are not a necessary condition of initial-letter biases. Just because people possess a universal dominant response tendency to prefer their name letters does not mean this preference is diagnostic of high implicit self-esteem. In the present study, less than one-third of participants possessed the favorable self-attitudes thought to drive name-letter preferences.

Jones et al. (2004) have suggested that important decisions—such as choices in mates, careers, and places of residence—are inherently threatening to the self-concept because these decisions carry costly consequences. As such, implicit egotism—as defined by the tendency to gravitate toward people, jobs, and cities/states that share our self-attributes—might serve the purpose of an unconscious “safety signal,” with presumably

the least threatening person an individual knows being him or herself. Thus, by association, self-related targets are deemed the “safe” and least-threatening choice among the larger array. Such a “threat-buffering function” view of implicit egotism says nothing about people’s favorable self-attitudes spilling over into evaluations of self-related objects. In fact, people who feel good about themselves *should not* be as threatened by important life decisions and should therefore not be as apt to rely on the unconscious “safety signal” broadcast by name letters, as compared to those who do not possess such favorable self-attitudes and self-confidence.

On the contrary, previous researchers have found temporary threats to the self-concept can serve as a motive for participants to self-enhance when confronted with name-letter stimuli, particularly among those with high explicit self-esteem. In the current study, however, evidence of implicit egotism’s function as a safety signal or similar self-regulating purpose following a threat to the self-concept was not found (even on the NLT which demonstrated a very strong name letter effect). This is unexpected and difficult to explain because measures of explicit self-esteem and the threat manipulation were identical to the ones used by previous researchers who observed such an interaction (Jones, Pelham, Mirenberg, & Hetts, 2002). Additional research is needed in order to further investigate whether preferring our name letters serves a self-regulating (or other) function and whether there are unconscious self-serving benefits to be had for exhibiting biases toward these self-related stimuli.

A trend for the additive effects of shared gender on initial-letter biases was found. Extensions and replications of these effects are needed to firmly establish this dual source of bias toward self-related attributes. Recently, the 281<sup>st</sup> real-life *New Yorker* cartoon

caption contest was won by film critic Roger Ebert, which might be less than surprising to those who allege Mr. Ebert had “a leg up” in the competition due to his celebrity. However, social psychologists—and implicit egotism researchers in particular—might argue he had two: his gender and his name letters he shared with *New Yorker* cartoon caption contest editor Robert Mankoff. Whether Robert’s self-relevant trait of gender and the name letters he shared with Roger—four letters in total—worked in tandem to land the famous film critic among the weekly contest’s finalists is unknown, but to scientific minds fascinated with dissecting humor, it’s great fun to speculate. After all, for participant judges in the current cartoon caption contest study, shared name letters with caption writers—and to some extent, shared gender—were indeed a laughing matter.



APPENDIX A  
ROSENBERG SELF-ESTEEM SCALE (RSES)

*Instructions:* Below is a list of statements dealing with your general feelings about yourself. If you strongly agree, choose **SA**. If you agree with the statement, choose **A**. If you disagree, choose **D**. If you strongly disagree, choose **SD**.

- |  |                  |
|--|------------------|
| 1. On the whole, I am satisfied with myself.                                   | <b>SA A D SD</b> |
| 2. At times, I think I am no good at all.                                      | <b>SA A D SD</b> |
| 3. I feel that I have a number of good qualities.                              | <b>SA A D SD</b> |
| 4. I am able to do things as well as most other people.                        | <b>SA A D SD</b> |
| 5. I feel I do not have much to be proud of.                                   | <b>SA A D SD</b> |
| 6. I certainly feel useless at times.  | <b>SA A D SD</b> |
| 7. I feel that I am a person of worth, at least on an equal plane with others. | <b>SA A D SD</b> |
| 8. I wish I could have more respect for myself.                                | <b>SA A D SD</b> |
| 9. All in all, I am inclined to feel that I am a failure.                      | <b>SA A D SD</b> |
| 10. I take a positive attitude toward myself.                                  | <b>SA A D SD</b> |

APPENDIX B

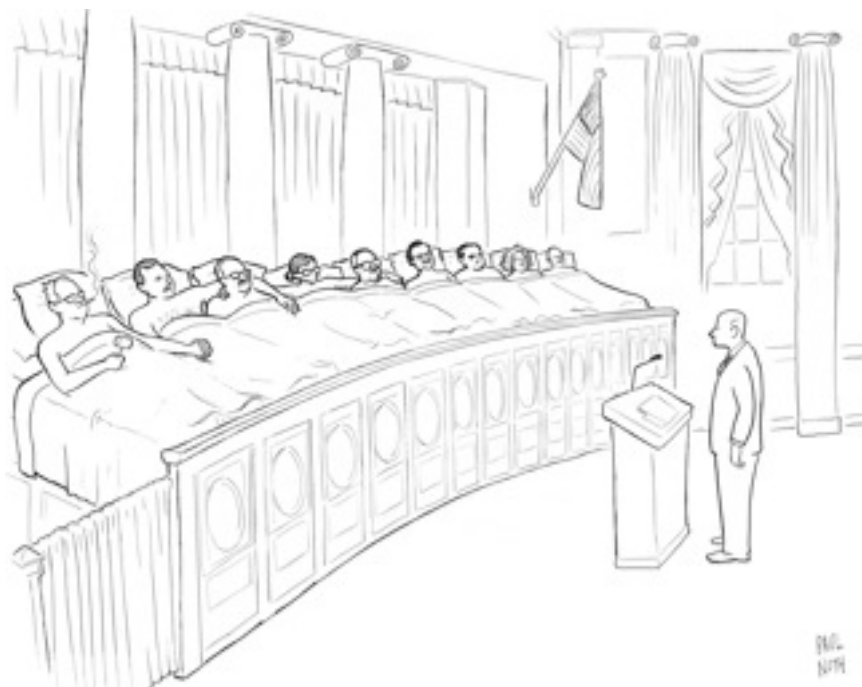
CARTOONS AND CAPTIONS

## Cartoon 1



submitted by	caption	(not at all humorous)					(extremely humorous)	
		1	2	3	4	5	6	7
<b>Kurt</b>	"No, honey, the cats are not on the bed."	○	○	○	○	○	○	○
<b>Vanessa</b>	"Hold on a minute, the cats got my tongue."	○	○	○	○	○	○	○
<b>Amanda</b>	"When TV went digital, I went feline."	○	○	○	○	○	○	○
<b>Ian</b>	"Let's stop buying the gourmet cat food."	○	○	○	○	○	○	○
<b>Jessica</b>	"No problems. They seem to be holding up okay."	○	○	○	○	○	○	○
<b>Frank</b>	"We won by a whisker."	○	○	○	○	○	○	○
<b>Thomas</b>	"My cats will do anything to get me off the phone."	○	○	○	○	○	○	○
<b>Greg</b>	"A few more and we can change that lightbulb."	○	○	○	○	○	○	○
<b>Helen</b>	"Sure it's cute, but I'm still allergic."	○	○	○	○	○	○	○
<b>Dennis</b>	"The gig's off. The beagles got the part."	○	○	○	○	○	○	○
<b>Pauline</b>	"Well, it's no dog and pony show."	○	○	○	○	○	○	○
<b>Bethany</b>	"It's hard to say. I'm more of a dog person."	○	○	○	○	○	○	○
<b>Regina</b>	"Yes, but can they balance a checkbook?"	○	○	○	○	○	○	○
<b>Nicole</b>	"Can't you just leave me a note like other wives?"	○	○	○	○	○	○	○
<b>Lauren</b>	"They're a class act, but they're union."	○	○	○	○	○	○	○
<b>Samantha</b>	"You should see how they respond to catnip."	○	○	○	○	○	○	○
<b>Eric</b>	"My cats describe the intruder as tall and thin."	○	○	○	○	○	○	○
<b>Wes</b>	"Well, what if they were wearing sweaters?"	○	○	○	○	○	○	○
<b>Christopher</b>	"You've gotta be kitten me."	○	○	○	○	○	○	○
<b>Matt</b>	"Actually, I don't miss cable."	○	○	○	○	○	○	○

## Cartoon 2



submitted by	caption	(not at all humorous)						(extremely humorous)	
		1	2	3	4	5	6	7	
<b>Scott</b>	"A show of hands, please, for those who want cookies and milk."	o	o	o	o	o	o	o	
<b>Gillian</b>	"Fankly, we're all tired of this argument."	o	o	o	o	o	o	o	
<b>Meghan</b>	"We'll take a two hour recess and reconvene after our nap."	o	o	o	o	o	o	o	
<b>Vince</b>	"Once upon a time, in a jurisdiction far, far away..."	o	o	o	o	o	o	o	
<b>Brian</b>	"Now that you're all comfortable with the facts in this case..."	o	o	o	o	o	o	o	
<b>Isabel</b>	"I don't want to make any blanket statements."	o	o	o	o	o	o	o	
<b>Andrew</b>	"The court has been shortsighted and short-sheeted."	o	o	o	o	o	o	o	
<b>Trisha</b>	"The facts of this case clearly present a wake-up call."	o	o	o	o	o	o	o	
<b>Danielle</b>	"I didn't know the Court could recline a case."	o	o	o	o	o	o	o	
<b>Peter</b>	"We've heard your argument... now we need to sleep on it."	o	o	o	o	o	o	o	
<b>Felicia</b>	"The court finds the defendant not quilty."	o	o	o	o	o	o	o	
<b>Emma</b>	"The decision is unanimous: we'd like to hear 'Goodnight Moon.'"	o	o	o	o	o	o	o	
<b>Logan</b>	"If the defense rests, then so will we."	o	o	o	o	o	o	o	
<b>Kristen</b>	"And in summation, sleep tight and don't let the bedbugs bite."	o	o	o	o	o	o	o	
<b>Ryan</b>	"In conclusion, I'd like to say, 'Nighty-night.'"	o	o	o	o	o	o	o	
<b>Chelsea</b>	"By unanimous decision, we have chosen sleep number five."	o	o	o	o	o	o	o	
<b>Jacob</b>	"If it pleases the court, I will now dim the lights."	o	o	o	o	o	o	o	
<b>Hank</b>	"And they lived happily ever after in eternal litigation. The end."	o	o	o	o	o	o	o	
<b>Whitney</b>	"I'm glad we've put this issue to bed."	o	o	o	o	o	o	o	
<b>Noah</b>	"And as evidence... the nine said mattresses missing their tags."	o	o	o	o	o	o	o	

APPENDIX C  
THE NAME LETTER TEST (NLT)

**Instructions:** This portion of the study is concerned with aesthetic judgments of lexical stimuli. While it might seem unusual to evaluate the letters of the alphabet in terms of their beauty, previous research has found such judgments to foster a better understanding of language and human emotions. Please estimate **how beautiful** you find each of the following letters using your “gut feelings.”

<b>Q</b>	(1)	(2)	(3)	(4)	(5)	(6)	(7)	<b>X</b>	(1)	(2)	(3)	(4)	(5)	(6)	(7)
<i>not at all</i>								<i>not at all</i>							
<i>beautiful</i>							<i>extremely</i>	<i>not at all</i>							<i>extremely</i>
							<i>beautiful</i>								<i>beautiful</i>
<b>D</b>	(1)	(2)	(3)	(4)	(5)	(6)	(7)	<b>J</b>	(1)	(2)	(3)	(4)	(5)	(6)	(7)
<i>not at all</i>								<i>not at all</i>							
<i>beautiful</i>							<i>extremely</i>	<i>not at all</i>							<i>extremely</i>
							<i>beautiful</i>								<i>beautiful</i>
<b>U</b>	(1)	(2)	(3)	(4)	(5)	(6)	(7)	<b>K</b>	(1)	(2)	(3)	(4)	(5)	(6)	(7)
<i>not at all</i>								<i>not at all</i>							
<i>beautiful</i>							<i>extremely</i>	<i>not at all</i>							<i>extremely</i>
							<i>beautiful</i>								<i>beautiful</i>
<b>Y</b>	(1)	(2)	(3)	(4)	(5)	(6)	(7)	<b>E</b>	(1)	(2)	(3)	(4)	(5)	(6)	(7)
<i>not at all</i>								<i>not at all</i>							
<i>beautiful</i>							<i>extremely</i>	<i>not at all</i>							<i>extremely</i>
							<i>beautiful</i>								<i>beautiful</i>
<b>I</b>	(1)	(2)	(3)	(4)	(5)	(6)	(7)	<b>S</b>	(1)	(2)	(3)	(4)	(5)	(6)	(7)
<i>not at all</i>								<i>not at all</i>							
<i>beautiful</i>							<i>extremely</i>	<i>not at all</i>							<i>extremely</i>
							<i>beautiful</i>								<i>beautiful</i>
<b>G</b>	(1)	(2)	(3)	(4)	(5)	(6)	(7)	<b>L</b>	(1)	(2)	(3)	(4)	(5)	(6)	(7)
<i>not at all</i>								<i>not at all</i>							
<i>beautiful</i>							<i>extremely</i>	<i>not at all</i>							<i>extremely</i>
							<i>beautiful</i>								<i>beautiful</i>
<b>R</b>	(1)	(2)	(3)	(4)	(5)	(6)	(7)	<b>F</b>	(1)	(2)	(3)	(4)	(5)	(6)	(7)
<i>not at all</i>								<i>not at all</i>							
<i>beautiful</i>							<i>extremely</i>	<i>not at all</i>							<i>extremely</i>
							<i>beautiful</i>								<i>beautiful</i>
<b>T</b>	(1)	(2)	(3)	(4)	(5)	(6)	(7)	<b>O</b>	(1)	(2)	(3)	(4)	(5)	(6)	(7)
<i>not at all</i>								<i>not at all</i>							
<i>beautiful</i>							<i>extremely</i>	<i>not at all</i>							<i>extremely</i>
							<i>beautiful</i>								<i>beautiful</i>
<b>P</b>	(1)	(2)	(3)	(4)	(5)	(6)	(7)	<b>Z</b>	(1)	(2)	(3)	(4)	(5)	(6)	(7)
<i>not at all</i>								<i>not at all</i>							
<i>beautiful</i>							<i>extremely</i>	<i>not at all</i>							<i>extremely</i>
							<i>beautiful</i>								<i>beautiful</i>
<b>M</b>	(1)	(2)	(3)	(4)	(5)	(6)	(7)	<b>V</b>	(1)	(2)	(3)	(4)	(5)	(6)	(7)
<i>not at all</i>								<i>not at all</i>							
<i>beautiful</i>							<i>extremely</i>	<i>not at all</i>							<i>extremely</i>
							<i>beautiful</i>								<i>beautiful</i>
<b>A</b>	(1)	(2)	(3)	(4)	(5)	(6)	(7)	<b>N</b>	(1)	(2)	(3)	(4)	(5)	(6)	(7)
<i>not at all</i>								<i>not at all</i>							
<i>beautiful</i>							<i>extremely</i>	<i>not at all</i>							<i>extremely</i>
							<i>beautiful</i>								<i>beautiful</i>
<b>B</b>	(1)	(2)	(3)	(4)	(5)	(6)	(7)	<b>W</b>	(1)	(2)	(3)	(4)	(5)	(6)	(7)
<i>not at all</i>								<i>not at all</i>							
<i>beautiful</i>							<i>extremely</i>	<i>not at all</i>							<i>extremely</i>
							<i>beautiful</i>								<i>beautiful</i>
<b>H</b>	(1)	(2)	(3)	(4)	(5)	(6)	(7)	<b>C</b>	(1)	(2)	(3)	(4)	(5)	(6)	(7)
<i>not at all</i>								<i>not at all</i>							
<i>beautiful</i>							<i>extremely</i>	<i>not at all</i>							<i>extremely</i>
							<i>beautiful</i>								<i>beautiful</i>

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## VITA

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